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PENSIONS AND WAGE PREMIA

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ABSTRACT

In this paper we use that the theory of compensating differentials to identify sources of heterogeneity in firms' costs of providing fringe benefits and hence heterogeneity in the magnitude of the compensating differential. We estimate the relationship between pensions and wages controlling for variations in the size of the compensating differential related to firm size or the presence of a union. Both firm size and unionism are commonly associated with the payment of wage premia and/or the presence of market power where the costs of fringe benefits to the firm may be less. Our results are consistent with these a priori expectations and suggest that the magnitude of the compensating differential is significantly higher in nonunion and in small firms.

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I. Introduction

Despite the widespread acceptance of the theory of compensating differences, empirical efforts to estimate the magnitude of these differences have met with only mixed success.[1] With the exception of fatal risks, economists have not been able to consistently find a significant relationship between onerous job characteristics and wages. These efforts to examine the relationship between wages and job characteristics have been hampered by difficulties in conceptualizing, let alone measuring precisely, the desired job characteristic. Efforts to estimate the extent of the compensating differential between wage and nonwage aspects of compensation (such as pensions and health benefits), which are conceptually easier to measure, have provided some evidence in favor of compensating differentials although the magnitude of the tradeoff is still subject to debate.[2]

One frequent explanation for this lack of consensus is that data problems have limited the ability of researchers to control for selectivity and unobserved ability differences. An equally important consideration is that in a world with heterogenous workers and firms, the theory of equalizing differences predicts that there may be a difference in compensating differentials across markets. If individual and firm-level differences exist in the value of costs of any given job attribute, and workers and

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[1] See Rosen (1986) for a review of the literature on the theory of compensating differences and empirical efforts to document their existence.

[2] See Montgomery, Shaw and Benedict (1992) and Smith and Ehrenberg (1983).

firms are not randomly assigned to markets, compensating differential theory is consistent with the finding of divergent estimates of the implicit hedonic locus. Thus, given the variations in tastes and technologies that are likely to exist in various markets in the economy, it is not surprising that researchers have gotten varying estimates of the extent of compensating differential.

In this paper we exploit the fact that the theory of compensating differentials can be used to identify sources of heterogeneity in firms' costs of providing fringe benefits and hence heterogeneity in the magnitude of the tradeoff. The few studies have attempted to test for heterogeneity in compensating differentials have concentrated on individual differences in tastes for risk.[3] In contrast, we estimate the relationship between pensions and wages taking into account the potential role of firm heterogeneity by controlling for variations in the size of the compensating differential related to firm size or the presence of a union. Both firm size and unionism are commonly associated with the payment of wage premia and/or the presence of market power where the cost of fringe benefits to the firm may be less. Our results are consistent with these a priori expectations and suggest that the magnitude of the compensating differential is significantly higher in nonunion and in small firms.

The results from this paper are significant for two reasons. First, they provide evidence on heterogeneity in the size of the compensating differential which

[3]See Hersch and Viscusi (1990). Viscusi (1983) also looks at union-nonunion differences in the compensating differential for job risks.

may, in part, account for differences in estimates across studies. Second, they expand our knowledge of the relationship between pensions and wages which has potential implications for the shape of workers age-earnings profiles and for the distribution of income. Regarding the latter, pensions are a large and growing component of compensation and work by McCarthy and Turner (1983), Moore (1987) and Even and Macpherson (1990) suggest that pensions are important for understanding race and gender wage inequality. Further, work by Allen, Clark and McDermed (1991), Ippilito (1991), Lazear (1981) and others suggest that pensions serve as a bonding device as workers invest in deferred compensation and therefore reduce their turnover.[4] This may have the effect of raising productivity by reducing shirking or promoting investments in firm-specific human capital. Pensions, however, serve as a bonding device only to the extent a worker experiences a capital loss from changing jobs. In firms or markets where workers "pay" for their pension accruals on a one-for-one annual basis there are no capital losses associated with turnover. Finally, work by Gustman and Steinmeier (1987) suggests that compensating differentials for pensions can provide evidence on market segmentation and the importance of efficiency wage theories. These issues are discussed at greater length below.

In Section II we outline the theoretical or conceptual framework that guides the analysis. In Section III we discuss the data set that is used in the estimation and the econometric consideration that are crucial in attempts to identify both the market

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[4]See Allen, Clark and McDermed (1991).

hedonic and the underlying firm variations. The empirical results are also presented. Finally, in Section IV conclusions and implications are discussed.

II. Conceptual Framework

The hedonic model of compensating wage differentials suggests that in competitive labor markets there will exist an entire locus of long run equilibrium wage-pension combinations if workers and firms value pension benefits differently from wage benefits. For example, older workers, or those in high marginal tax brackets, or those with low rates of time preference, may prefer to receive a significant portion of their compensation as pensions. Similarly, the implicit price to the firm of a dollar of pension benefits may be different than the implicit price of a dollar of wages if, as previous research has suggested, pensions raise workers' productivity. To see this point, assume our firm has the following short run production function:

(1) $q = f(n\ e)$

$f_e > 0, f_n > 0$

where n is employment and e is effort.

Following Lazear (1981) we assume that effort is related to the amount of deferred compensation or pension the firm offers to workers.

(2) $e = e(P)$

$e'_P > 0$

where P = pension benefits.[5]

The profit maximizing level of labor compensation is:

(3) $e(P)f_n = w+P$

Totally differentiating (3) we can solve for the firm's pension-wage tradeoff:

(4) $$\frac{\partial w}{\partial P} = \frac{\partial e}{\partial P} f_n - 1$$

In markets where effort is not a function of deferred compensation $(\frac{\partial e}{\partial P} = 0)$ we get the standard prediction that the pension-wage tradeoff is one-for-one. The actual tradeoff will vary even in competitive labor markets to the extent pensions are used to prevent shirking. In effect pensions, by reducing shirking or turnover costs, have

[5] We are assuming here that only pensions affect effort but our model could easily be modified to allow both pension and wage effects on effort. As long as the effect of pensions on effort is greater than that of wages our predictions about the slope of the pension-wage tradeoff would be unaltered.

a lower implicit price to the firm than do wages.[6]

Thus, the extent to which increased pension benefits are offset with reduced wages will vary across firms or industries where turnover costs or shirking is an important consideration. One might expect that in markets or for firms in which firm-specific human capital or training costs are unimportant, or in which monitoring can be easily achieved, that increased pensions would be offset almost dollar-for-dollar by reduced wages. However, in markets where the reverse is the case then we might expect a smaller or even positive relationship between wages and pensions. Thus, the magnitude of the compensating differential should vary systematically across markets or firms where the productivity-enhancing effects of pensions are important.

Work by Oi (1983) and others suggest that turnover costs or monitoring problems should be more acute for large firms. Consequently, big firms may use deferred compensation more than small firms. Even and MacPherson (1991) present evidence that large firms are more likely to offer pensions to comparably skilled workers than are small firms. Of course, the mere presence of a pension at big firms need not imply any productivity effect from pensions. Brown and Medoff (1990) found that wages are also higher at big firms so, given the tax advantages of pensions for high-wage workers, it is not surprising that they are more likely to be

[6]Work by Salop and Salop (1976) and Akerlof and Katz (1988) also implies that total compensation will also be increased by pension productivity effects associated with worker screening and motivation. It is difficult to disentangle whether this represents an "efficiency wage" payment or simply an increase in the market clearing level of compensation per unit of enhanced productivity.

covered by a pension. The increased prevalence of pensions at big firms could thus simply reflect the preferences of higher wage workers for pensions. If tastes on the part of workers account for the correlation between pensions and firm size, however, there would be no reason for the magnitude of the tradeoff of pensions for wages to vary by firm size. Thus, to adequately test the notion that pensions are used as a productivity enhancement device one needs to look at whether workers or firms "pay" for these pensions via reduced wages. Firms will only pay for some or all of the cost of pensions if they do not have productivity-enhancing characteristics, so that the amount that the firm pays varies with the magnitude of the productivity-enhancing effect.

Gustman and Steinmeier (1987) have offered a related argument for why the pension benefits might vary by firm size. They suggest that pensions serve primarily as a form of efficiency wage premia and hence raise total compensation. They argue that it is the presence of wage premia not deferred compensation that generates the reduction in worker turnover associated with the presence of pensions. If labor markets are not competitive or firms pay efficiency wages, then it is unclear to what extent, if any, increased pension benefits in these firms would be offset by lower wage payments. Consequently, we might again tend to expect that the pension-wage tradeoff may be less than one-for-one in markets where compensation premia are present.[7]

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[7] Another reason for a tradeoff in this range is that, as Allen and Clark (1987) suggest, if pension plans are underfunded the calculated pension benefit may overestimate the short run cost of the plan.

Previous research has shown that the extent of coverage and level of pension benefits also varies across the union and nonunion sectors.[8] To the extent that unions' demands reflect the preferences of older more senior workers or overall compensation is higher, it is not surprising that pension wealth is higher in the union sector. The fact that the level of pension wealth is higher in the union sector need not imply that the compensating differential associated with pensions varies across union status. Only if the implicit price to the firm of pensions varies with union status will there be cross-section variation in the size of the compensating wage differential. There are several reasons, to believe that this might be the case. First, to the degree unions help defray the administrative costs of pensions, the implicit price of pensions to unionized firms will be reduced.[9] Second, if turnover costs (per worker) are higher in the union sector then pensions might play a bigger role in bonding workers to the firm. Of course since quits rates are higher for nonunion workers the expected cost of a dollar of pension benefits could well be less in the nonunion sector. In either case, turnover costs considerations will make the implicit price of pensions vary by union status. Finally, to the degree union provide better information on the nature of benefits they may serve to prevent firms from charging "too high" of a price for these benefits. Work by Viscusi (1983) has found that the size of the compensating wage differential for job risks is larger for union workers

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[8]See Freeman (1985) for evidence that suggests that unions alter the extent of coverage and the characteristics of pension plans.

[9]Unions pay the administrative costs of pensions in trucking, construction and mining industries. Thanks to Steve Allen for pointing this out to the authors.

compared to that received by nonunion workers. This result is interpreted to reflect either superior information or taste differences across union status. It does suggest that unionized workers face different compensating differentials than nonunion workers.

To test these predictions we estimate the market wage-pension tradeoff, or the hedonic price equation using the lag-linear functional form in recognition of the widespread evidence that wages are distributed log-linearly (and as used by Smith and Ehrenberg 1983). Let compensation equal productivity for individual i in market j where j=1,,3 for union, nonunion big-firm, and nonunion small-firm or:

(5) $W_{ij}(1+b_j p_{ij}) = A \exp(\delta' X_{ij} + e_{ij})$

where W is wage income, p is the ratio of pension income to wage income, X is a vector of individual characteristics, e is an i.i.d. error term, and productivity is distributed exponentially with technology shifts parameter A.[10] The coefficient b_j varies across markets and will equal one if there is a one-for-one pension-wage tradeoff. Taking natural logs and assuming that $\ln(1+p)=p$ for small values of p (and suppressing ln A in the constant term contained in X), the market HPE to be estimated is:

(6) $\ln W_{ij} = \delta' X_{ij} - b_j \, p_{ij} + e_{ij}$.

[10]Although more restrictive, this three-market model is similar in spirit to the dual labor market model developed by Dickens and Lang (1988). In their model they have a union sector and then a nonunion primary and secondary. Jobs at large firms may resemble those in the primary sector of a dual labor market model in the sense of offering high pay and being rationed. Thus, our framework could be thought of as one in which we explicitly sort jobs into the primary and secondary sectors based on firm size characteristics.

Potential biases in the estimation of this equation will be discussed below.

III. Empirical Results

The 1983 Survey of Consumer Finances collected extensive data on a nationally representative sample of 3824 randomly selected households. Of this total, 1066 households were matched to the Pension Provider Survey (PPS), which surveyed pension providers for those respondents who indicated that they were covered by a pension. Thus, this is one of the few data sets containing pension values for a randomly selected sample of workers. After eliminating observations for whom data is missing, or who were not working in 1983, the sample size is 3029 individuals.

The pension accruals were based on firm benefit formulas which incorporate firm-specific data on vesting age, Social Security offsets, early retirement provisions, profit sharing and type of annuity. The expected tenure with the current firm was based on the worker's response to expected quit date questions (so vesting occurs as long as their expected tenure exceeds the vesting period). In addition, the benefit calculations require a nominal interest rate assumption which we set at the 1982 thirty-year T-Bill rate of 10.85 percent. Expected inflation was assumed to average 6.85 percent in 1982 and real interest rates were 4.00 percent. Productivity growth was assumed to be 2 percent and the inflation and discount rates are the same as those for the pensions' expected present values. The value of the annual pension accrual is the difference in the expected present value of pensions in 1983 minus that in 1982.[11]

In Table 1 we present mean values of pensions accruals (relative to income) for the different sectors as well as means for the percent of the work force in each sector receiving a pension. For each skill (education) group we see that unionized workers are about twice as likely to receive pensions as the average worker. Further, among nonunion workers those in large firms (> 100 employees) are six to twelve times more likely to receive pensions than those in small firms. In fact, only about 2 percent of nonunion workers in small firms are covered by a pension. It should be noted that not only are unionized workers more likely to receive a pension but the value of their pension (as a percentage of income) is generally higher than those received by nonunion workers. Finally, these means suggest that pension coverage rises with skill or education but that the value of benefits relative to income is highest among those with 12-15 years of education.

Table 2 contains the results of estimating equation (6) that introduces sectoral differences in the compensating differential for pensions. It is possible that the returns to other worker characteristics may also vary across union status and firms size and if these are correlated with the presence of pensions we would get a biased estimate of the pension-wage tradeoff. Thus, we permit the displayed coefficients (in Table 2) to vary across sectors - those not displayed do not differ across sectors (differences were rejected by F-statistics).

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[11] The mean value of the annual pension accrual ($2293 for those having a pension) is comparable to those found by others in the literature (Smith and Ehrenberg, 1983; Moore, 1987).

$$\ln W_{ij} = \delta_j' X_{ij} - b_j \, p_{ij} + e_{ij}.$$

The education variables are years of education by subgroup (e.g., EDUC <12 is years of education for workers with less than 12 years of education).

The coefficients on the personal characteristics are the right signs and are consistent with those found in previous studies.

The estimates of the compensating differentials vary significantly across sectors ($F(2,3029)=4.31$) and are significantly different from zero. The values of the tradeoff are also consistent with the predictions of our theory, in that unionized workers face the smallest tradeoff while those nonunion workers in small firms face a tradeoff that is five to six times bigger. These result demonstrate that previous studies that limit samples to large firms (Moore 1987) or to unionized workers (Ehrenberg 1980) will underestimate the pension-wage tradeoff as a whole, partially explaining previous results that find small or negative differentials. Our results are also consistent with other studies that find wage premia in large firms and unions as higher wages are not totally offset by lower pension values. That is, unionized workers and those in big firms "pay" less for their pensions in terms of foregone wages. It should be emphasized that these results cannot tell us whether this is because of rent-sharing or pension induced productivity effects. They do suggest however that the value of the compensating differential does vary systematically across markets.

Several studies have noted that OLS estimation of the hedonic wage equation may yield biased estimates. Epple (1987) and Biddle and Zarkin (1988) suggest that omitted variable bias may affect the hedonic estimates if there are unobserved productivity factors that are correlated with the hedonic variable:

(7) $\quad \ln W_i = \delta' X_i - b\, p_i + u_i + e_i$

where u_i is an individual-specific intercept. If u_i and p_i are positively correlated, the estimated hedonic coefficient b will be biased upward.

However, in our specification of the HPE it is not possible to sign the direction of this potential omitted variable bias. Specify pension benefits as a function of wages:

(8) $\quad P_i = \gamma_1' Z_i + \gamma_2 W_i + \gamma_3 W_i^2 + v_i$

where P is the level of the pension benefits, W is wage income, and Z is a vector of other variables that affect pension benefits. Divide equation (8) by W to yield an expression for our pension variable:

(9) $\quad p_i = \gamma_1' Z_i/W_i + \gamma_2 + \gamma_3 W_i + v_i/W_i$

Because increases in W_i affect the pension variable both positively (γ_3) and inversely (γ_1' Z_i), it is no longer possible to sign the direction of the omitted variable bias imposed by the omission of u_i from equation (7). That is, if the cov(W,u)>0, the cov(u,p) will be positive or negative depending on the relative magnitudes of γ_3 and γ_1 and Z_i.

Smith and Ehrenberg (1983) and Moore (1987) discuss a second problem with the pension variable, arising from the technical bias in firms' pension benefit formulas. Benefits are generally calculated as a function of the worker's wage income in the years just prior to retirement:

$$(10) \quad P_i = k_i \, [E(W_i^f)]$$

where annual pension accruals are a positive function of the expected future wages, $E(W^f)$, and pension plan generosity, or replacement rate k. Because our pension measure, p, divides the pension level by the wage, the direction of the technical bias is uncertain for the same reason the direction of the omitted variables bias is uncertain.

To test for the possibility of technical and omitted variables bias, we instrument p in the wage equation (7) with variables that are unlikely to be correlated with these causes of bias. To reduce technical bias, the instruments for the pension value are: pension characteristics (the final year replacement ratio, vesting status); firm-specific characteristics that are likely to be correlated with the generosity of the pension plan (firm size and we match the 3-digit industry capital-labor ratio from the 1982 Census of Business and the 1982 Source Book: Statistics of Income); and measured individual variables in the wage equation, such as tenure.

Unfortunately, these variables explain pension values very poorly - the R^2 for a regression of nonzero pension values on the listed variables is only 0.025. The instrument list for omitted ability bias excludes the pensions characteristics lists (because the probability of pension coverage may be correlated with omitted ability), including only linear and nonlinear interactions between worker and firm characteristics. The R^2 for the pension regression on these characteristics is 0.015.[12] Not surprisingly, compensating differentials estimated with such poor instruments are universally insignificant. Given these lack of precision in the instrumental regressions and the fact that we cannot sign the directions of the potential biases in the pension coefficient, the OLS results remain our preferred specification.

Because of the potential importance of this omitted ability bias problem we investigate another avenue to try and determine the sensitivity of our results to this issue. It would seem that if unobserved ability was more important in unionized or large firms (and hence generating our results) then one might expect that the size of the compensating wage differentials associated with other job characteristics would also vary across sectors in the same way as we observe for pensions. To test this we included in our basic equation the number of days lost due to nonfatal injuries in each industry as a measure of on the job risks. As expected, the coefficients on the injury rate were positive and significant in each sector (see Appendix table 1). Further, neither the relative size nor the significance of the pension coefficients are affected by the inclusion of this other measure. Most importantly we could not reject the hypothesis that the compensating differential for risk were the same in each sector. Thus, although unobserved ability may be affecting the magnitude of our observed compensating wage differential, it seems that it is not generating the relative pattern that we observe across sectors. This pattern could instead reflect the

[12]Because pension data is only available for one year, we cannot first difference the wage regression.

differential productivity enhancing effects of pensions.[13]

There is another concern regarding the OLS estimates that warrants attention. OLS estimates may yield biased results to the that degree they fail to control for unobserved worker or sector attributes which affect both sectoral choice and sectoral income. That is, if the likelihood that a worker is employed in a large or union firm is correlated with the value of their pension we would get a biased estimate of the pension-wage tradeoff. To control for selectivity bias, we introduce the simultaneous estimation of sector-specific income equations and sectoral choice using the following two-step multichoice model:

$$(11) \qquad \ln W_{ij} = \delta'_j X_{ij} - b_j p_{ij} + e_{ij}.$$

$$(12) \qquad I^*_{ij} = \gamma' Z_{ij} + u_{ij}$$

where I is a latent polychotomous index of sectoral choice with endogenous determinants Z_{ij} that may contain elements of X; W_j is the wage that is observed

[13]As another check we also interacted our pension variable with education group in regression for large and unionized firms. If unobserved ability was generating our results we might expect all education groups within a sector to have the same compensating differential. As with our basic results the estimated differentials for each group were bigger in large nonunion firms than in union firms. Further, within each sector the compensating differential increased with education. Again this heterogeneity seems inconsistent with the hypothesis that unobserved ability is driving our results. These results are available from the authors upon request.

when the jth sector is chosen, and we assume that $E(e_j \mid X_j , Z_j) = 0$ (Maddala (1983)). A worker will be observed in the jth sector, (j = union, large nonunion, small nonunion) if it is the sector in which utility is highest or:

$$(13) \qquad I=j \ \ iff \ \ \acute{\gamma}Z_j-\acute{\gamma}Z_k>u_k-u_j \ \ for \ \ \begin{matrix} k=1,.,3 \\ k\neq j \end{matrix}$$

this implies

$$(14) \qquad I=j \ iff \ I_j^*>Max \ I_k^* \quad (k=1,..,3,k\neq j)$$

so that if

$$\epsilon_j\equiv Max \ I_k^*-u_j \quad (k=1,..,3 \ k\neq j)$$

then

$$(15) \qquad I=j \quad iff \quad \epsilon_j<\acute{\gamma}Z_j$$

Assuming ϵ_j is distributed i.i.d. with the extreme-value distribution with the cumulative distribution function:

$$(16) \qquad F(u_j<c) = \exp[-\exp(-c)].$$

then

$$(17) \qquad F_j(\epsilon) = Prob(\epsilon_j < \epsilon) = \frac{\exp(\epsilon)}{\exp(\epsilon) + \sum_{k \neq j}^{k=1,3} \exp(\hat{\gamma} Z_k)}$$

Using a procedure similar to that used in binary selection models, the two-stage method can be used to estimate the second-step wage equation.

$$(18) lnw_j = \delta_j' X_j - bp_{i_j} - \sigma_j \rho_j \, \phi[J_j(\gamma z_j)]/F_j(\gamma z_j) + \nu_j = \delta_j' X_j - bp_{i_j} - \sigma_j \rho_j \hat{\lambda}_j +$$

where $\sigma_j^2 = var(u_j)$, ρ_j is the correlation coefficient between u_j and ϵ_j^* ,

and $\epsilon_j^* = \Phi^{-1}[F_j(\epsilon)]$.

Now $\hat{\gamma}$ can be substituted for γ from the first stage multinomial logit model to estimate selectivity-corrected sectoral wage regressions (18).

The results from estimating the selection equations are presented in Table 3 while the new wage equations are presented in Table 4. The multilogit results conform to expectations: they demonstrate that older workers, or men with multiple children, or more educated workers, are more likely to hold jobs in large firms or in the union sector. However, the coefficients on the selectivity correction terms,

$\hat{\lambda}_j$, in the wage regressions are insignificant in all sectors.

Thus, after controlling for selectivity bias, we see that the magnitude of the pension-wage tradeoff is smallest in the union sector and largest for workers in small-nonunion firms, and all are significantly different from zero. The tradeoff is over five times greater for nonunion workers at small companies than it is for unionized workers and twice as big as it is for nonunion workers at big companies.

IV. Conclusions

In this paper we exploit the fact that the theory of compensating differentials can be used to identify sources of heterogeneity in firms' costs of providing fringe benefits and hence heterogeneity in the magnitude of the compensating differential. We estimate the relationship between pensions and wages taking into account the potential role of firm heterogeneity by controlling for variations in the size of the compensating differential related to firm size or the presence of a union. Both firm size and unionism are commonly associated with the payment of wage premia and/or the presence of market power where the costs of fringe benefits to the firm may be less. Our results are consistent with these *a priori* expectations and suggest that the magnitude of the compensating differential is significantly higher in nonunion and in small firms.

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Appendix Table 1

	NONUNION SMALL	UNION	NONUNION LARGE
PEN	-.074	-.168	-.491
	(-2.11)	(-2.04)	(-2.60)
DUR	.0018	.0014	.0019
	(6.23)	(4.73)	(3.69)

NOTE: DUR is the number of days absent per nonfatal injury that results in any loss of worktime of a day or more. Data are from BLS Occupational Injuries and Illnesses in the US by Industry and from Occupational Safety and Health Statistics of the Federal Government. The other included variables are the same as in Table 2.

Table 1

Mean Pension Values
(Pension as a % of Income)

	NONUNION UNION	NONUNION LARGE		SMALL	TOTAL
EDUC <12	.070	.096		.014	.082
	(.085)	(.165)		(.009)	(.123)
EDUC 12-15	.164	.066		.047	.120
	(.538)	(.117)	(.089)	(.385)	
EDUC 16+	.145	.102		.025	.106
	(.248)	(.198)	(.024)	(.201)	

Pension Coverage Rates

EDUC <12	23.7	12.4	1.0	12.0
EDUC 12-15	24.8	17.6	2.5	15.5
EDUC 16+	39.3	23.7	1.9	24.0

Table 2

ln (Income) Regressions

	NONUNION LARGE	NONUNION SMALL	UNION
PEN	-.096 (-2.43)	-.235 (-2.25)	-.511 (-2.76)
TENURE	.018 (2.63)	.023 (4.62)	.032 (3.36)
$TENURE^2$	-.0003 (-1.72)	-.0003 (-2.08)	
EXPER	.021 (3.27)	.029 (7.19)	.022 (4.78)
$EXPER^2$	-.0004 (-3.04)	-.0005 (-5.05)	-.0004 (-4.36)
MALE	.205 (6.89)	.188 (6.67)	.237 (5.22)
BLACK	-.126 (-2.91)	-.156 (-3.74)	-.215 (-3.81)
HOURS	.821 (16.75)	.908 (44.85)	.785 (33.02)
EDUC <12	.026 (1.67)	.030 (2.39)	.083 (5.49)
EDUC 12-15	.031 (2.56)	.038 (4.09)	.074 (6.55)
EDUC >16	.029 (3.04)	.048 (6.47)	.065 (7.12)

R^2 = .682

Other variables in the regressions include marriage, city size, four regions, and five occupational dummies. T-statistics are in parentheses.

Table 3

Multinomial Logit Selection Equations

	UNION VERSUS NONUNION, SMALL	NONUNION, BIG VS. NONUNION, SMALL
EDUC <12	.050 (1.03)	.030 (0.69)
EDUC 12-15	.082 (2.27)	.054 (1.64)
EDUC 16+	.116 (4.83)	.080 (3.05)
EXPER <5=1	-1.27 (-7.17)	-.371 (-2.64)
EXPER 5-9=1	-.85 (-5.72)	-.283 (-1.36)
EXPER 10-14=1	-.512 (-3.33)	-.190 (-1.36)
EXPER 15-19=1	.014 (0.09)	-.021 (-0.13)
MANAGERS	-1.12 (-6.25)	.346 (2.23)
TECH,SALES,CLERICAL	-.74 (-4.62)	.170 (1.17)
SERVICE	-.45 (-2.48)	.091 (0.55)
CRAFT	.023 (0.14)	-.174 (-1.06)
BLACK	.73 (4.13)	.055 (0.38)
MARRIED	.037 (0.02)	-.08 (-.72)
MALE	-.150 (-1.09)	-.107 (-0.96)
NO. OF CHILDREN, MALE	.164 (2.58)	.134 (2.23)
NO. OF CHILDREN, FEMALE	.05 (0.71)	.041 (0.72)
HEALTH IS POOR	.04 (0.58)	-.036 (-0.53)

NORTHEAST	.30 (1.83)	.370 (2.47)
NORTHCENTRAL	.26 (1.67)	.157 (1.11)
SOUTH	-.319 (-2.00)	.543 (3.91)
CITY LARGE	.194 (1.71)	.080 (0.78)
CONSTANT	-.390 (-0.76)	-.348 (-0.75)

LOG Likelihood: - 3195

Table 4

ln (Income) Regressions
(with selectivity correction)

	NONUNION LARGE	NONUNION SMALL	UNION
PEN	-.094 (-2.41)	-.217 (-2.21)	-.526 (-2.74)
TENURE	.017 (2.75)	.023 (4.58)	.032 (3.39)
$TENURE^2$	-.00035 (-1.85)	-.0003 (-2.04)	
EXPER	.024 (3.18)	.033 (6.72)	.026 (4.02)
$EXPER^2$	-.00046 (-2.94)	-.0005 (-5.21)	-.0005 (-3.73)
MALE	.205 (6.89)	.188 (6.65)	.230 (5.01)
BLACK	-.080 (-2.90)	-.111 (-3.30)	-.188 (-3.35)
HOURS	.824 (16.78)	.900 (39.18)	.793 (28.30)
EDUC <12	.026 (4.69)	.028 (2.25)	.090 (5.13)
EDUC 12-15	.032 (2.66)	.035 (3.61)	.082 (5.94)
EDUC >16	.031 (3.10)	.046 (6.17)	.073 (6.85)
λ	-.088 (-.079)	.353 (1.81)	.158 (0.84)
R^2	.552	.716	.614

t-statistics are in parentheses. Other variables in the regressions include marriage, city size, four region and five occupational dummies.