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capital-intensive firms. Such firms may encourage production efficiency and worker attachment by paying higher wages. However, if part-time jobs are inherently less important to their holders than full-time jobs, the wage increase necessary to attach part-timers may be greater than that for full-timers, making the former relatively more costly. Without further data we cannot disentangle these effects. Similarly, special characteristics of the people who supply part-time work may explain some of the effects detected in this study. Most part-time workers are married women or teenagers (Deuterman and Brown, 1978). These workers may tend to prefer flexible work schedules or loose attachment to the firm, arrangements which are often more easily obtained in the comparatively informal atmosphere of small firms. Our data did not provide evidence on this and other aspects of the work environment, some of which may influence the sorting of part-time and full-time workers among firms.[3] The inability of this study to address all of these issues illustrates the paucity of information about the demand side of the part-time labor market.

[3] Several suggestions about alternative interpretations of the results were made by an anonymous referee whom I wish to thank.

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LABOR TURNOVER BIAS IN ESTIMATING WAGES

John J. Beggs and Bruce J. Chapman*

Abstract—Cross-sectional earnings function analyses typically do not recognise the potential for misrepresentation arising from a relationship between unobserved ability and the probability of labor turnover. Our point is that if individuals staying with the firm for relatively long periods of time are more, or less, able on average than recent hires, the coefficient estimated on tenure is necessarily biased. This is an important issue for wage determination modelling given the relevance of tenure to wages. In this note we examine the theoretical basis of our claim, and propose and implement a solution to the problem with a novel use of instrumental variables on a large sample of workers employed in the Australian government.

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I. Introduction

In this note it is postulated that previous cross-sectional research using the earnings function is inadequate in its treatment of the relationship between unobserved ability and the probability of labor turnover. This relationship implies that individuals observed at a point in time are an unrepresentative sample of the cohort originally joining the firm, because of the nature and incidence of past quits and permanent lay-offs. Ordinary regression estimates of the earnings function are therefore inconsistent. In particular, this is the case for the time-on-the-job coefficient, one of the more important

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parameters in the analysis of, for example, sex and race wage differentials. A novel use of instrumental variables estimators is proposed as a solution to the problem.

The note proceeds as follows. Section II presents a brief and selective review of studies relevant to the analysis and examines the theoretical issues. In section III a solution to the problem is explained and as an example the technique is applied to a large sample of Australian government workers; the results are reported in section IV. A concluding section summarises the major findings.

II. An Overview of the Literature

Bloom and Killingsworth's (1982) primer on the use of regression in pay discrimination research identifies the following problem. They refer to the testimony of a statistical expert witness, Dr. Paul L. Meier, in the trial of Presseisen v. Swarthmore (1977), who

> questioned the reliability of cross-section studies of Swarthmore faculty who were employed at Swarthmore as of a given date on the grounds that such studies necessarily omitted "inactives"—faculty who were employed as of some previous date but were no longer employed as of the date referenced by the study. (p. 337)

Our goal is to examine the implications of this view.

The issue is that because of attrition some proportion of the original cohort of workers joining a firm will have left by the time the cohort is observed in the cross-section. This constitutes a problem for econometric investigation of wage determination if workers' unobserved ability is not independent of the turnover decision. If this is the case, there is a correlation between unobserved ability and time in the firm (tenure). That is, as the original cohort ages, workers remaining become increasingly more homogeneous and increasingly less representative of the original group as either high or low unobserved ability workers quit. A form of selectivity bias is present and the parameter estimated for tenure in the wage equation is inconsistent.

For want of an adequate methodological treatment, research has proceeded in this area as if the problem did not exist. For example, Blinder (1974), Filer (1983), Abowd and Killingsworth (1983) and Johnson and Lambrinos (1985) attempt to explain wage differentials between groups without reference to the effects of nonrandom quitting. All three analyses include tenure exogenously in the wage estimation, with the variable found to have significant explanatory power.

The problem applies also to less specific applications of the earnings function. For example, Medoff and Abraham's (1981) rejection of the efficacy of human capital theory in an explanation of earnings relies substantially on their finding that company experience is a significant salary determinant holding constant direct measures of productivity. But because the turnover selection process can generate upward or downward bias in returns to experience (Beggs and Chapman (1986)) their exogenous treatment of tenure implies their case is yet to be established.

No wage determination exercise can avoid the problem, since even investigations using time in the labor force are inappropriate. This is because this variable is likely to be positively correlated with tenure, an issue analysed by Chapman and Tan (1980). Recently, Beggs and Chapman (1986) have developed a model of worker self-selection which has important implications for the issue. The essence of this framework is as follows.

One way of understanding job turnover is to assume that firms and workers are initially poorly informed about the suitability of the worker in the current job, but over time both parties become more aware of the appropriateness of continuing the employment relationship. This is the essence of job mismatch models (Jovanovic (1979)) in which jobs are "experience" goods, where agents learn of the match of their ability with particular job requirements through having experience on the job. Extra dimensions can be added to this framework. They are that, over time, workers gain, (i) greater insight into their general ability through experiencing a particular job, and (ii) knowledge of which employers reward their type of ability more.

The basic issues of this model can be illustrated simply, with the formal analysis being available in Beggs and Chapman (1986). Workers' wage streams are made up of an initial wage plus a payment for intangible ability. At the point of first hire the wage offer is for measured ability only, and wages will be relatively high, initially, in the first job. Through time employers learn about workers' intangible ability and offer wages to them commensurately. Given and alternative wages will vary in such a way as to encourage or discourage labor turnover.

The firm's cross-sectionally observed wage structure is affected by which workers have quit in the time between hiring and sample observation. If workers with relatively low ability leave, the situation embodied in conventional job mismatch theory, the cross-sectionally observed wage–tenure profile is steeper than if no quitting occurred. The converse is true if high ability workers leave, a case resulting from the original firm paying a relatively low return to ability. The two possibilities mean that worker self-selection has ambiguous implications for wage equations estimated on job stayers. The issues are addressed empirically below.

III. Econometric Methodology

The essential point is that if workers learn of their intangible worth with time on the job and, on the basis

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of that knowledge decide whether or not to quit, the usual assumptions of the regression model are not satisfied. If better workers (those with high intangible skills) quit, this induces a negative correlation between the disturbances and the tenure variables, resulting in downward biased estimates of the regression coefficients on tenure. The opposite is true if poorer workers are more likely to quit.

In this section a novel use of the instrumental variables (IV) technique for estimating the wage equation where the available data have been subjected to sample self-selection is explained. To demonstrate the strategy take the wage equation

$$W = X\beta + u \tag{1}$$

and rewrite it in deviation form as

$$w = x\underset{\sim}{\beta} + (u - \bar{u})$$

where $\bar{u}$ is a vector containing the means of the in-sample disturbances and the parameter vector is partitioned as:

$$\beta = \begin{bmatrix} \beta_0 \\ \underset{\sim}{\beta} \end{bmatrix}$$

with β_0 the intercept term and $\underset{\sim}{\beta}$ the slope coefficients. The slope coefficients can be estimated consistently by identifying instruments, Z, which are asymptotically uncorrelated with the in-sample disturbances. That is:

$$\operatorname{plim} Z'(u - \bar{u}) = 0. \tag{2}$$

This requirement is slightly different from the usual condition placed on instruments, namely that they be asymptotically orthogonal to the true disturbances of the regression model. The importance of this distinction is that in the IV procedure proposed here the regression intercept term is not consistently estimated by:

$$\hat{\beta}_{0IV} = \overline{W} - \overline{X}\underset{\sim}{\hat{\beta}}_{IV} \tag{3}$$

where $\overline{W}$ and $\overline{X}$ are the means of the wage and explanatory variables, respectively. This is because $\overline{W}$ is not the unconditional mean of the dependent variable in the absence of sample self-selection. In many IV applications the slope coefficients are the central concern and inability to estimate the intercept consistently is of no consequence. But in some exercises, such as comparisons of male–female or black–white wage differentials, it is important to predict the average wage level without self-selection. In cases of this type it is central that the intercept term in the wage equation be estimated consistently.

In the problem studied here, a simple procedure is proposed. The idea is to select a region of the regressor space in which no turnover self-selection is present and re-estimate the model. The estimate of the intercept term from this subsample regression is then combined with the previously estimated $\hat{\beta}_{IV}$ slope coefficients to provide a consistent overall set of regression coefficients.[1]

In the situation at hand the economic model implies that the sample self-selection of the type discussed is related to tenure. This means that the wage equation can be estimated consistently only for that part of the sample where tenure equals zero. But, obviously, this regression does not reveal how the firm rewards its employees for tenure, a coefficient which is obtained consistently from the IV estimates.

Taking the intercept term from the subsample regression and combining it with IV estimates of the slope coefficients is not the most efficient available procedure. Since the intercept term itself is a long way from the mean of the data, it is estimated with a considerable amount of sampling variability. Hence a slightly different method was employed, which is now explained.

Let $\overline{W}^S$ and $\overline{X}^S$ denote the sample means of the wage and explanatory variables in the subsample not subject to turnover selectivity bias, and let $\hat{\beta}_S$ be the ordinary least squares parameter estimates from this sample. Denote

$$\overline{W}^S = \overline{X}^S\hat{\beta}_S. \tag{4}$$

The predicted wage at the mean of the subsample is simply the average wage in the subsample. Denote

$$\hat{W}^S_{IV} = \overline{X}^S\hat{\beta}_{IV} \tag{5}$$

where $\hat{W}^S_{IV}$ is the predicted wage from the IV estimates evaluated at the mean of the regressors in the subsample not subjected to self-selection. Asymptotically the estimable slope coefficients $\hat{\beta}_S$ and the corresponding elements of $\hat{\beta}_{IV}$ are identical. The difference between $\overline{W}^S$ and $\hat{W}^S_{IV}$ can only be due to the asymptotic inconsistency in the estimate of the intercept coefficient. That inconsistency is estimated here as

$$\operatorname{est}(\hat{\beta}_{0S} - \hat{\beta}_{0IV}) = \overline{W}^S - \hat{W}^S_{IV} \tag{6}$$

which is the difference between the mean wage in the subsample and the predicted wage by the IV procedure at the mean of the subsample. The actual predicting equation used in the analysis is thus:

$$\hat{W}_i = \hat{\beta}_{0IV} + (\overline{W}^S - \hat{W}^S_{IV}) + \hat{\beta}_{1IV}X_{1i} + \cdots + \hat{\beta}_{kIV}X_{ki}. \tag{7}$$

[1] The suggestion is trite in the case where there is a large enough subsample of the regressors, where there is no self-selection and where there is sufficient variability in the regressors themselves. In such a case one would only run the regression on the subsample not subject to self-selection and not be concerned with the IV procedure.

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TABLE 1.—STATISTICAL CHARACTERISTICS OF THE DATA

	Men		Women	
	Mean	Standard Derivation	Mean	Standard Derivation
SAL ($A)[a]	4236.80	1362.80	3040.40	931.10
AGE (years)	25.90	6.47	22.32	4.07
YOS (years)	12.47	1.12	12.28	0.90
OJEX (years)	4.15	3.95	2.19	2.50
TEXP (years)	8.43	6.38	4.56	3.70

[a]To facilitate understanding of the orders of magnitude involved, note that in 1969 $A1 = $US1.1. From 1969 to 1986 Australian wages increased about six-fold in nominal terms.

IV. An Application

The data chosen to illustrate the issue are the total set of full-time Australian government workers all employed in 1969 in what is essentially the same job, the clerical/administrative (CA) occupation. These are very useful data for several reasons. First, the ability requirements for success in the job are fairly clear-cut. They are skills concerned with the preparation of government reports and management of programs and personnel. The implication is that the usual misspecification associated with applying a single earnings function to workers with different human capital formation paths is avoided. Secondly, there are a large number of observations allowing fairly confident statistical inference. The point of the exercise is to illustrate the potential importance of the theoretical and econometric issues with cross-section data, those commonly used for earnings function analysis.

To focus on a group most likely to have been involved in job shopping, the analysis was restricted to workers aged less than 45 and with less than 16 years experience in the CA. This resulted in a sample of 14,870 men and 4,941 women, with information available on age (*AGE*), years of schooling (*YOS*),[2] time in the CA (*OJEX*) and annual salary (*SAL*).[3] The availability of age and *YOS* information allowed the calculation of total time in the labor force, *TEXP*, given by (age minus *YOS* minus 5). The interpretation of *TEXP* as total time in the labor force assumes that all schooling was completed before labor force participation began, that work immediately followed the completion of schooling, and that labor force participation was continuous once commenced. For women in particular this results in important econometric and interpretation difficulties (Chapman (1985)). Consequently, women's *TEXP* was adjusted downwards slightly in order to gain a closer approximation of true experience.[4]

The statistical characteristics of the data are shown in table 1.

These data allowed the estimation of several models, all of which are variants of the traditional wage equation. Several modifications were incorporated to the usual form. First, for flexibility the equation was estimated with fourth order polynomials in the tenure and general labor market experience variables.[5] Second, a quadratic was included on the schooling variable to allow non-linearities. Third, returns to formal education were allowed to depreciate by interacting schooling (quadratically) with length of time since its completion. This permits a test of the hypothesis that formal schooling becomes less important as a promotion determinant as the vintage of education increases.

The above innovations led to the estimation of the following model:

$$\begin{aligned}\log(SAL_i) = {} & \alpha_0 + \alpha_1 YOS_i + \alpha_2 YOS_i^2 + \alpha_3 OJEX_i \\ & + \alpha_4 OJEX_i^2 + \alpha_5 OJEX_i^3 \\ & + \alpha_6 OJEX_i^4 + \alpha_7 TEXP_i + \alpha_8 TEXP_i^2 \\ & + \alpha_9 TEXP_i^3 + \alpha_{10} TEXP_i^4 \\ & + \alpha_{11} TEXP_i \cdot YOS_i + \alpha_{12} TEXP_i^2 \\ & \cdot YOS_i + \epsilon_i. \end{aligned} \tag{8}$$

[2]The years of schooling variable was computed by inferring time equivalents from educational qualifications.

[3]Annual salary data as such were unavailable from the sample. They were inferred by taking the midpoint of the (fairly narrow) salary range of each individual's job class.

[4]The adjustment estimate was calculated from the data under a set of restrictive assumptions. Time in the labor force before joining the APS was assumed to overstate true experience by $1/\alpha$, with $0 < \alpha < 1$, where α is the proportion of women remaining in employment each year from 1969 to 1974. Given that it is possible to identify which individuals were not employed in the CA in 1974 but nevertheless were in 1969 an approximation of α is given by $(1 - x/y)/5$ where x is the number of leavers by 1974 over the original population y. In other words, the average general labor market experience accumulation of women prior to 1969 is assumed to be given by the proportion of each year, on average, that the original population remained in employment. For the example, α is 0.904.

[5]We adopted this approach because, while the results of estimating the usual earnings function were statistically satisfactory — in terms of explanatory power and significance tests — a particular characteristic was discomforting. Predicted

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TABLE 2.—OLS AND IV ESTIMATES OF THE WAGE EQUATION

Variable	Men		Women	
	OLS	IV	OLS	IV
INTERCEPT	5.508	4.953	3.647	4.521
	(21.28)	(9.71)	(4.94)	(3.39)
OJEX	0.122	0.244	0.109	0.185
	(28.26)	(2.26)	(15.41)	(1.12)
$OJEX^2$	−0.0220	−0.0382	−0.0174	−0.0378
	(15.65)	(1.19)	(6.18)	(0.59)
$OJEX^3$	0.00190	0.00224	0.00124	0.00389
	(11.90)	(0.67)	(3.37)	(0.52)
$OJEX^4$	−0.0000574	−0.0000411	−0.0000291	−0.000134
	(9.98)	(0.37)	(1.99)	(0.49)
TEXP	0.260	0.291	0.369	0.320
	(23.55)	(1.60)	(16.63)	(2.40)
$TEXP^2$	−0.0218	−0.0322	−0.0345	−0.0298
	(27.70)	(2.54)	(20.25)	(3.15)
$TEXP^3$	0.000820	0.00170	0.00136	0.00132
	(21.36)	(3.45)	(14.38)	(3.00)
$TEXP^4$	−0.0000129	−0.0000283	−0.0000236	−0.0000204
	(17.80)	(3.04)	(11.14)	(2.37)
YOS	0.192	0.223	0.428	0.276
	(5.03)	(3.59)	(3.92)	(1.43)
YOS^2	−0.00214	−0.00318	−0.0101	0.00455
	(1.54)	(1.11)	(2.54)	(0.64)
$TEXP \cdot YOS$	−0.00617	−0.000303	−0.0119	−0.00590
	(7.69)	(0.02)	(6.98)	(0.60)
$TEXP \cdot YOS$	0.000253	0.000210	0.000576	0.0000768
	(7.12)	(0.32)	(6.27)	(0.15)
R^2	0.684	0.200	0.696	0.160
F	2681.6	310.1	940.9	77.4
N	14,870		4,941	

Note: Absolute t-statistics are in parentheses.

The predictions of this model are that $\alpha_1, \alpha_3, \alpha_5, \alpha_7, \alpha_9, \alpha_{12} > 0$ and $\alpha_2, \alpha_4, \alpha_6, \alpha_8, \alpha_{10}, \alpha_{11} < 0$.[6]

The IV procedure was made operational by the assumption that the learning associated with intangible ability is solely a function of tenure. Thus, the self-selection bias in the wage equation is ascribed to the correlation between the disturbance of the wage equation and the tenure variables. The wage equation disturbances were assumed to be asymptotically uncorrelated with educational qualifications and outside experience. A caveat is that learning about intangible ability may be associated with these variables. However, if the dominant effect is due to the learning associated with tenure the assumption provides a workable basis on which to proceed.

The IV method used was as follows. All the variables in the basic wage equation which contain tenure ($OJEX^{1-4}$, $TEXP^{1-4}$, $TEXP \cdot YOS$ and $TEXP^2 \cdot YOS$) were instrumented from the schooling and outside experience variables. The generalised instruments were obtained by reference to the unit record data on the type of educational qualification held by each individual (e.g., high school, 3-year Bachelor's degree, Honor's degree, Master's degree, etc.) and using 16 dummy variables for each year of "outside" work experience (*TEXP* minus *OJEX*). The correlation between tenure and the schooling and "outside" experience variables is relatively low, hence the final instrumental variables estimates of the regression equation are subject to a reasonable amount of sampling variability.[7] The OLS and IV estimates of the wage equation for men and women are presented in table 2.

The OLS results reported in table 2 are as predicted. The coefficients on the experience and schooling vari-

wages turned down at tenure levels of around 10 years, and did so rapidly after this point. While a turning point is commonly found in cross-sectional estimates, the rapidity of decline was unsatisfactory. We recommend this type of diagnostic testing for other researchers in the area.

[6] The conventional model in the literature sets $\alpha_2 = \alpha_5 = \alpha_6 = \alpha_9 = \alpha_{10} = \alpha_{11} = \alpha_{12} = 0$. Results for this and an intermediate range of models are available from the authors on request.

[7] The R^2 statistics for the regression of the tenure variables on the 16 outside experience dummies and the 9 schooling dummies lie between 0.06 and 0.20. In cross-sectional studies it is rarely possible to find the highly correlated instruments which are familiar in time series macroeconomic applications.

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TABLE 3.—PERCENTAGE INCREASE IN SALARY ASSOCIATED WITH A ONE-YEAR INCREASE IN INDEPENDENT VARIABLES, CALCULATED AT THE MEAN (FROM TABLE 2)

Variable	Men		Women	
	OLS	IV	OLS	IV
OJEX	5.09	9.27	6.45	10.31
TEXP	4.20	4.79	7.21	8.29
YOS	11.46	11.77	14.38	13.92

TABLE 4.—OLS SELECTION-FREE AND IV PREDICTED STARTING WAGES

Men			Women		
(i) Average Starting Wage, W^S	(ii) IV Predicted Starting Wage W^S_{IV}	(iii) Adjustment Factor (i − ii)	(i) Average Starting Wage, W^S	(ii) IV Predicted Starting Wage, W^S_{IV}	(iii) Adjustment Factor (i − ii)
2770.4	2314.5	455.9	2289.7	2257.6	32.1

ables display concavity, having negative signs on the even-order polynomials. Interacting schooling with time since entering the workforce shows the anticipated depreciation with depreciation occurring at a decreasing rate. All other coefficients give plausible results and their implied sizes are reported in table 3. The R^2s of 0.68 and above are high, indeed unusually so, for such cross-sectional estimations.

The IV coefficient estimates of the wage equation are similar in sign to those found for the OLS although the standard errors are large, a result probably attributable in part to the narrow set of instruments available. The important finding concerning the relative parameter sizes is stressed below.

The OLS and IV estimates reveal that, for these data, treating tenure exogenously leads to a substantial underestimate of the true effect on wages of experience on-the-job. At the mean for males and females, respectively, a one-year increase in tenure increased wages by 9.3% and 10.3% in the IV estimation, compared to 5.1% and 6.5% from the usual technique. The results thus indicate that workers of both sexes with relatively high levels of intangible ability were more likely to quit. Clearly, for these data, labor turnover was not random. This suggests that earnings functions estimated as if tenure is exogenous have within them the scope for serious misinterpretation of relationships.

The result of table 3 implies the possibility of important differences in the OLS and IV-corrected predictions of average wages. The adjustment necessitated on the intercept term of the IV equation to ascertain the true gap proceeded as follows.

The OLS equation (8) was estimated for the sample of men and women workers for whom *OJEX* equals 0, using the mean of the data for other variables, to derive a selection-free estimation of the starting wage. These data were used to adjust the IV equation by the amounts shown in table 4.

The results of table 4 demonstrate that, for these data, the IV technique understates the selection-free starting wage by about $455 and $32 per year for men and women, respectively. That there is an understatement is consistent with the finding that the IV *OJEX* slope coefficients are higher that the OLS *OJEX* slope coefficients for both men and women.

The estimates may now be used to investigate the empirical significance of the inappropriate application of OLS to the question of predicted average wages, an important issue for the measurement of sex and race discrimination. The OLS and IV-corrected wages predicted for males are $4,237 and $4,693, respectively, and for females the figures are $3,040 and $3,072. That is, the incorrect method understates true male and female predicted wages by 10.8% and 1.1%, respectively. This implies that the male–female average wage differential is increased from $1,197 to $1,621, or from 39.4% to 52.8%. This is an important variation highlighting the relevance of the suggested correction for these data. We stress that results could be quite different for other samples.

V. Conclusion

It has been argued that if a relationship exists between labor turnover and intangible ability, the OLS parameter estimated between tenure and wages is inconsistent. No a priori predictions can be made about the sign of the bias. For the data examined it was found that relatively able men and women workers were more likely to quit, a result implying that OLS cross-section estimations not correcting for this type of self-selection systematically underestimate the returns to on-the-job

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experience for both sexes. Other results may be true for different firms.

A technique was developed to correct the wage equation and estimates of predicted average wages for the presence of non-random quitting. For the sample considered several points stand out. First, OLS rates of return to on-the-job experience were substantially less than those derived with the selectivity-corrected process. Secondly, the OLS technique understates importantly the male–female wage gap predicted at the mean. Finally, the recommended technique allowed insight into the labor turnover process, a possibility previously not exploited with cross-section data.

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A JOB-CHARACTERISTICS THEORY OF RETIREMENT

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Abstract—This paper argues that certain fundamental job characteristics determine both retirement and pensions accumulated in anticipation of retirement. Data on job characteristics, average retirement ages on different jobs, and pensions are used to test this hypothesis. The results show a statistically significant relationship between job characteristics such as intense physical demands, stress, and repetitive working conditions, and early retirement. The results also show that retirees from jobs with such characteristics have relatively high pensions at early retirement ages, suggesting that their employment contracts anticipated early retirement. Thus, the causal relationship between retirement and pensions may be the opposite to the one normally accepted.

Introduction

Considerable econometric attention has been focused recently on the relationship between pension benefits and retirement age.[1] Most of this work has treated the characteristics of a worker's pension plan exogenously, without inquiring into the determinants of the structure of such plans. Implicitly, the causal arrow between pensions and retirement has been assumed to point from pensions toward retirement.

In a general equilibrium model of work, compensation, and retirement, causation may well run from desired retirement age to pensions. If jobs involve differing working conditions, and if human productivities in the different dimensions of work and tolerances for various conditions change unevenly with age, then the expected retirement age will differ from job to job. Now, if tax advantages and other factors lead workers to finance retirement primarily through pensions, then the pension plans associated with different jobs will also

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[1] Major recent econometric studies of the retirement decision include Gordon and Blinder (1980), Hall and Johnson (1980), Mitchell and Fields (1984), Hamermesh (1984), Fields and Mitchell (1985), and Gustman and Steinmeier (1985b). All of these studies show that retirement behavior is significantly affected by pension structure.

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