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ABSTRACT

A well-known, if underappreciated, finding in the mobility literature is that turnover is much lower in jobs covered by pensions than in other jobs. This could result from capital losses for job changes created by most benefit formulas, the tendency of turnover-prone individuals to avoid jobs covered by pensions, or higher overall compensation levels in such jobs. A switching bivariate probit model of pension coverage and turnover is developed to estimate the effect of each of these factors. The results show that capital losses are the main factor responsible for lower turnover in jobs covered by pensions, but self-selection and compensation levels also play an important role. This is the first direct evidence that bonding is important for understanding long-term employment relationships.

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I. INTRODUCTION

Each month between 4 and 5 percent of all workers in manufacturing leave their jobs, a fact frequently cited as evidence of the tremendous amount of flexibility in the U.S. labor market. Yet among workers who are age 30 or above, Hall (1982) has shown that 40 percent are working in jobs that will ultimately last 20 years or more. This apparent paradox is easy to reconcile. Most young workers go through a period of job shopping before they settle into a long-term job. Hall finds that the typical American worker will eventually hold 10 jobs, eight of them by age 40.

Although this profile of mobility over the life-cycle has been well documented in the literature, the determinants of long-term employment contracts are still not well understood. One simple explanation is mover-stayer heterogeneity. The odds of job changing can vary across workers because of factors such as attitudes and on-the-job behavior which are unobservable to the person analyzing the data. Although heterogeneity is no doubt an important factor, it is highly unlikely that it is the only factor influencing job mobility because it assigns a passive role to employer behavior.

Economic models of long term contracts focus on two mechanisms. Long-term jobs can result from decisions by firms to pay workers more than what they can earn elsewhere. A variety of rationales for such behavior have been offered in the literature, including (but not limited to) economizing on hiring and training costs (Becker (1975)), preventing shirking (Shapiro and Stiglitz (1984)), and optimal matching of workers and jobs (Jovanovic (1979)). The other mechanism that can produce long-term employment relationships is bonding. Up-front fees or bonds are rarely observed in modern American labor markets, but steep age-earnings profiles and deferred compensation are equivalent to bonds in terms of their effects on behavior. Bonding discourages quits (Salop and Salop (1976)) and layoffs for shirking (Lazear (1979,1981)) by imposing exit costs on workers who leave the firm. In return, the worker receives higher lifetime earnings and a commitment from the firm that he will not be terminated until the end of the contract, even though in some periods he is paid more than his output. This promise is enforced in the labor market by the firm's concern about its reputation (and ability to write such contracts in the future).

Although there are literally dozens of studies of mobility in the economics literature, no empirical study has examined whether bonding directly lowers turnover probabilities and most studies ignore nonwage compensation. The purpose of this paper is to start filling this gap in the literature. Under reasonable assumptions, defined benefit pension plans have the same characteristics as bonds in the Lazear model. This paper estimates the bonding component of pensions and shows how it is related to mobility.

The relative impact of wage premiums and bonding on job duration is an important issue with both theoretical and policy implications. The leading criticism of efficiency wage models has been that the same productivity gains could be obtained through bonding without the adverse side effects that supracompetitive wages have on employment and hours[1]. Evidence concerning

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[1] In response, advocates of the efficiency wage approach have pointed to sizable firm expenditures on monitoring as evidence that bonding must be limited by some constraints. These issues are discussed in Katz (1986) and Murphy and Topel (1988).

the magnitude of bonds and the impact of such bonds on worker behavior would indicate whether this criticism of efficiency wage models has any merit.

Pension bonding is also relevant to the financial question of what explicit and implicit obligations the employer faces under the pension contract. Bonds are not observed directly and do not even exist under a strictly legal interpretation of the pension contract (Bulow (1982)). Evidence that estimates of bonds are correlated with worker behavior (in this case mobility) would strengthen the empirical case that the employer's implicit liabilities exceed those on the balance sheet.

One policy concern is that labor market mobility has fallen below the efficient rate because of the pension bond that workers lose if they leave the firm. This is not a new concern; it has been with us at least since the time Ross (1958) examined the "new industrial feudalism" issue. The recent emphasis on insufficient mobility is attributed, in large part, to structural changes in the U.S. economy and the accompanying shifts in the demand for labor. The latest evidence on mobility trends compiled by Murphy and Topel (1987) shows that the rate at which workers move across different sectors of the economy has declined considerably since 1974. Some have suggested that pension regulations be changed to encourage more mobility. Another controversial policy issue is the question of whether workers are entitled to a share of any excess assets from a terminated pension plan as repayment for bonds on which the plan has defaulted.

To increase our understanding of the process through which workers enter long-term employment relationships, the paper begins by estimating job retention rates in the May 1979 and 1983 Current Population Surveys (CPS). Hall has used similar data to show how tenure is related to age, sex, and race; Hashimoto and Raisian (1985) have used the 1979 CPS to show how tenure is related to firm size. Our results show that pension coverage has been underappreciated as a determinant of job duration; it could very well be the most important factor.

Why do workers stay so long in jobs covered by pensions? One possibility is that firms with pensions pay more total compensation than other firms, resulting in lower quit rates. Second, salary-based defined benefit pension formulas impose sizable capital losses on those who leave. The capital loss not only discourages quits, but also lowers layoffs, at least among firms concerned about their labor market reputations. In addition, the bonding mechanism acts as a self-selection device to screen out workers who are likely to quit or be fired. After discussing these competing explanations in some detail, this paper develops an econometric model to estimate the relative importance of each factor in the decision to remain on the job.

II. PENSIONS AND JOB DURATION

Although the issue of how pensions affect mobility has been studied before, the emphasis has been on estimating a pension coverage intercept in a turnover equation. The questions of whether the impact of pension coverage on mobility varies with tenure and how much pension coverage matters relative to other variables have received little attention. In this section we examine these questions using data from the CPS and the 1975-1982 Panel Survey of Income Dynamics.

Two measures of pension coverage are available in the CPS. One indicator is based upon whether the employer offers a retirement plan (to which the employer contributes) and does not consider whether workers actually participate in the plan. We will call this measure of coverage "pension provision." The other indicator is based upon whether the employee is actually included in the retirement plan. We will use the term "pension participation" to refer to this measure.

Reasons for nonparticipation at establishments where pensions are provided are quite varied. In some cases, not all jobs at an establishment are covered by the retirement plan. As long as job durations in such situations are considerably shorter than those in most other jobs, the pension provision measure understates the impact of pensions on job duration. In other cases, the worker is a new employee and, although not yet eligible to participate, would be covered by the retirement plan in a fairly short time should he stay. This creates a bias in the pension participation measure because it categorizes workers with little tenure, who will eventually participate in a pension on that job, as not being covered. In Table 1 we report data on job duration based on the provision measure.[2]

Length of service for workers with employers providing pensions is remarkably longer than that of other workers. In 1979 the mean duration of workers in jobs where pensions are provided was 8.3 years as compared to 4.0 years for workers in other jobs. The figures for 1983 are comparable; 8.8 years

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[2]The distributions based on the participation measure, available from the authors, show an even wider difference in job duration between workers participating in pensions and other workers.

for workers in jobs where pensions are provided and 4.1 years for workers in other jobs.[3]

Data on spells in progress could provide misleading estimates of the impact of pensions on completed spells and mobility rates if new hire rates in earlier years vary by pension coverage. Mobility can be estimated by combining the 1979 and 1983 CPS files to calculate the retention rate, the ratio of 1983 to 1979 employment. Employment estimates are based on the May supplement sampling weights. The somewhat unorthodox choice of years of service categories in Table 2 minimizes measurement error caused by the tendency of respondents to report years of service in multiples of five. No adjustments can be made in the CPS data for the creation or termination of pension plans, thereby introducing an additional degree of error into the analysis.

Between 1979 and 1983 those in jobs where pension plans are provided were much more likely to stay in their jobs than other workers, as shown in the first two rows of Table 2. The retention rate for workers with thirty years or less of service was 61 percent where pensions were provided in contrast to 41 percent for other workers. Part of this difference results from the fact that most new employment spells are found in establishments where no pensions are provided. If persons with zero years of service are excluded from the analysis, the retention rate for those in jobs where no pensions are provided increases to

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[3]These are employment spells that are still in progress; rough estimates of the duration of completed spells can be obtained by doubling mean tenure, an approximation discussed in the context of unemployment duration by Salant (1977).

50 percent, still substantially below the 66 percent rate for those working where pensions are provided.[4]

To determine how the impact of pension coverage on mobility varies with years of service, Table 2 reports retention rates by pension provision status for those with less than four years of service in 1979 and by pension participation status for those with four or more years of service. The rationale for this approach is that any bias in estimates of pension impact based on the participation measure should be present only for new workers, whereas the pension provision measure includes workers who would never be covered by their employer's plan.

The impact of pensions on mobility declines with years of service. The difference in mobility rates between workers covered by pensions and other workers is greatest in both absolute and relative terms for those with eight years or less of service in 1979. Yet even among workers with 9 to 23 years of service, the retention rate is still much greater for those participating in pensions than for those who are not. Retention rates for those with 24 to 28 years of service are about the same for pension participants and nonparticipants, as both groups begin to retire in significant numbers.

Pension coverage is correlated with a host of other variables that influence job duration, including age, sex, firm size, industry, occupation, and unionization. To determine how much pension coverage affects retention rates, holding these other variables constant, Table 3 reports some simple length-of-

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[4]When retention rates are broken down by years of service and pension participation, the retention rate for pension participants becomes twice as high as the retention rate for nonparticipants (71 versus 34 percent).

service and mobility results from our own analysis of the PSID and the May 1983 CPS Pension Supplement, along with the findings of six other studies.[5] The entries in Table 3 compare the impact of pension coverage with that of $1 an hour additional compensation, one additional year of service, and unionization.

In our analysis of the 1983 CPS, years of service among males is 35 percent longer for those whose employers provide pensions; among females, 30 percent longer. Union membership has a more modest impact on length of service -- 15 percent longer among men and 21 percent longer among women. An extra dollar of average hourly compensation is associated with an even more modest 6 to 8 percent increment in length of service. Establishment and firm size dummies are also included in the model. The results for men are:

Size category	Establishment size coefficient	Firm size coefficient
25-99	.102	-.023
100-499	.109	-.032
500-999	.189	-.016
1000+	.163	.127

A man employed in an establishment with 1000 or more workers has spent on average 34 percent more years with his employer than a person employed in a

[5]McCormick and Hughes (1984) and Ippolito (1985) are excluded from the table because they interact pension coverage with other variables and do not report sample means for all variables. Although Wolf and Levy (1984) estimate hazard functions for length of service separately for workers covered and not covered by pensions, we exclude their results from the table because age is the only other variable in their analysis.

firm with fewer than 25 workers, a difference comparable to the impact of pension coverage. In the female sample, the firm and establishment size dummies were not significantly correlated with length of service.

Three other studies summarized in Table 3 also show that pensions are strongly associated with length of service. Both Freeman (1980) and Leigh (1979) find that pension coverage is the strongest correlate of length of service; union membership has a sizable but more modest impact. In Rebitzer's (1986) study of the 1979 CPS, firm and establishment size are the strongest correlates of years of service, but the impact of pension coverage was still very sizable relative to the other variables in the analysis.

Pensions also have a very pronounced effect on mobility. In the PSID, the odds of leaving one's job between 1975 and 1982 are 15 percentage points lower for those covered by pensions than for other workers, an impact five or more times larger than that of other variables.[6] Mitchell's (1982) analysis of the 1973-77 Quality of Employment Survey (QES) also finds that among men pension coverage is the strongest correlate of the probability of a job change. In Mitchell's analysis of women and in Ippolito (1987), pension coverage and union membership have comparable impacts on the odds of changing jobs. Freeman's analysis of the National Longitudinal Survey (NLS) of mature men between 1969 and 1971 is the only study to examine job changes that did not find a significant pension impact. Freeman's evidence is consistent with the

[6]PSID respondents are defined as being covered by a pension if they answer affirmatively to the question "Are you covered by a company retirement plan?"

finding in Table 2 that the impact of pensions on mobility is much smaller among senior employees.

The correlation between pensions and quits is not as overwhelming as that between pensions and mobility. In the PSID, quit rates are 9 percentage points lower for workers covered by pensions. In four other studies, three out of six estimates fail to find any relationship between pensions and quits. Unionization is correlated with lower quit rates in every estimate except the PSID, suggesting that union membership is a better predictor of quit probabilities than pension coverage.

Further evidence that pensions have a larger effect on total separations than on quits comes from Department of Labor (1963), which compared annual separation rates in 1955 by pension coverage, broken down by age group and firm size. The difference in quit rates by pension coverage is much smaller than the difference in separation rates, as shown below:

	Separation rate	Quit rate
Under 45, pension	42	23
Under 45, no pension	80	33
45-64, pension	16	5
45-64, no pension	46	15

To summarize, we have shown that length of service and job retention rates in the 1979 and 1983 CPS are both considerably greater for persons covered by pensions than for other workers. We also have found that the impact of pension coverage on mobility varies with tenure. It is greatest at low levels of tenure, but it is sizable at all tenure levels. This suggests that different mechanisms may be at work at different career stages. Pension coverage remains strongly correlated with mobility when we attempt to control for covariates including age, firm or establishment size, unionization, occupation, industry, and sex. In most of the evidence presented here, pension coverage has been the strongest correlate of mobility and length of service, a point that previous studies have not emphasized. To understand lifetime jobs in the U.S., one must understand pensions.

III. HOW PENSIONS AFFECT MOBILITY

Pensions can reduce mobility by imposing a capital loss on those who leave. Most workers covered by pension plans have their benefits determined by a formula that multiplies final earnings by a constant and years of service. As long as workers expect their nominal earnings to rise over time, such formulas discourage mobility by penalizing workers who leave their jobs. For instance, consider a case in which a worker earns \$20,000 after twenty years and \$40,000 after forty years in the labor market and is covered by a pension plan that annually will pay him when he retires 1.5 percent of final earnings for each year of service. If he stays on the same job throughout this period, his annual benefit will be \$24,000 (.015 x 40 x \$40,000). However, after his twentieth year, if he moves to a new employer who has an identical pension plan and pays the same wage, his benefit will be only \$18,000 (.015 x 20 x \$20,000 + .015 x 20 x \$40,000).

Despite this characteristic of the benefit formula, the existence of a capital loss hinges upon how much workers pay in each year (via reduced earnings) for their pensions. Under the legal interpretation of the pension contract described in Bulow (1982), workers pay only for the benefits to which they are legally entitled. In the case above, a worker with twenty years of service pays for a pension based on a final salary of $20,000 rather than one based on an expected final salary of $40,000. Under the legal pension contract, this worker has nothing to lose if he leaves the firm.

Ippolito shows that there must be an implicit contract between the firm and the worker for a capital loss to exist.[7] Under this contract the worker expects to continue to be employed with the same firm until retirement and pays each year for a pension based on his expected final earnings ($40,000 in the example). The present value of his expected pension benefits is called the "stay pension." Assume for simplicity that worker i survives to retirement with certainty and that pension benefits are paid in a lump sum (B_i) at retirement date R and are based on the formula $B_i = AS_iY_i(t)$, where $Y_i(t)$ indicates earnings at each date, S_i indicates years of service, and A is a constant reflecting plan generosity. At time t ($t<R$), the stay pension is

$$(1)\quad SP_i = AS_iY_i(R)e^{-r(R-t)},$$

where $Y_i(R)$ is expected final earnings. If he leaves the firm, the worker receives only the benefits to which he is legally entitled, $AS_iY_i(t)$. This will be

[7]The concept of such a capital loss was introduced by McCormick and Hughes (1984) and was integrated into models of financial and labor market behavior by Ippolito (1985, 1987).

less than $AS_iY_i(R)$ as long as $Y_i(t) < Y_i(R)$. The present value of $AS_iY_i(t)$ is called the "leave pension,"

(2) $LP_i = AS_iY_i(t)e^{-r(R-t)}$.

The capital loss (CL_i) is the difference between the stay and the leave pension,

(3) $CL_i = AS_i[Y_i(R) - Y_i(t)]e^{-r(R-t)}$.

The traditional explanation given for lower mobility under pensions has been the impact of vesting rules. A person who leaves a job before he is vested is not entitled to pension benefits from that job upon retirement. The impact of vesting is expressed by (3) by assuming $Y_i(t) = 0$ for unvested workers.[8]

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[8]Before the Employee Retirement Income Security Act (ERISA) was passed in 1974, there were no restrictions on vesting provisions. In some cases a worker had to stay with the firm until he reached retirement age to be vested. Kolodrubetz and Landay (1973) found that in a 1972 sample restricted to workers aged 50 or above with 10 or more years of service, only half those covered by pensions were fully vested. In such a regulatory environment it is quite easy to see how this explanation became conventional. After ERISA was enacted, most firms adopted policies to vest workers fully after 10 years of service. Such vesting policies are likely to have little effect on turnover because the gain in pension wealth at the point of vesting is actually quite small for most workers, as shown by Kotlikoff and Wise (1985). Schiller and Weiss (1979) examine the impact of vesting on mobility in the pre-ERISA era; Wolf and Levy (1984) examine mobility before and after ERISA.

Estimates of this loss based on representative samples of actual pension plans are reported in Allen, Clark, and McDermed (1988). Using pension data from the 1983 Employee Benefit Survey of Medium and Large Firms and data on workers from the May 1983 CPS, Allen, Clark, and McDermed estimate that the pension loss roughly is between half and two-thirds of earnings for workers aged 40 to 55 in most occupations and industries.[9]

When one also takes into account the shorter time horizon over which these workers can recoup these losses through higher earnings at some alternative employment, the capital loss could conceivably be a powerful impediment to mobility and shirking within this age group. Furthermore, under Lazear's model, if a firm is concerned about its reputation in the labor market and its ability to write underpay-early/overpay-later contracts in the future, the capital loss should also be associated with fewer layoffs. This theory is consistent with the higher retention rates for pension participants among workers with nine or more years of service in Table 2.

This still leaves unaddressed the fact that most mobility is concentrated among young workers with relatively little service. Table 2 demonstrated that pension coverage drastically reduces mobility among those with fewer than four years of service. These workers face negligible capital losses if they leave their jobs. However, the prospect of suffering a sizable capital loss in the future operates as a self-selection device matching stable workers with jobs covered by pensions.

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[9]A different measure of the capital loss associated with leaving one's job that explicitly incorporates assumptions about the odds of staying on the job an extra year is developed in Lazear and Moore (1988).

This can be seen through the following simple example. Suppose there are two types of workers, with quit probabilities m and s, with m > s. These differences could arise from expected differences in the value of nonmarket time or in mobility costs. A firm that must invest a great deal in worker-specific training will want to attract the s-applicants. This can be done by setting up a compensation system that includes a pension with delayed vesting and an earnings-based formula.[10] This will discourage the m-applicants. It will also encourage the s-applicants if they end up receiving more compensation than they could elsewhere, as they would under certain efficiency wage models. The result is a set of employees with lower initial odds of leaving the firm. A similar argument is presented formally in Viscusi (1985) with several interesting extensions involving the precision of prior quit probability assessments and on-the-job learning. The self-selection of m and s-applicants generated by bonding can account for the relatively large retention rates among workers in Table 2 who have little tenure on jobs covered by pensions, but it cannot account for the impact of pensions on turnover among more senior workers (all of whom should be s-applicants).

Besides bonding, the other rationale for the greater length of service and lower mobility rates observed under pensions is that it merely reflects a higher overall level of compensation. Economic models of mobility imply that quit rates are inversely related to the difference between compensation at one's current and one's best alternative job. Assuming that alternative compensation is held constant through variables representing human capital and other

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[10]It can also be done with any other deferred compensation scheme.

individual characteristics, higher wage rates and pension coverage both should be associated with lower quit rates.

All of the studies summarized in Table 3 included wages or earnings as a control variable. It could be that if one could hold total current compensation constant, pension coverage would no longer be related to mobility. However, high levels of current compensation relative to alternative compensation may also be associated with greater layoff probabilities, making the direction of the overall linkage between pay levels and mobility ambiguous, a point raised in Antel (1985). Another problem with this rationale is that it cannot by itself explain why the impact of pensions on mobility in Table 2 falls with tenure. Pension compensation is either a constant or rising share of total compensation, so one would expect the impact of pensions on mobility either to stay the same or to increase with seniority if compensation levels were mainly responsible for the correlation between pension coverage and mobility.

IV. EMPIRICAL MODEL

The strategy used in this paper to model the sorting of workers into jobs by pension coverage is to estimate a three equation model of pension coverage, turnover in pension jobs, and turnover in jobs without pensions. Sorting is based on observable and unobservable characteristics of the worker and the job. Even when deferred compensation eliminates many turnover-prone applicants for a job, firms continue to have an incentive to select those with the lowest odds of turnover other things equal. Variables such as schooling and age could act as signals of employment stability that influence how workers are matched with employers providing job-specific training. This requires a model that will treat sorting based on observable characteristics as well those that are unobservable (at least in our data set).

Separate estimation of pension and turnover equations identifies sorting effects (e.g., married applicants are more likely to be hired into pension jobs where turnover odds are low for all workers) versus effects that are conditional on sorting (e.g., married workers have lower turnover odds). Separate mobility equations by pension status also distinguish whether turnover is lower in pension jobs because the intercept of the mobility function is lower or because the responsiveness of mobility to exogenous variables is lower. For instance one would expect an m-applicant to have higher turnover odds than an s-applicant and may even be more responsive to a dollar of additional hourly compensation than an s-applicant. These are testable hypotheses in our approach. Correlations between the mobility equations and the pension coverage equation are estimated to determine the role of unobservables in the sorting process.

The model also estimates how the possibility of suffering a capital loss of pension wealth influences the sorting process. Presumably persons who anticipate a large capital loss on a pension job will prefer to take a job without a pension. This hypothesis can be tested by including the expected cost of leaving a pension job as a right-hand variable in the pension equation.

The model consists of three equations

(4) $P_i^* = \beta_{11}' X_{1i} + \beta_{12} \hat{T}_{Pi} \cdot CL_i + \epsilon_{1i}$

(5) $T_{Pi}^* = \beta_{21}' X_{2i} + \beta_{22} CL_i + \epsilon_{2i}$

(6) $T_{Ni}^* = \beta_{31}' X_{3i} + \epsilon_{3i}$,

where P_i^* is a latent variable indicating the odds that a particular worker will be covered by a pension; T_{Pi}^* and T_{Ni}^* are latent variables indicating the odds of leaving a job with a pension (P) and without a pension (N); X_{1i}, X_{2i}, and X_{3i} are vectors of control variables; $\hat{T}_{Pi} = \text{Prob}(T_{Pi}^* > 0)$; and $\epsilon_{1i} \sim N(0,\sigma_1^2)$, $\epsilon_{2i} \sim N(0,\sigma_2^2)$, $\epsilon_{3i} \sim N(0,\sigma_3^2)$, $E(\epsilon_{1i}\epsilon_{2i}) = \sigma_{12}$, $E(\epsilon_{1i}\epsilon_{3i}) = \sigma_{13}$, and $E(\epsilon_{2i}\epsilon_{3i}) = 0$.

Equation (4) indicates how expected mobility costs influence the assignment of workers to jobs by pension coverage. A more structural model that would take into account the linkage between pension choice and the job matching process is beyond the scope of this paper because pension coverage is known at only one point in time in our data.

Workers are assumed to take jobs with pensions if the value of the pension-covered job (including compensation and nonmonetary considerations) exceeds that of a job without a pension. They are more likely to be covered by pensions when they face high marginal tax rates. Pensions delay taxes until retirement, a time when most people expect to be in lower marginal tax brackets. The value of tax postponement typically is greater for those currently in high tax brackets. Total compensation and the marginal tax rate are included in X_{1i} for this reason. Firms have an incentive to provide pensions when monitoring costs are high (to reduce shirking) or when they have high hiring and training costs (to reduce turnover). To control for these factors, years of schooling and industry and occupation dummies are included in X_{1i}. Unionized establishments are also more likely to provide pensions, as discussed in Freeman (1985). Sorting into pension jobs is likely to be based on observable characteristics such as age, schooling, race, marital status, and number of children, all of which are included in X_{1i}. Tenure is also included to account for the fact that pension participation at some employers depends on years of service.

When workers choose jobs with pensions, they also choose a set of specific pension plan characteristics, including the benefit formula determining CL_i. Key variables in this matching process include worker attitudes toward risk bearing, the amount of information they have about the plan, and their expected odds of staying with the plan. Workers who do not expect to stay at a pension job should be less likely to be observed in such jobs because holding other things equal, including total compensation before taxes (regular earnings and pension compensation combined), they would expect a capital loss.

In a world of perfect information, workers would base their pension choice decision on the expected cost of leaving a pension job, the product of $\tilde{T}_{Pi}$ and CL_i. In practice it is much more likely that workers know $\tilde{T}_{Pi}$ than the magnitude of the function generating CL_i before (or even after) accepting a job.[11] Accordingly we also consider a variant of (4) where pension coverage is related solely to $\tilde{T}_{Pi}$:

$$(4') \quad P_i^* = \beta_{11}' X_{1i} + \beta_{12} \tilde{T}_{Pi} + \epsilon_{1i}$$

and estimate (4') jointly with (5) and (6). In this formulation the capital loss affects pension coverage only indirectly through its effect on turnover.

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[11]Many workers do not know the most basic characteristics of a pension such as whether early retirement is provided or whether employees contribute to the plan, as shown in Mitchell (1988).

Observable factors in the process matching workers with employers by pension coverage are modelled in (4), whereas the role of unobservable factors is estimated by σ_{12} and σ_{13}. The signs of σ_{12} and σ_{13} cannot be predicted ex ante. Generally one would expect more stable workers to be in pension jobs, making $\sigma_{12} < 0$, but if observable factors play the predominant role in the matching process, the value of σ_{12} could be zero. The sign of σ_{13} could be negative if workers with high odds of pension coverage are more stable regardless of whether they are covered by a pension. Alternatively, it could be positive if there is queuing for pension jobs. In this latter case, those with high odds of pension coverage who land in a job without a pension are more likely to leave because they prefer a job with a pension.

The impact of CL_i and compensation levels on turnover in pension jobs is estimated in (5). The validity of the bonding hypothesis is indicated by β_{22}. The hourly compensation variable in X_{2i} and X_{3i} is calculated on an after-tax basis. To determine whether persons with pensions stay on their jobs longer merely because they are paid better, we include pension compensation (PC_i) in the compensation variable. Pension compensation is the change in pension wealth obtained by staying an extra year with an employer (dLP_i/dS_i). We assume that a person on the margin of leaving a job will define PC_i according to the "leave pension" definition in (2).

The impact of vesting rules on mobility is an interesting policy question. This issue is examined by including two binary variables in X_{2i}: one indicating whether the respondent was vested at the beginning of the sample period and another indicating whether the respondent became vested during the sample period. We assume that all workers are 100 percent vested at 10 years of service but are completely unvested beforehand. This assumption was the predominant practice under ERISA in the sample period examined below. Worker and job characteristics, along with regional dummy variables, are also included in X_{2i} and X_{3i}.

An alternative model for estimating pension effects on mobility has been developed by Gustman and Steinmeier (1987, 1990). They estimate the effect of lifetime compensation premiums on mobility using the 1983 Survey of Consumer Finances (SCF) and the 1984-1985 Survey of Income and Program Participation. The premium estimate is the difference in lifetime wage and pension income between the current and alternative job. It contains the pension bond and the efficiency wage premium. Their approach differs from ours in two important respects. First, Gustman and Steinmeier do not model the sorting of workers into jobs by pension coverage. Second, Gustman and Steinmeier model alternative compensation explicitly, equating it to income on the new job, a variable observable only for job changers, who are not likely to be a random subsample, and not known even by them except on an ex post basis. Because of the inherent difficulties in estimating lifetime alternative compensation ex ante, we restrict our attention to the linkage between current pay levels and mobility.

Parameter estimates for (4), (5), and (6) can be estimated using standard maximum likelihood methods. The parameters of (5) and (6) are identified through sample separation and (4) is identified through exclusion restrictions. In the data, only the dichotomous counterparts of P_i^*, T_{Pi}^*, and T_{Ni}^* are observed,

$$P_i = 1 \text{ if } P_i^* > 0, 0 \text{ otherwise.}$$

If covered,

$$T_{Pi} = 1 \text{ if } T^*_{Pi} > 0, \text{ 0 otherwise,}$$

and if not covered,

$$T_{Ni} = 1 \text{ if } T^*_{Ni} > 0, \text{ 0 otherwise.}$$

Therefore the variances of the disturbance terms are normalized to one. The resulting joint probability distribution of (P, T_P, T_N) is given as

$$P_{11} = \text{Prob}(P = 1, T_P = 1) = F[\beta'_{11}X_{1i} + \beta_{12}\tilde{T}_{Pi} \cdot CL_i, \beta'_{21}X_{2i} + \beta_{22}CL_i; \rho_{12}]$$

$$P_{10} = \text{Prob}(P = 1, T_P = 0) = F[\beta'_{11}X_{1i} + \beta_{12}\tilde{T}_{Pi} \cdot CL_i, -(\beta'_{21}X_{2i} + \beta_{22}CL_i); -\rho_{12}]$$

$$P_{01} = \text{Prob}(P = 0, T_N = 1) = F[-(\beta'_{11}X_{1i} + \beta_{12}\tilde{T}_{Pi} \cdot CL_i), \beta'_{31}X_{3i}; -\rho_{13}]$$

$$P_{00} = \text{Prob}(P = 0, T_N = 0) = F[-(\beta'_{11}X_{1i} + \beta_{12}\tilde{T}_{Pi} \cdot CL_i), -\beta'_{31}X_{3i}; \rho_{13}],$$

where F(.) denotes the bivariate cumulative normal distribution, ρ_{12} denotes the correlation coefficient between P and T_P, and ρ_{13} denotes the correlation coefficient between P and T_N. The likelihood function to be maximized is

$$L(\beta_{1.},\beta_{2.},\beta_{3.}) = \prod_i P_{11}^{P_i T_{Pi}} \cdot P_{10}^{P_i(1-T_{Pi})} \cdot P_{01}^{(1-P_i)T_{Ni}} \cdot P_{00}^{(1-P_i)(1-T_{Ni})} \quad .$$

The results reported below are estimated with the Davidon, Fletcher, and Powell method.[12] Estimates of the asymptotic covariance matrix are computed using the Berndt et al. estimator.

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[12]Experimentation with other algorithms (Newton's method), various convergence criteria, and alternative starting values showed them to be quite stable with respect to the optimization procedure.

V. DATA DESCRIPTION

In most longitudinal micro data sets there are annual observations for each variable, allowing the researcher to use fixed effects or pooled time-series and cross-section estimation procedures. This cannot be done here because no longitudinal data set reports pension coverage on a regular basis. The PSID reports pension coverage in 1975, but not again until 1984. Because we are concerned with job duration, we focus on the PSID data between 1975 and 1982, the last year of data available at the time this study was started. The alternative was to consider a shorter interval. We chose not to do this because the turnover rate for persons covered by pensions is extremely low.

Unfortunately, the PSID provides little information on pension plan characteristics. The PSID reports in the 1975 survey whether the respondent is "covered by a company retirement plan." The pension benefit formula and eligibility criteria for benefits are needed to calculate CL_i.

These variables were imputed from the 1983 Employee Benefit Survey for Medium and Large Firms (EBS) in the following way. Plans in the EBS were sorted into eight industry and three occupational classifications.[13] Within each industry-occupation cell there are as many as five different types of pension formulas (e.g., simple earnings-based or dollar per year of service). The formula type that covered the largest proportion of participants within each cell

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[13]In the construction industry we did not make any breakdowns by occupation because of the extremely small number of observations in the EBS.

was assumed to apply to all participants in that cell. The mean parameter values for that formula type are used as the estimate of the benefit formula for all PSID respondents in a given industry-occupation category. For earnings-based formulas, the key parameters are the generosity factor (percentage of average earnings) and the length of the salary averaging period. Age and service requirements for normal retirement in the EBS were assumed to be equal to cell means, based on all plans in the cell regardless of formula type.

Given the benefit formula, it is straightforward to calculate PC_i and CL_i with information on age, earnings history and years of service given by PSID respondents. The value of a life annuity beginning at the normal retirement age was calculated and discounted to age at the time of the survey. The discount factor used was 9 percent; this value corresponds to long-term top-grade bond rates during this period. Survival probabilities by sex and race were set equal to those for the U.S. population for 1981.[14] The sample is restricted to whites and blacks because survival probabilities for other racial groups are not reported. We assume no postretirement adjustments in benefits are provided. To calculate $Y_i(R)$ we assume that earnings are projected forward to the age of eligibility at the interest rate, following Ippolito (1985).[15]

The PSID sample consists of private wage and salary workers who were employed at the time of the 1975 survey and reported earnings. The sample is restricted to heads of households under age 55 in 1975 who were working 35 or

[14]U.S. Bureau of the Census (1984), Table 104, p. 70.

[15]In dollar-per-year-of-service formulas, we assume the dollar amount in the formula is projected forward at the interest rate.

more hours per week.[16] Those working in occupations or industries not represented in the EBS files were deleted.[17] Observations with missing values of the variables were also dropped from the sample. Out of the 6,740 households in the PSID, 1,111 observations met our criteria.[18] The focus of the empirical analysis of the PSID below will be on whether the respondent left his 1975 employer by 1982.[19] We also will examine terminations resulting from layoffs and decisions to quit to find another job.[20] All independent variables are based on information from the 1975 survey except for the pension variables, which rely on earnings data from as far back as the 1970 survey, and

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[16]A small percentage of respondents claim to be retiring at the time they leave their 1975 job, but in subsequent surveys many of them report that they are working full time. The sample restriction on age removes almost all of the respondents who seem to have permanently retired from the sample.

[17]Persons employed in the following PSID industrial categories are excluded: agriculture, forestry, fishing, and trade, NA whether wholesale or retail. Occupations excluded include farmers, farm managers, and miscellaneous occupations.

[18]Splitoff families are excluded from the sample to facilitate data base management. Persons who were self-employed in any year of the survey are excluded to focus the analysis on wage and salary workers.

[19]Turnover is defined according to the reason given by the respondent why he left his previous job. If that job ended because of a plant closing, strike, layoff or dismissal, quit, or other reason or because it was a temporary job that had been completed, the dependent variable is set equal to one.

[20]Layoffs are defined as terminations resulting from plant closings, layoffs, or dismissals. If a person quit his previous job and was either employed or unemployed at the time he was surveyed, he is labelled below as someone who quit to find another job. If a person quit and was out of the labor force at the time of the survey, he is labelled below as someone who quit to exit the labor force.

the federal marginal tax rate, which comes from the 1976 survey but is based on 1975 income. Average hourly compensation equals the sum of average hourly earnings, as reported directly on the survey, and average hourly pension compensation, which is obtained by dividing PC_i by the product of average weekly hours and weeks worked.

The other variables in X_i include union status, age, years of service and its square, number of children, years of schooling, sex, race, marital status, occupation, industry, and location. They are intended to control for observable forms of heterogeneity.

To summarize the significant patterns in the data, Table 4 reports mean values of selected variables by age group and mobility. The share of workers who left their 1975 job ranges from 66 percent for the under 25 group to 35 percent for the 45-54 group. Within each age group, those who stay with their 1975 employers (stayers) are much more likely to be covered by a pension than those who leave their 1975 employers (movers). Movers tend to have lower average hourly compensation and fewer years of service than stayers; movers are also more likely to be females or blacks. The rate of union membership is much greater for stayers than for movers among workers under 35, but is about the same for stayers and movers among older workers.

The mean value of the capital loss variable is quite low for the younger age groups, especially relative to the number of years they are likely to remain in the labor force. For instance, among workers covered by pensions, the mean capital loss is $821 for those under age 25 and $2,926 for those between 25 and 34. Assuming persons in the latter group work another thirty years, this amounts to $260 per year of remaining worklife at an 8 percent discount rate.

The capital loss averages $6,526 for the 35-44 group; if they work twenty additional years, this amounts to $664 per year of remaining worklife. For the 45-54 group the mean capital loss is $8,503 for those covered by pensions, meaning that workers in this age group would have to earn $1,267 extra per year on any alternative job to offset the capital loss over a ten-year period. The mean capital loss of stayers is much greater than that of movers in all four age groups; the difference is largest in absolute dollar terms in the 45-54 group.

Sample means by pension coverage are reported in Table 5. Workers covered by pensions are older, have more years of service, have more schooling, and are more likely to be union members that other workers. In terms of demographic characteristics, workers covered by pensions are more likely to be white, male, married, and to have more children. Compensation is much larger for workers covered by pensions.

We examined a number of surveys besides the PSID and CPS for assessing the effect of pensions on job mobility. Because of its unique and detailed pension information, the 1983 SCF appeared to be a likely candidate for inclusion in our study. However, this is a cross-sectional data set with limited information on employment history. The lack of data on earnings and pension characteristics on previous jobs is a serious shortcoming in employing the 1983 SCF for estimating turnover equations. In addition, Gustman and Steinmeier (1987) had to impute pension benefit formulas and eligibility criteria for 41 percent of the workers covered by pensions in their analysis of 1983 SCF because pension plan descriptions were not provided.

Recently, a 1986 follow-up to the SCF became available. The possibility of a panel survey that contained detailed pension information on the 1983 job suggests that the two SCF surveys would be an excellent data source for estimating job change equations. Unfortunately, the sample design of the 1986 survey did not include a reinterview of all 1983 respondents and there was a relatively high attrition rate (32 percent). Furthermore, about one-third of the attrition consists of households that were "not located" for the 1986 survey. Since change of residence is likely to be highly correlated with job change, the sampling has probably systematically eliminated many job changers from the 1986 sample. While we are still exploring the usefulness of both the 1983 SCF alone and the 1983-1986 surveys jointly, it is our current assessment that the gain in pension information on these surveys is outweighed by the cost of missing data on previous jobs in the 1983 survey and missing respondents in the 1983-1986 panel.

VI. RESULTS

Before reporting the results for the complete system of equations, it is instructive to examine the results for some simpler models. First consider a simple turnover equation with X_{3i} and a pension coverage dummy as right-hand side variables. Next replace the pension coverage dummy with CL_i and then in a third equation include both CL_i and the pension dummy. When these models are estimated over the sample used for the complete system, one obtains the following probit coefficients (standard errors are in parentheses and partial derivatives evaluated at the sample means are in brackets):

Pension coverage	Capital loss	Hourly compensation	Log likelihood
-0.322	--	0.061	-682.2
(0.098)		(0.030)	
[-0.128]		[0.024]	
--	-0.035	-0.049	-683.1
	(0.013)	(0.031)	
	[-0.014]	[-0.020]	
-0.260	-0.024	-0.042	-679.9
(0.104)	(0.013)	(0.031)	
[-0.103]	[-0.010]	[-0.016]	

In this model a \$1000 increase in CL_i is associated with a 1.0 to 1.4 percentage point reduction in turnover. The addition of CL_i to the equation reduces the pension coefficient by about 20 percent. The coefficient of the hourly compensation variable is quite sensitive to the inclusion of CL_i in the equation, but all other coefficients are fairly stable. The results for CL_i are quite similar in a bivariate probit model of pensions and turnover where β_{12} in (4) is restricted to equal zero and the parameters of turnover equations (5) and (6) are restricted to be the same. The estimated correlation between the pension coverage and turnover error terms is -0.140 with a standard error of 0.063, indicating that, even after controlling for observable characteristics, workers who are most likely to be observed in pension jobs are still less likely to leave those jobs.

The main results for the complete systems reported in Table 6 are as follows:

1. *A key reason why lower turnover is observed among workers covered by pensions is the prospect of capital losses of pension wealth.* The impact of a \$1000 increase in the capital loss is a 1.8 percentage point reduction in the odds of turnover. To put this result in perspective, note that the average worker covered by a pension faced a \$5024 capital loss in 1975 if he left his job. This translates into a turnover probability differential of 9 percentage points. The seven-year turnover rate was 39 percent for persons covered by pensions and 61 percent for persons not covered. Thus the capital loss accounts for about 41 percent of the turnover difference between these two groups.

The capital loss effect in the complete system is twice as large as in the simple single equation estimates noted above. This results from allowing the parameters of (5) and (6) to be different from each other to reflect sorting. To see this, compare the results in Table 5 to those obtained from a bivariate probit model consisting of a pension choice equation (4) (with $\beta_{12}=0$) and (5) estimated over all workers with ρ_{12} as a free parameter. In this model the CL_i coefficient is -.023, comparable to the simple probit results above, but a likelihood ratio test rejects the restrictions that $\beta_{21}=\beta_{31}$ and $\beta_{12}=0$. The simpler restriction that $\beta_{12}=0$ cannot be rejected.

2. *There is self-selection of workers with low odds of turnover into jobs covered by pensions but it is based on observable rather than unobservable characteristics.* Turnover in our sample is lower for men, whites, union members, and those who are married than for women, blacks, nonunion workers, and those who are not married. In addition turnover falls rapidly with

years of service.[21] Sex, race, union membership, and marital status are associated with pension coverage in the estimates of (4), whereas none of these variables are associated with turnover conditional on pension coverage in the estimates of (5) and (6).

Another indication of self-selection is that the response to changes in right-hand side variables varies considerably by pension coverage. The hypothesis that the coefficients of the two equations are equal is rejected by a likelihood ratio test. Turnover in jobs covered by pensions is mainly related to age, industry, and CL_i; turnover in other jobs is strongly related to years of service, region, and occupation. The intercept of the turnover equation is also much greater for workers without pensions.

The difference in the parameters between (5) and (6) is much more important than the difference in the means of the control variables between workers with and without pensions for explaining the observed difference in turnover rates. Predicted turnover for workers covered by pensions using the parameters of (5) is 34 percent, whereas predicted turnover for those without pensions based on the same parameters is 40 percent. Workers with pensions would have had a turnover rate of 56 percent, based on the parameters of (6), not much smaller than the 65 percent turnover rate predicted for those without pensions.

[21]These results are based on a simple probit equation estimated across our entire sample without any pension variables. Details are available upon request.

There is no correlation between the unobservables in the pension coverage and either of the turnover equations. The restriction that $\rho_{12}=\rho_{13}=0$ cannot be rejected by a likelihood ratio test at a 5 percent confidence level.

3. *The expected magnitude of the capital loss has little effect on the sorting of workers by pension coverage.* Despite this evidence on self-selection, the data do not indicate any direct link between the size of the expected capital loss and the odds of being in a job with a pension. Neither $\tilde{T}_{Pi} \cdot CL_i$ nor $\tilde{T}_{Pi}$ are significantly related to pension coverage, and contrary to expectations the coefficient is positive. The mean partial derivative of P_i with respect to CL_i in (4) is also positive. A more complete model that takes into account income profiles along with the transaction cost of switching out of a pension job is probably needed to make more progress on this issue. Also, even though years of service, compensation, industry and occupation are all "held constant" in (4), large values of CL_i reflect pension generosity as well.

4. *Pension compensation has little effect on turnover, but the higher overall level of compensation associated with pension coverage leads to lower turnover.* Pension compensation is a very small component of total compensation for most workers covered by pensions. The mean value of hourly pension compensation for such workers is 7.3 cents in 1975; for all but a few workers, it is well under a dollar. This fact alone makes it unlikely that the extra compensation associated with pension jobs explains much of the observed turnover difference.

Even though pension compensation is a relatively small component of total compensation, the gap in total compensation between jobs with and without pensions is quite large. The effect of a change in compensation on the odds of turnover must be evaluated using the entire system (4)-(6). Based on the coefficients from System A in Table 6, a one dollar increase in before-tax compensation is associated with a 1.8 percentage point decrease in the odds of turnover (almost exactly the same as our estimate in panel B of Table 3). If the \$2.05 difference in compensation between jobs with and without pensions were eliminated, the gap in turnover rates by pension coverage would narrow by 3.7 percentage points, accounting for 17 percent of the turnover difference between the two groups.

5. *There is no change in turnover at the point of vesting.* The restriction that the two vesting dummy variables in (5) equals zero cannot be rejected at conventional significance levels. This is not surprising in that CL_i takes into account changes in LP_i at vesting. The partial effect of tenure on mobility in (5) does switch from negative to positive at 10 years of service, but the magnitude of this derivative is quite small in this range (e.g., -.014 at 10 years and .008 at 15 years).

Beyond these key issues addressed by the results in Table 5, two other important questions can be addressed with our model and data. First, what types of turnover does the capital loss affect? Because we are unable to reject the hypotheses that $\beta_{12}=0$ and $\rho_{12}=\rho_{13}=0$, this question can be answered by simply estimating (5) for the two different types of turnover: quitting to take another job and layoffs. The results, reported in Table 7, show that the impact of CL_i clearly is greatest on layoffs. Layoff probabilities fall by 1.3 percentage points with a \$1,000 increase in CL_i. This finding is consistent with models of pensions being part of an implicit contract where bonding prevents shirking and reputational concerns prevent employers from pocketing CL_i by firing their workers. The much smaller effect of CL_i on quits is puzzling, although previous studies summarized in Table 3 tend to find pension coverage to be more strongly correlated with mobility in general than quits in particular.

Second, what is the association between mobility and wages and how does it vary by pension coverage? Between 1975 and 1982, average hourly earnings grew by 103 percent for workers who stayed at a 1975 job that was covered by a pension, whereas earnings growth was slightly lower (95 percent) for those who left such a job. Among those not covered by a pension in 1975, movers enjoyed wage growth of 113 percent versus 98 percent for stayers. In contrast to Gustman and Steinmeier (1987, 1990), the breakdowns of earnings growth by mobility and pension coverage in our sample shows no evidence that workers covered by pensions are collecting rents.

The estimates of CL_i are based on average pension characteristics in a given industry by occupation category, not the actual pension for the worker. The variation of CL_i across the persons in the sample reflects their seniority, age, and earnings history as well as estimates of the benefit formula and eligibility criteria that are measured with an unknown amount of error. If errors in measuring these provisions of the pension are sufficiently large, then CL_i is little more than a nonlinear combination of past earnings, seniority, and age. A skeptic could argue that there is really no new information in CL_i that is not already contained elsewhere in the model.

This criticism can be addressed in two ways. First, if there is no additional information in CL_i, then the hypothesis that its coefficient is zero should not be rejected. One would also expect the estimated effects of compensation and seniority to be considerably larger without CL_i in the model. Second, if CL_i contains no information about pension bonding and is merely picking up the effects of compensation and seniority on turnover, then it should have the same effect in (6) as it does in (5) -- CL_i should be a good predictor of turnover in jobs without pensions and the coefficients of compensation and seniority in (6) should diminish. This can be determined by adding CL_i to (6) and testing whether its coefficient is zero.

The data strongly reject both of these conjectures. Exclusion of CL_i from (4) and (5) is rejected by a likelihood ratio test at the five percent confidence level. In the restricted specification the compensation coefficient in (5) remains negligible in magnitude and is estimated with little precision. The magnitude and precision of the tenure variables do increase, however. When CL_i is added to (6), its coefficient (S.E.) is -.008 (.053) and the compensation and seniority parameters are unaffected. The conclusion to draw from these exercises is that CL_i contains useful information for predicting turnover that is not contained in other variables in the model and that information is relevant only for jobs covered by pensions.

VII. CONCLUSION

The mechanisms through which pension coverage reduces labor mobility are the large capital losses in pension wealth associated with leaving a job and the matching of more stable workers with firms that provide pensions and higher overall levels of compensation in those jobs. These conclusions follow from econometric evidence from the PSID. They are also consistent with the evidence in Table 2 on how pensions affect mobility at different tenure levels.

Do the capital loss results imply that Ross was premature in dismissing the new industrial feudalism hypothesis? Are many workers trapped in jobs they do not like because they fear they will lose tens of thousands of dollars in pension wealth? The answers to these questions hinge on whether capital losses reduce mobility mainly through quits or layoffs. Our results indicate that the capital loss has a much larger effect on layoffs than quits. In addition, previous studies summarized in Table 3 tend to find a much stronger correlation between coverage and mobility in general than between coverage and quits. If the quit-layoff distinction is a meaningful one, these results imply that pension plan parameters such as benefit formulas and vesting criteria have had a relatively small direct impact on quit decisions. The lower quit rate for workers covered by pensions seems to come mainly from self-selection. This suggests that public policy measures that have been proposed to increase labor market flexibility by changing the terms of pension contracts to make pension wealth more portable across employers will have little impact on mobility.[22]

These findings on pension capital losses are the first evidence that bonding directly affects employee behavior and that it is an important cause of lifetime jobs in the U.S. Extreme conclusions on the roles of bonding and efficiency wages in explaining job duration cannot be drawn from this result. The result certainly does not imply that firms are unconstrained in setting bonds. It does

[22]For a discussion of such proposals, see Choate and Linger (1986) and Clark and McDermed (1988).

imply that bonding should be taken seriously as an important component of the implicit pension contract.

These results are consistent with models of implicit contracts for lifetime jobs. Although employers seemingly would gain tremendous windfalls by terminating workers who have paid for stay pensions but legally stand to collect leave pensions, this opportunistic behavior apparently is being held in check by some force such as the cost of finding and training replacements or concern over labor market reputation. Reputational considerations also seem to explain why many employers provide large postretirement increases in benefits even when they are not contractually obliged to do so.[23] However, every firm does not keep its implicit pension promises; the question of when firms decide to renege on implicit contracts requires further study.

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[23]Postretirement increases in pension benefits are examined in more detail in Allen, Clark, and Sumner (1986).

Table 1. Percentage distribution of years of service by pension provision status, private wage and salary workers, May 1979 and 1983 CPS

	1979			1983		
Years of service	All	Pension not provided	Pension provided	All	Pension not provided	Pension provided
0	26.4	34.3	16.9	21.6	30.7	10.4
1	9.7	11.8	8.0	9.6	12.4	6.8
2	11.0	13.3	9.8	12.0	14.0	10.1
3-4	12.3	13.6	12.0	15.4	15.8	15.6
5-9	17.0	14.7	19.9	17.7	14.4	21.6
0-14	9.7	5.9	12.9	9.6	6.0	13.4
5-19	5.0	2.5	7.1	5.9	3.0	9.1
20-29	6.3	2.6	9.2	5.5	2.3	8.9
30+	2.6	1.4	4.2	2.7	1.4	4.1
Mean	6.2	4.0	8.3	6.3	4.1	8.8
Employment (millions)	67.9	24.7	38.1	71.3	30.0	35.8

Source: Calculated from May 1979 and 1983 CPS public use tapes, using sampling weights for pension supplement.

Note: Totals for all workers in each year includes workers who did not report pension coverage. Over half of these workers have zero or one year of service with their employers.

Table 2. Four-year retention rates by pension coverage, May 1979 and 1983 CPS

Years of service		Four-year retention rate	
1979	1983	Covered by pension	Not covered by pension
0-30	4-34	.608	.406
1-30	5-34	.656	.497
0-3	4-7	.490	.313
4-8	8-12	.715	.473
9-13	13-17	.748	.641
14-18	18-22	.756	.714
19-23	23-27	.804	.688
24-28	28-32	.581	.601

Source: Same as Table 1.

Note: Retention rate estimates for those with four or more years of service in 1979 are broken down by pension participation status; all other estimates are broken down by pension provision status.

Table 3. Estimates of the impact of pension coverage and other variables on length of service and mobility

Source (if published elsewhere)	Data Set	Dependent variable[a]	Impact on dependent variable of Pension coverage	$1 in average hourly compensation[b]	One additional year of service[c]	Union membership
			A. Length of Service Estimates			
	CPS,1983					
	Males	Ln(years of service)	.30	.06	n.a.	.14
	Females	Ln(years of service)	.26	.08	n.a.	.19
Rebitzer (1986)	CPS,1979					
	Males	Years of service	.08	.37	n.a.	.39
	Females	Years of service	.90	.49	n.a.	1.57
Freeman (1980)	NLS mature men 1969-71	Years of service in 1969	4.65	.43	n.a.	2.96
Leigh (1979)	NLS mature men, 1969-71					
	Whites	Years of service in 1969	.61	.44	n.a.	2.31
	Blacks	Years of service in 1969	3.71	.43	n.a.	3.58
			B. Mobility estimates			
	PSID, 1975-82	Probability of job change	-.15	-.02	-.03	n.s.
Ippolito (1987)	CPS,1979-80	Probability of job change	.06	-.001	-.03	-.06

Table 3. (continued)

Mitchell (1982)	QES,1973-77					
	Males	Job change probit	.63	-.10	-.04	n.s
	Females	Job change probit	-.40	n.s.	n.s.	-.46
Freeman (1980)	NLS mature men,1969-71	Total separations logit	n.s.	n.s.	-.16	-.63
			C. Quit Estimates			
Mitchell (1982)	QES,1973-77					
	Males	Quit probit	-.44	n.s.	-.11	-.39
	Females	Quit probit	n.s.	n.s.	n.s.	-.62
Freeman (1980)	NLS mature men,1969-71	Quit logit	n.s.	n.s.	-.24	-1.85
Viscusi (1979)	NLS mature men,1969-71	Quit logit	-.83	n.s.	-.17	-1.62
Leigh (1979)	NLS mature men,1969-71					
	Whites	Quit probit	-1.53	-.10	--	-.87
	Blacks	Quit probit	n.s.	n.s.	--	-1.01

[a]When authors estimated probit equations and reported derivatives of the equation at the sample means, the derivatives are reported in the table. Otherwise, the actual probit (or logit) equation coefficients are reported and the dependent variable is referred to as probit (or logit).

[b]In studies where this variable is expressed in logs, we assume that this is equivalent to a .18 change in logs (the logarithmic difference between $5 and $6).

[c]In studies where this variable is entered quadratically, we estimate the derivative in the neighborhood of five years of service and 35 years of age.

Note: n.s. indicates the coefficient was not statistically significant at a 10 percent confidence level using a two-tailed test; n.a. indicates the variable is not applicable; -- indicates the variable was not included in the study.

Table 4. Means of selected variables, 1975-1982 PSID, by age group and mobility status

	Less than 25			25 to 34			35 to 44			45 to 54		
	All	Stayers	Movers	All	Stayers	Movers	All	Stayers	Movers	All	Stayers	Movers
Covered by pension	.60	.77	.52	.69	.77	.61	.73	.81	.60	.74	.78	.65
Capital loss in $1000 (if covered)	.82	1.19	.54	2.93	3.32	2.42	6.53	7.20	5.10	8.50	9.83	5.51
Average hourly compensation	3.92	4.66	3.54	5.16	5.45	4.88	5.68	5.96	5.24	6.19	6.57	5.47
Union member	.31	.44	.25	.32	.39	.25	.39	.38	.41	.36	.36	.38
Years of service	1.80	2.19	1.61	3.77	4.44	3.10	8.01	8.88	6.66	11.43	12.42	9.56
Male	.88	.89	.88	.89	.91	.87	.86	.87	.85	.86	.87	.85
Black	.34	.28	.37	.30	.29	.31	.31	.30	.34	.26	.22	.32
N	169 (100%)	57 (34%)	112 (66%)	404 (100%)	203 (50%)	201 (50%)	250 (100%)	152 (61%)	98 (39%)	288 (100%)	188 (65%)	100 (35%)

Table 5. Means and standard deviations by pension coverage

	Not covered	Covered
Age	34.02 (10.30)	36.21 (9.94)
Tenure	4.40 (5.06)	7.29 (7.26)
Union	.17 (.37)	.43 (.49)
Number of children	1.44 (1.49)	1.60 (1.56)
Black	.35 (.48)	.27 (.45)
Years of schooling	11.57 (2.86)	12.21 (2.80)
Male	.79 (.41)	.91 (.28)
Not married	.31 (.46)	.17 (.38)
Federal marginal tax rate	.18 (.08)	.23 (.08)
Hourly compensation before tax	3.99 (2.06)	6.04 (2.95)
Average hourly earnings	3.99 (2.06)	5.95 (2.88)
Hourly compensation after tax	3.16 (1.42)	4.55 (1.89)
Capital loss (in $1000)	1.94 (3.30)	5.02 (6.15)
Already vested		.26 (.44)
Became vested		.27 (.44)
Changed job	.61 (.49)	.39 (.49)
Quit	.31 (.46)	.19 (.39)
Layoff	.26 (.44)	.15 (.36)
N	337	774

Note: Standard deviations appear in parentheses.

Table 6. Estimates of three equation model of pension coverage and mobility.

Equation	System A[a] (4)	System A[a] (5)	System A[a] (6)	System B[a] (4')	System B[a] (5)	System B[a] (6)
Age	-.006	-.013	-.006	-.008	-.012	-.005
	(.006)	(.006)	(.011)	(.007)	(.006)	(.012)
	[-.002]	[-.004]	[-.002]	[-.002]	[-.004]	[-.001]
Tenure[b]	.016	-.043	-.181	.029	-.034	-.188
	(.034)	(.056)	(.064)	(.035)	(.056)	(.053)
Tenure squared/100	-.002	.207	.685	-.012	.191	.696
	(.121)	(.148)	(.224)	(.130)	(.148)	(.210)
	[.004]	[-.005]	[-.035]	[.007]	[-.003]	[-.040]
Union	.599	-.057	.085	.590	-.110	.001
	(.119)	(.154)	(.424)	(.120)	(.174)	(.515)
	[.144]	[-.019]	[.026]	[.155]	[-.037]	[.0003]
Number of children	.040	.012	-.038	.045	.008	-.040
	(.038)	(.035)	(.069)	(.039)	(.036)	(.068)
	[.011]	[.004]	[-.012]	[.012]	[.003]	[-.013]
Black	.247	.017	-.122	.229	.042	-.155
	(.121)	(.127)	(.281)	(.121)	(.132)	(.285)
	[.059]	[.006]	[-.038]	[.059]	[.014]	[-.050]
Years of schooling	.019	-.008	-.0002	.022	-.019	-.005
	(.025)	(.025)	(.049)	(.025)	(.027)	(.052)
	[.005]	[-.003]	[-.0001]	[.006]	[-.006]	[-.002]
Male	.379	.322	.755	.489	.249	.712
	(.219)	(.238)	(.439)	(.236)	(.253)	(.551)
	[.109]	[.102]	[.240]	[.139]	[.082]	[.226]
Not married	.252	.314	.470	.372	.265	.448
	(.192)	(.180)	(.390)	(.212)	(.182)	(.458)
	[.065]	[.109]	[.139]	[.091]	[.093]	[.139]

Table 6. (continued)

Equation	System A[a] (4)	System A[a] (5)	System A[a] (6)	System B[a] (4')	System B[a] (5)	System B[a] (6)
Hourly compensation before tax	.162 (.033) [.042]			.174 (.032) [.046]		
Federal marginal tax rate	.006 (.009) [-.003]			.004 (.009) [.001]		
T_{Pi}				-.007 (.008) [-.002]		
$T_{Pi}\ CL_i$	.149 (.101) [.013]					
Hourly compensation after tax		.030 (.043) [.010]	-.178 (.195) [-.055]		.010 (.050) [.003]	-.217 (.208) [-.069]
Already vested		.090 (.393) [.052]			-.035 (.393) [-.012]	
Became vested		.153 (.222) [.031]			.098 (.223) [.034]	
CL_i		-.054 (.017) [-.018]			-.056 (.018) [-.019]	
ρ_{12}		.388 (.370)			.198 (.479)	
ρ_{13}			.133 (.892)		-.084 (1.104)	
log $\mathcal{L}$		-1177.0			-1177.6	

[a]Berndt et al. asymptotic standard errors appear in parentheses and sample means of the derivative of the probability function are in brackets. Each equation contains 5 occupation and 6 industry dummies; (5) and (6) also contain 3 region dummies.

[b]Partial derivatives for tenure are reported beneath the tenure squared estimates.

Table 7. Capital loss results from single equation estimates of (5), by type of turnover.

Dependent variable	CL	$\mathcal{L}$
1. All turnover	-.056 (.018) [-.021]	-469
2. Quit to find another job	-.021 (.023) [-.005]	-316
3. Layoffs	-.067 (.026) [-.013]	-294

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