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PENSIONS, EFFICIENCY WAGES, AND JOB MOBILITY

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ABSTRACT

This paper finds that compensation premia and not pension backloading are responsible for the low mobility rates from jobs with pensions. Compensation premia, which may represent efficiency wages, are calculated as the difference in compensation between the current job and the best alternative job, allowing for the fact that such premia are observed only for job changers. The amount of pension backloading is calculated from data provided by employers to the Survey of Consumer Finances, greatly improving the precision of measurement over past efforts. This finding has important implications for labor market analysis and for policies concerning pension regulation.

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I. Introduction

It is a well-established empirical result that individuals in jobs with pensions have very low mobility rates from those jobs (Bartel and Borjas, 1977; Mitchell, 1982, 1983; McCormick and Hughes, 1984). For example, according to data from the Survey of Consumer Finances, an individual remained in the same job over a five year period 91% of the time if the job involved a pension but only 44% of the time if the job was not covered by a pension. This paper will explore empirically the causes of the negative relation between pension coverage and job mobility. In particular, is the relation due to some feature of the pension itself, or is it the result of some other characteristic of pension-covered jobs?

One explanation for this negative relation is that defined benefit plans, which are the predominant form of pensions, calculate benefits from formulae using job tenure and/or wages, and such calculations typically cause benefits to accrue disproportionately in the later years of employment (Bulow, 1981, 1982; Kotlikoff and Wise, 1985, 1987). This "backloading" of benefits produces a potentially large cost of separation for a worker who has accumulated a significant amount of tenure on the job, and this cost in turn discourages mobility from jobs with a pension of this type (Ippolito, 1986; Allen, Clark and McDermed, 1986).[1] An alternative explanation is that pension jobs may offer individuals a compensation "premium" over and above what the individual could obtain on most other jobs. Payment of compensation premia would make it relatively unlikely that those covered by pensions will be attracted to another job by a superior offer. A rationale for such premia is provided in the recent literature on efficiency wages.[2]

The lack of mobility among pension covered employees has recently created concern among policy makers that the design of pension benefit formulae is responsible for reducing the mobility of this group. A related concern is that pension backloading unduly penalizes those who do move. Despite theoretical arguments that backloading enhances an individual's productivity over the course of the job, there is no empirical evidence which links pensions or specific pension plan characteristics directly to productivity. Only indirect evidence of any such linkage is available (Allen and Clark, 1987). If pension backloading does create large barriers to mobility, and if these barriers are not justified by the productivity enhancing effects of the backloading, then policy makers may find it desirable to introduce regulations to discourage the practice of backloading pension benefits. However, this line of reasoning assumes that backloading is the major cause of low mobility rates among pension-covered workers, and to date the empirical validity of this assumption has by no means been established.

Although it has long been recognized that the financial incentives created by pensions may affect mobility, many studies of mobility simply ignore pensions. Those that do pay attention to pensions typically relate mobility or job tenure to pension coverage or to vesting status and plan characteristics, not to the value of the potential loss which a pension can generate if an individual moves to another job. The exception is Allen, Clark and McDermed (1986, 1987), which uses the dollar value of pension incentives in an equation for turnover. However, that study fails to consider the possibility of compensation premia on pension jobs, and as a result it is unable to assess the relative importance of pension backloading and compensation premia in discouraging turnover in pension-covered jobs.

The major finding of this study is that it is not the backloading of pension benefit formulae that is responsible for the negative relation between pension coverage and mobility. Those who are covered by pensions receive a higher level of compensation on their jobs than do those without pensions, and at least a part of this appears to be a compensation premium over and above what they could obtain elsewhere. It is this premium, rather than the pension loss from moving, that accounts for the lower mobility of pension covered workers. Therefore, although a reduction in backloading could be accomplished, for example, by mandating defined contribution plans, such a policy would not produce major changes in job mobility among pension covered workers unless the premia paid to these workers were to fall drastically as a result.

The organization of this paper is as follows. The next section presents descriptive statistics on the relation among mobility, pensions, and compensation. Section III introduces the analytical model used in the study, and the following section outlines the econometric method used to estimate the model. Section V discusses the empirical implementation of the model, followed in Section VI by estimates of the model. Section VII presents the results of two sets of simulations, the first of which indicates how well the model tracks the mobility behavior of different groups and the second of which analyzes the nature of the influence of pensions on job mobility. This section also examines potential reasons for the difference in findings from those of Allen, Clark and McDermed (1987). A final section summarizes the study and discusses implications of the findings for labor market analysis and for pension regulation and policies.

II. Descriptive Statistics.

This section will present some basic descriptive statistics pertinent to the relations among pensions, compensation and job mobility. The source of the data is the Survey of Consumer Finances (SCF), which is a single random cross-section sample of households taken in 1983. Detailed employment information was obtained only for the head of the household and the spouse, but since the focus of this study is middle-aged males, few of whom are not household heads, the household orientation of the survey should not present a major problem. The basic sample consists of 602 non-agricultural private-sector full-time male employees who were 30 to 50 years old in 1978 and who were not in the special high-income supplement to the SCF sample.[3] 44 of these observations are eliminated because of faulty information on experience, industry or occupation, leaving a final sample of 558.

Pensions and Mobility.

Table 1 presents rather striking results on the relation between pensions and mobility. An individual is considered to be a mover if he took a new job during the five year period immediately preceding the survey, that is, during the period 1978-83.[4] Since the individuals in the sample were 30-50 years old in 1978, the mobility being considered occurs after the turbulence at the beginning of a working career but before retirement decisions become dominant. The pension status information in the table refers to the status at the beginning of the period. Among the entire group of 558 individuals in the final sample, 28% of them moved at some point during the five year period and 72% remained with the same employer. These figures vary dramatically with pension status, however. Among individuals with pensions in 1978, over 91% remained with the employer over the next five years, while among those without pensions only about 44% stayed with the employer.

Before considering the remainder the table, it is important to mention a non-trivial problem with the SCF. The SCF, while fundamentally a cross-section survey, did attempt to gather information on each individual's job history. To be specific, after gathering information about the current job, the survey inquired about the individual's longest previous job and also about any other jobs covered by pensions. This means that information was not always collected about the job held in 1978. Since the job history did inquire about pension jobs, the pension status of the 1978 job can always be established, but unless the job was covered by a pension, was the same as the 1983 job, or was the longest prior job, no information was collected. This is the case for the 1978 jobs of about 15 percent of the sample, all of whom were movers without pensions in those jobs.

Returning to the table, the bottom part of the table reports on personal and job characteristics for those with and those without pensions in their 1978 jobs. The job characteristics reported in this table pertain to the longest job, so that comparable information is available for the entire sample. The longest job usually coincides with the job held in 1978, but even in cases where it does not we would expect the job characteristics of the 1978 job to be fairly well correlated with the characteristics of the longest job. In this table, it is evident that although pensions are strongly correlated with tendencies toward mobility, differences in other job and personal characteristics may also explain the lack of mobility from jobs covered by pensions. Although the average age, experience, and education do not differ greatly between those covered and not covered by pensions in 1978, those covered are much more likely to have held their longest jobs with manufacturing firms that were large and unionized. All three of these characteristics tend to be associated with higher compensation, which provides a further incentive to stay in the job.

Moreover, those with pensions are about 9 percentage points more likely to own a home, which would inhibit geographic mobility and may inhibit job mobility as well. Thus, while pensions are a strong potential determinant of mobility, other variables are closely related to mobility as well.

Mobility and Compensation.

Table 2 documents the compensation levels in the 1978 job and in the alternative job for various groups. The first two columns include all individuals for whom compensation can be calculated for the job held in 1978, a group comprising 396 individuals. The last two columns are limited to individuals who changed jobs between 1978 and 1983, and for whom compensation can be calculated in both jobs.

Compensation is calculated as the average per hour amount of wages plus increases in pension values between 1978 and either the individual's expected date of retirement from full-time work or the normal retirement age specified in the individual's pension plan (if he had one), whichever is earlier. If the individual did not provide an expected retirement age, the terminal date for the compensation calculations is taken to be the normal retirement age in the pension plan if the individual had one and age 65 if he did not. Real wages each year are imputed on the basis of a regression of log wages on experience, experience squared, tenure, tenure squared, interactions of both experience and tenure with education, union status and firm size, and a set of other standard explanatory variables including marital status, race, sex, health status, union status, firm size, SMSA residence, industry categories (8), and geographical regions (4). A nominal wage profile is created by using the observed wage and extrapolating to dates before and after on the basis of the estimated coefficients for the experience and tenure variables in the wage equation, taking into account the general growth of nominal wages. Pension values are calculated by applying the resulting wage profile to the individual's own pension. All compensation amounts reported in Table 2 and elsewhere in this paper are discounted to 1983 and expressed in 1983 dollars.[5]

The first two columns of the table look at all pension covered and non-covered individuals for whom compensation amounts are available for the 1978 job. The compensation differential between the two groups is quite substantial, with the log compensation amounts translating to $14.45 and $8.19, respectively. It must be noted that compensation amounts are unavailable for a nontrivial fraction of the sample, raising the possibility of selection problems. Wages are missing for about 15 percent of the 1978 jobs that are observed, and among individuals without pensions in 1978, the 1978 job is not included in the job history about 37% of the time. A comparison between the observed variables and the corresponding variables in Table 1, however, suggests that the selection process for observing compensation has not had a large effect, at least as far as these variables are concerned.

The next two columns look at job movers with and without pensions for whom observations on compensation both on the 1978 and the 1983 job are available. The compensation amounts in the original job do not appear to be very different from the amounts given in the first two columns, which included the movers and stayers combined. However, the compensation amounts in the 1983 jobs tell a much different and very interesting story. Movers without pensions in their 1978 job <u>gained</u> an average of about 17% in compensation, while movers with pensions in their 1978 job <u>lost</u> an average of 17% of their compensation. It should be noted that movers with pension in 1978 may be more likely than movers without a pension to have been separated via a layoff rather than a quit. Nevertheless, these figures would certainly suggest the possibility that the alternative job opportunities of individuals with pensions are not as renumerative, relative to the original job, as the opportunities of those without pensions. Although movers with pensions originally had a 76% compensation advantage on their 1978 job over those without pensions, the advantage dropped to only 25% in the 1983 job. This pattern is also found in the degree to which movers found new jobs with pensions. Among movers from pension jobs, 43.3% found new jobs with pensions, while among movers from jobs without pensions the figure was 37.5%.[6] It appears that movers from pension jobs were not much more successful at finding new jobs with pensions than were movers from jobs without pensions.

Table 3 attempts to decompose the differential between the 1978 job compensation of those with and without pensions. The first and last figures in the first column indicate the total per hour compensation from 1978 until retirement for the two groups; these are simply repeated from Table 2. The second figure in the column asks the question: For those with pensions, what would the compensation until retirement have been if pension amount over the lifetime of the job had been held constant, but pension value had accumulated in the fashion of a defined contribution plan? In essence, the difference between this number and the number immediately above it is the pension loss which occurs because actual pension plans accumulate value disproportionately at the end of the job rather than smoothly over the life of the job. The loss amount, amortized over the time until retirement, is only about 3%, but because these individuals have approximately 22 years until retirement, the lump sum value of the loss is a little over $17,000 in 1983 dollars. In comparison, Allen, Clark, and McDermed (1987) find an average loss for 35-44 year olds (our sample is 30-

50 years old) of $6530 in 1974 dollars, which translates to about $12,000 in 1983 dollars. This is probably some understatement because their procedure does not catch the sizable spikes in incremental pension value which occur at the early and normal retirement ages in many pension plans (Kotlikoff and Wise, 1987; and Gustman and Steinmeier, 1987), and it also seems likely that pension generosity increased to some degree between 1974 and 1983. All things considered, then, the findings of the two studies appear to agree reasonably well as to the general order of magnitude of the pension loss, and it would certainly not appear that the loss estimates used in this study are too small relative to the Allen, Clark and McDermed figures.

In Table 3, the lump sum loss is amortized over the remaining time until retirement. This would appear to be the appropriate figure to use in a study of mobility, since the question is: If the individual is separated from the present job, what would the new job have to pay in order to enable the individual to earn as much as he could if he continued in the present job? Since mobility rates from pension jobs are so low, most individuals in pension jobs will be in those jobs until close to their retirement age. Thus, although a typical individual may suffer a $17,000 pension loss if he leaves his current job, he has 22 years in which to make up this loss in the new job. At 2000 hours per year, this works out to about 38 cents per hour, which is about 2.7% of compensation. The dollar value of the pension loss may look sizable, but it is a relatively small component of the value of the job. More sizable is the value of the pension itself, as implied by the third figure in the first column. This figure gives mean log hourly earnings until retirement for individuals with pensions, excluding the value of the pensions. The difference between this figure and the one above it indicates that over the life of the job, pensions contribute about 11.9% of compensation, a percentage that agrees reasonably well with previous work (Allen, Clark and McDermed, 1986; and Gustman and Steinmeier, 1987).

The second column of Table 3 indicates the components of the differential (calculated as differences between adjacent values of log compensation in the first column), and the third column gives the percentage of the total differential of 0.568 which is accounted for by each component. The results are striking. Of the total value until retirement of a typical job, the backloading of pension benefits accounts for only about 5% of the gross difference between pension jobs and non-pension jobs. The value of pensions themselves accounts for a much larger share, almost 20%, but by far the largest part of the difference, at over 75%, is due to the fact that pension jobs typically pay much higher wages than do non-pension jobs.

III. A Model of Mobility.

This section will introduce the mobility model to be employed in the empirical analysis. The model consists of a set of four equations describing the two compensation opportunities, the mobility decision, and a selection equation. The first is a probit equation governing the mobility decision:

$$(1) \qquad M = \alpha[\ln(C_a) - \ln(C_c)] + X_1'\beta_1 + \epsilon_1$$

M is a latent variable which is positive if a job change occurs during the period and negative if it does not. $\ln(C_a) - \ln(C_c)$ is the "compensation gain" if an individual changes jobs, with C_c and C_a being the compensation levels in the current and alternative jobs in 1978. X_1 is a vector of exogenous explanatory variables, and ϵ_1 is a normally distributed random error term.

The compensation levels in the current and alternative jobs are given by the equations

$$\ln(C_c) = X_2'\beta_2 + \epsilon_2 \tag{2}$$

$$\ln(C_a) = X_3'\beta_3 + \epsilon_3 \tag{3}$$

where the X's are vectors of explanatory variables and the ϵ's are the error terms for these equations. The compensation levels include both the value of wages and the increments in the values of pensions in the two jobs. As in the last section, C_c is calculated as the average hourly amount of wages plus pension accruals from the start of the period (1978) to either the individual's expected retirement age or to the age of normal retirement in the individual's pension plan (if he has one), whichever is earlier. C_a is calculated similarly between the same dates. The rationale is that given the positive effects of tenure on wages and the typical backloading of pension benefits toward the end of the job, if an individual will ever find it advantageous to switch jobs before the normal retirement age, it will be as soon as possible. Hence the appropriate comparison is the total compensation in the two jobs from the current date until the normal retirement age in the plan or until the individual's expected retirement date if that is earlier.

As mentioned earlier, the 1978 job is not observed for all individuals because of the manner in which the job history was collected. C_c cannot be calculated for these individuals, and this requires another equation to describe whether or not it is observed. This selection equation may be written as

$$(4) \qquad I = X_4'\beta_4 + \varepsilon_4$$

where the 1978 job is observed if I is positive and is not observed otherwise. To be valid as a selection equation, X_4' must include all the explanatory variables in both the mobility and compensation equations above. Similarly, C_a is not observed for individuals who remained in their 1978 jobs at least through 1983. Normally this would require another selection equation, but the selection equation in this case is the mobility equation, which is already included in the model. Hence a separate additional equation is not required.

The model is regarded as having a correlation matrix Σ among the various ε's which is completely free. For estimation, it will be helpful to substitute from equations (2) and (3) for $\ln(C_a)$ and $\ln(C_c)$ in equation (1). This yields

$$(1') \qquad M = \alpha(X_3'\beta_3 - X_2'\beta_2) + X_1'\beta_1 + \alpha(\varepsilon_3 - \varepsilon_2) + \varepsilon_1$$

IV. The Estimation Procedure.

To facilitate the presentation of the estimation procedure, a slight change in notation will be convenient. Specifically, denote the compound error term from the mobility equation (1') to be ε_1^*, as follows

$$\varepsilon_1^* = \alpha(\varepsilon_3 - \varepsilon_2) + \varepsilon_1$$

Also, to simplify notation, let ε_i^* be equal to ε_i for the remaining equations, and denote the correlation matrix for the ε_i^* as Σ^*. The correlation matrix Σ for the original ε_i can be derived from Σ^* by straightforward calculations, if desired.

Estimates of the model are obtained by maximum likelihood. It is assumed that the error terms ϵ_i are statistically independent of the explanatory variables in the X vectors in the various equations. The likelihood function for the model is simply the product of the probability densities for the individual observations. The form of these probability densities depends on which compensation values are observed. There are three possible cases, as follows. First, consider the case where the 1978 job is included in the job history and where the individual did change jobs during the period, so that both compensation values are observed. The probability density of this observation is given as

$$Pr_j = \int_{-\hat{M}}^{\infty} \int_{-\hat{I}}^{\infty} f(\underline{\epsilon}^*) \, d\epsilon_4^* \, d\epsilon_1^*$$

where $\hat{M}$ and $\hat{I}$ are the deterministic parts of equations (1′) and (4), respectively, and j indexes the individual. $\underline{\epsilon}^*$ is a vector of the four ϵ_i^*'s, with ϵ_2^* and ϵ_3^* taking on the values solved for from equations (2) and (3) using the observed compensation values and the values of the explanatory variables.

The second case arises if the 1978 job is not observed and the individual changed jobs during the period. This implies that C_c, the compensation value for the 1978 current job, is not observed but that C_a, the compensation value for the alternative job, is observed. The probability density for the observation for this case is

$$Pr_j = \int_{-\hat{M}}^{\infty} \int_{-\infty}^{-\hat{I}} f(\epsilon_1^*, \epsilon_3^*, \epsilon_4^*) \, d\epsilon_4^* \, d\epsilon_1^*$$

This expression is different in two respects from the expression in the previous example. First the limits in the second integral are changed to reflect the lack of observation for the 1978 job. Also, the expression has in effect been integrated out with respect to ϵ_2^*, the residual in the equation determining the unobserved compensation value in 1978.

The third case occurs if the 1978 job is observed and the individual remained in the job over the period. Here, C_c is observed but C_a is not. The probability density of the observation is

$$Pr_j = \int_{-\infty}^{-\hat{M}} \int_{-\hat{I}}^{\infty} f(\epsilon_1^*, \epsilon_2^*, \epsilon_4^*)\, d\epsilon_4^*\, d\epsilon_1^*$$

This is the same kind of integral as in the previous case, except for obvious changes in the integration limits and the substitution of ϵ_2^*, which can be computed in this case, for ϵ_3^*, which cannot.

There is one data problem which is relevant to the estimation procedure. The data necessary to construct wage information for a particular job are missing in about 15 percent of the cases. An ideal solution would be to use separate selection equations for these cases, but doing so would increase the dimensionality of the cumulative normal to be evaluated by two dimensions and would make the estimation procedure computationally much more difficult. Instead, we make the assumption that the process inducing the omissions is orthogonal to the explanatory variables and error terms in the various equations. The likelihood function can then be integrated out with respect to the error terms associated with missing wages. This would cause ϵ_2^* or ϵ_3^*, depending on which wage is missing, to be dropped from the appropriate probability density formula for the observation. For example, suppose that the 1978 job is observed and a job change did occur, but that the 1978 wage is missing. The probability density for the individual in this case would be

$$Pr_j = \int_{-\hat{M}}^{\infty} \int_{-\hat{I}}^{\infty} f(\epsilon_1^*, \epsilon_3^*, \epsilon_4^*)\, d\epsilon_4^*\, d\epsilon_1^*$$

If instead the 1978 job is not observed, a job change did occur, and the alternative job wage is missing, the probability density would be

$$Pr_j = \int_{-\hat{M}}^{\infty} \int_{-\infty}^{-\hat{I}} f(\epsilon_1^*, \epsilon_4^*) \, d\epsilon_4^* \, d\epsilon_1^*$$

Having constructed the log-likelihood from the sum of the logs of the probabilities of the individual observations, maximum likelihood estimates are obtained by maximizing this function with respect to the parameters in the model, which include α, the β's, and the elements of the correlation matrix Σ. The maximization technique is a scoring algorithm with a linear search along the indicated direction in combination with the Berndt-Hall-Hall-Hausman routine for evaluating the expected second derivative matrix. This algorithm also provides asymptotic standard errors for the estimated parameters of the model.

V. Empirical Specification.

The variables included in the model are indicated along the left side of Table 4. Since most of these variables are standard, discussion of them will be brief. It may be noted that several of the employment-related variables, including industry, occupation, union status and firm size, refer to the longest job. This is necessitated because information on these variables for the 1978 job is not available for everyone in the sample, as noted previously. It is hoped that since the longest job variables refer to a major labor force experience, they will be fairly indicative of the 1978 job as well. Another temporal mismatch imposed by the nature of the data set concerns several of the household variables, including marital status, the presence of children, home ownership, and whether the wife was employed. The values of these variables pertain to 1983, the year in which the data set was collected, although again it would be preferable to use values for 1978 if that were possible.

The mobility equation is identified by excluding the firm size and pension variables. These two variables are expected to influence primarily compensation levels, and they should not have a direct effect on mobility except to the extent that they are associated with higher or lower compensation in the two jobs. There are reasonable grounds to debate the exclusion of these two variables, however, and the potential consequences of this choice are discussed in a later section. With regard to the two compensation equations, identification is achieved by excluding the household variables.[7]

Before moving to the estimates for the model, one problem with the dependent variables should be noted. The pension provider data in the survey are available only for jobs held in 1983, and even then there are some individuals who claim they are eligible for pensions but for whom pension provider records are absent in the data set, and other individuals for whom pension provider records are available but seriously deficient in some critical regard. No pension provider information at all is available for the small group of individuals who indicated a pension in 1978 but changed jobs before 1983. One way to deal with the problem is to treat missing pension provider information simply as one more cause for missing compensation observations and to deal with it in exactly the same way as missing wage information was dealt with above. However, in this case another option is used, namely to impute pensions based on pension provider information for other individuals in the sample with similar industry, occupation and union status. Specifically, the sample is divided into cells according to three-digit industries, three occupations, and union status. For any individual who is missing pension provider information, the pension component of compensation is taken as the weighted average compensation which would result from other pensions in the same cell. If there are no pension provider observations in that cell, then cells are collapsed to two-digit or one-digit industries, as required, until a non-empty cell is found. In total, imputations are made for about two-fifths (41%) of the pension values. 23% are imputed on the basis of three-digit industry cells and 7% and 9% on the basis of two-digit and one-digit industry cells, respectively, all using the correct occupation category and union status. The remaining 2% are imputed by collapsing across occupations, with 1% using three-digit industry and 1% using two-digit industry cells. In no cases are union plans used to impute nonunion plans, or vice-versa.

VI. Empirical Results.

The remainder of Table 4 presents estimates of the model. The first column reports estimates for what might be called a reduced form mobility equation, in which mobility is related to the exogenous variables in the model. This equation is estimated as a single equation probit. The most notable feature of these estimates is the overwhelming impact of the pension variable. At the means of the other explanatory variables, mobility is estimated to be about 54 percentage points lower for individuals with pensions, an effect which is in line with the descriptive statistics presented earlier. Home ownership also has a significant and sizable effect on mobility, reducing it by 23 percentage points at the means. The remaining significant variable is experience. The positive sign on this variable might at first seem surprising, but an additional year of experience is accompanied by another year of age and by a reduction in the number of years until expected retirement, so that the total effect of becoming another year older is close to nil. More surprising, perhaps, is that several variables which might be expected to affect mobility do not show up well in this data. For example, unions should depress mobility both because of high union wages and because of improvement in the employees' "voice" in the firm, and yet the estimated coefficient is insignificant and positive. Large firms usually provide high wages and enhanced promotion opportunities, both of which should reduce mobility, and yet this coefficient is also insignificant and positive. Both of these are positively correlated with pension status, and it would appear that pension status is the variable to which the effect is overwhelmingly attributed.

The remaining columns of the table present the results of the maximum likelihood estimator of the full model, as described in Section IV.[8] With this procedure, the maximum of the likelihood function is achieved at the boundary defined by the requirement that the estimated correlation matrix for the error terms (Σ^*) be positive definite. The standard errors Table 4 are therefore the result of a constrained estimation, with all of the correlation parameters except the correlation between the selection and mobility equations being treated as free, and this last correlation calculated as the value necessary to just meet the positive semidefiniteness requirement. As a result of this procedure, no standard error is estimated for this correlation, as indicated in the table, and the estimated standard errors for the remaining correlations should be interpreted with this constraint in mind.

In the mobility equation, only two of the variables are significant at standard levels. One of these, however, is the compensation gain variable which is of particular interest in this study, and its impact on mobility is by far the largest. These estimates imply that a 10% gain in this variable would result in a 8.3 percentage point increase in job mobility. The other significant variable is home ownership, which is estimated to reduce job mobility considerably. Among the remaining variables, union membership has the expected negative effect on mobility and is sizable, but not significant.

In the 1978 job compensation equation, the significant coefficients all have the expected signs, and quite a few of the coefficients are significant. Compensation is positively related to education, firm size, union membership, management/professional occupations, pension eligibility, and SMSA residency, and is lower for blacks. For compensation in an alternate job, only education has a clearly significant impact, in the positive direction, but several others hover close to significance, including union membership, SMSA residency, and white collar occupations.

The 1978 job selection equation confirms the impression that average compensation values calculated in Table 2 are not much affected by the inclusion or exclusion of the 1978 job from the job history. Only one variable, home ownership, is significant in this equation, and this variable is not a determinant of compensation. Also, the correlations between the error term in this equation and the error terms of the two compensation equations are not significant, implying that the observability of the 1978 job does not much influence compensation values through unobserved factors either.

The other correlation estimates contain few surprises. The error terms in the two compensation equations are moderately correlated at 0.45, and the correlation is significant. Also significant and relatively large, at 0.61, is the correlation between the errors in the mobility equation and the alternative job compensation equation. This would suggest that individuals whose high alternative job compensation is high for unobserved reasons are also more likely to change jobs anyway for unobserved reasons. Most striking among the correlations in Σ^*, however, is the exceedingly high negative correlation between the error terms in the 1978 job selection equation and the mobility equation. Not surprisingly, the same unobserved factors which predispose an individual toward changing jobs also make it more likely that the job history will miss the job held in 1978.

VII. Simulations

This section explores the results of two sets of simulations with the model. These simulations are performed as follows. For each individual, the nonstochastic parts of equations (1') and (4) are calculated as $\hat{M}$ and $\hat{I}$, respectively. Given $\hat{I}$ and the observation about whether or not the 1978 job is actually observed, it is possible to calculate the range of values for ϵ_4^* that is consistent with this choice (If the 1978 job involved a pension, $\hat{I}$ is infinitely positive and the whole range of ϵ_4^* is consistent with observation). Also, if compensation in either the current or alternative job is observed, it is possible to calculate the values of ϵ_2^* and ϵ_3^*, given the coefficient estimates and the values for the explanatory variables in equations (2) and (3). From these values for ϵ_2^* and ϵ_3^* and the range of values for ϵ_4^*, it is possible to calculate the conditional mean and variance for ϵ_1^* using standard multivariate normal formulae. The projected probability that the individual would have changed jobs in the five-year interval is simply the integral of the probability density for ϵ_1^* above the value of $-\hat{M}$. The simulated mobility rate for the sample is the sample average of the mobility probabilities for the individuals in the sample. In the results presented below, the simulated mobility rates are weighted averages, but some runs with unweighted averages gave very close to the same results.

Model Validation.

The first question of interest is how well the model predicts the actual mobility rate, both for the sample as a whole and for important subgroups within the sample. The simulations relevant for this question are presented in Table 5. The first column in the table gives the observed mobility rate for the group in question during the five-year period, while the middle column gives the mobility rate which the model would simulate using the explanatory variables for that particular group.

Two things are evident from the table. First, the simulations capture the disparities in mobility rates among very different groups rather well. For example, the actual five-year job mobility for individuals in the sample is about 47 percentage points higher for individuals without pensions than for individuals with pensions, and the entire amount of this differential is reflected in the simulated mobility rates. This is particularly encouraging because pensions are <u>not</u> an explicit explanatory variable in the mobility equation in the model. A similar result holds, though less dramatically, when the mobility rates are compared between union members and others. In this case, the mobility rate among union members is a bit over 12 percentage points lower than for individuals not in unions, and again the whole amount of this differential is reflected in the simulation results. Thus, the simulation model does appear to do a good job of predicting differences in the mobility rates of various groups, even of some groups not explicitly represented in the mobility equation itself.

The second thing evident from the table is that the simulations are consistently one to two percentage points high. The reasons for this are not entirely clear, but it may have something to do with the fact that the likelihood maximization occurred along a boundary. In any case, this overestimation is quite small when compared to the large scale differences in mobility rates among different groups, and it does not seem to have affected the ability of the model to predict these differences among groups successfully.

Policy Simulations of Pension Effects.

Table 6 reports on simulations which are intended to shed some light on the question: What is the role of pensions in affecting job mobility? The first part of the table conducts the following hypothetical experiment. For each individual in a pension covered job, calculate the value of the pension rights at the individual's expected retirement date or the plan's normal retirement age, whichever is earlier. Now suppose that a defined contribution plan of the same value were paid to the individual, so that the backloading which typically occurs in pension plans is eliminated. For most pension-covered individuals, this would mean that a greater percentage of the value of the pension would be accrued earlier and a smaller percentage later. In terms of the model, this would lower the value of C_c, which would presumably reduce incentives to remain on the same job. By how much would this change in the time path of compensation increase job mobility?

The answer to this question is given in the top part of Table 6, both for the sample of individuals covered by pensions as a whole and for some specific subgroups. For the group as a whole, the simulated five year mobility rate with the observed compensation is 10.3%, and with the pension changes described in the last paragraph, the mobility rate rises only to 11.6%. This small effect is in line with the fact that backloading raises compensation between 1978 and the expected retirement age by only about three percent, which is a relatively small amount. This in turn occurs because even for individuals who stay until retirement, pensions account for only about one-eighth of the total value of compensation. Also, most of the individuals in the sample are still a number of years away from retirement, and for them the elimination of the capital loss in the hypothetical experiment, pro-rated over the remainder of their years in the job, is relatively small.

One might expect the effect to be larger among the older individuals in the group, who are nearer the retirement ages specified in the plans and for whom the incentive effects of backloading should be greater. The second row in the table presents the results for this group. They are indeed more affected by the backloading, but the increase in their mobility rates is only about 2.5 percentage points over the five year period. This is still very small when compared to the nearly 50 percentage point difference in the mobility rates of those with and without pensions. Nor is the effect very large among either union or non-union members as a group, as indicated by the next two rows in the table.

The next line of the table investigates the question: What might happen if pensions were to be eliminated altogether from the compensation of individuals in pension covered jobs, and no compensating adjustments in wages or other benefits were made? This serves to eliminate both the effect of the pension backloading as well as the effect of the pension value itself from the mobility rate. As the numbers in the table indicate, eliminating pensions completely would have over four times the effect on mobility as compared simply to eliminating the backloading of pensions, but the effect is still relatively small compared to the total differential in mobility between those with and without pensions.

In the last line of the table, we analyze the effects of changing the compensation gain measure in the mobility equation by that amount which is indicated by the estimated pension coefficients in the compensation equations. The estimates of the compensation equations suggest a pension in the 1978 job reduces $\ln(C_a)$, the alternative job compensation, by 0.362. Similarly, a pension in the 1978 job increases $\ln(C_c)$, the compensation in that job, by 0.392. Hence, the effect of a pension on the difference, which is the compensation gain, is to lower it by 0.754. In other words, if the pension variable had a value of zero rather than one, the gain measure would be higher by 0.754. To simulate the effect of this change, the value of $\hat{M}$ in the simulations is simply increased by 0.754 times the coefficient of the compensation gain variable. The results of this simulation are indeed striking. Job mobility would increase to 52.3%, which is very close to the measured mobility for those without pensions, as reported in Table 1.

This result is sensitive to the estimated difference in pension coefficients in the two compensation equations, and these coefficients, particularly in the alternative job equation, are not precisely estimated. However, although the estimated difference of 0.754 may seem rather large, it is within the ball park when compared to the difference in the mean logs of compensation in the data, which from Table 2 is given as 0.568. Further, even if the difference and its associated impact on job mobility were cut in half, it would be very difficult to avoid the conclusion that the reason that mobility in pension jobs is so low is mostly because pension jobs pay a high compensation premium. Only a relatively minor role appears to be played by the fact that most pensions are typically backloaded and concentrate their benefits toward the end of the job.

Specification and Bias.

A major innovation in this model is the use of a specific equation for opportunity compensation so as to be able to construct the compensation gain measure in the mobility equation. An alternative approach, which does not require a separate equation for opportunity compensation, would be to include separately variables for current compensation and pension capital loss, as in Allen, Clark and McDermed (1987). To investigate potential biases in these approaches, note that the mobility equations in both are nested within the more general specification

$$M = \beta_1 C_o + \beta_2 C_p + \beta_3 C_l + \ldots$$

where C_o is the opportunity compensation, C_p is the compensation "premium," that is, the difference between current and opportunity compensation, and C_l is a measure of potential pension capital losses. The approach taken by Allen, Clark and McDermed, which includes current compensation in the equation without separating it into its opportunity compensation and compensation premium components, effectively imposes the constraint $\beta_1 = \beta_2$. Our own approach combines compensation premia and pension capital losses into the compensation gain variable $-(C_p + C_l)$, where C_p is a present discounted value stream. This implies the constraint $\beta_2 = \beta_3$. For reasons of indentifiability, we also omit opportunity compensation from the mobility equation, but since there is no particular reason why opportunity compensation per se should influence mobility, this exclusion should be innocuous.[9]

We believe that in the presence of premia, the constraint $\beta_2 = \beta_3$ rather than $\beta_1 = \beta_2$ is the appropriate one. From the employee's side, the compensation factor which should most strongly influence the decision to change jobs is the difference between the current job and the best alternative. Part of this difference is the pension capital loss which the individual will keep if he stays on the current job but will forfeit if he goes. Another part is the capitalized value of any premia which the individual enjoys on the current job. The individual loses both parts if he departs, and there is no reason why he should give the premium part any less weight than the capital loss part in his decision. On the employer's side, firm reputation effects may inhibit firms from laying off employees with large potential capital losses from pensions. If the purpose of potential capital losses is to induce employees to stay with the firm and/or not to shirk, however, the potential loss of the compensation premium upon separation should have the same effect. Layoffs by the firms reduce the effectiveness of either type of potential loss in discouraging mobility and shirking, and hence it should the total size of the loss, rather than the division between pension loss and premium loss, which should govern the firm's incentive not to tarnish its reputation.

There is a growing body of evidence to indicate that the possibility of large premia should at least be considered. For example, Krueger and Summers (1987) find that after standardizing for the usual human capital variables, the wage premia in two-digit industries have a standard deviation of about 15%, and that the structure of these premia is highly stable over time. In our own work, the descriptive statistics in Table 2 above suggest premia of comparable magnitude in pension jobs. Moreover, premia of this magnitude are very large in comparison to pension capital losses, since the capital losses tend to be only 3% or so of the present discounted value of compensation for most workers with pensions.

If these premia exist and are of sizable value, then the imposition of the constraint $\beta_1 = \beta_2$ may strongly bias estimates of β_1 and β_2, and of β_3 as well. A reasonable presumption is that β_1, the coefficient of the opportunity compensation, is small, since as noted before there is no particular reason why opportunity compensation per se should influence mobility. Imposing the constraint $\beta_1 = \beta_2$ when the true relation is $\beta_1 < \beta_2$ would likely result in an estimated coefficient which is less than the true value of β_2. Further, if the premia and pension capital losses are positively correlated, as seems likely, then the capital loss variable would pick up explanation which should be attributed to the premia. Thus, imposing $\beta_1 = \beta_2$ is expected to bias β_2 downward and β_3 upward.

These effects can be seen in the results of the study by Allen, Clark and McDermed (1987). They employ a mobility equation with current compensation and pension capital loss variables, which imposes $\beta_1 = \beta_2$. As expected from the preceding analysis of bias, their estimate of β_3, the effect of backloading, is much larger than ours. In a sample from the Panel Study of Income Dynamics (PSID), they estimate that backloading reduces seven-year mobility rates by 18.6 percentage points (p. 28), and in another sample from the National Longitudinal Survey (NLS), they estimate the reduction in five-year mobility rates to be 8.1 percentage points (p. 31). Also, their estimate of β_3 is implausibly large relative to the estimate of β_2, further suggesting the possibility of biases of the type discussed above. In their PSID results, for example, they find that an increase of $1000 in pension capital losses would decrease mobility by 3.7 percentage points, while an increase of $1 per hour in compensation would decrease mobility by 1.1 percentage points. This means that a $1000 capital loss is estimated to have the same impact on mobility as a $3 per hour compensation premium, despite the fact that over the 15-20 years until retirement, the $3 per hour premium will amount to around $100,000. Given that one might expect capital losses and premia to have comparable effects, such results suggest that the biases from imposing $\beta_1 = \beta_2$ may be large.

With regard to other issues of specification, the assumptions that have been made in this study have generally been such that the results would be expected to be biased in the direction of finding larger effects of pension backloading than in fact exist. First, the pension variable is omitted completely from the mobility equation. This forces the very substantial pension effect in the reduced form equation to work entirely through the compensation gain measure, thus probably biasing this coefficient upward if pensions do have a direct effect in the mobility equation or if they are proxying for other variables which should enter the that equation. Secondly, the firm size variable is also omitted from the mobility equation. Since firm size is positively related to compensation premia, this omission would tend to bias the coefficient of the compensation gain measure upward as long as the direct effect of firm size on mobility is negative. This would be expected, for example, if the greater advancement opportunities in large firms reduce mobility incentives there. Third, primarily due to a lack of identifying variables, the model has treated a dollar's worth of expected pensions as having the same value as a dollar of earnings. In fact, most people would argue that because of the greater uncertainty regarding the eventual receipt of pensions, they should be valued at some lesser amount. This would mean, however, that even more of the effect of pensions on mobility should be attributed to the wage premia and even less to pensions and backloading than we have in fact found. Finally, we have not included the possibility that those individuals in pension jobs have a lower inherent propensity to change jobs, as in Allen, Clark and McDermed (1987). This means that the part of the effect of pensions on mobility which is in fact due to heterogeneity is instead attributed to compensation and backloading, again tending to overstate the effect of backloading. Hence, we conclude that our small estimates of the effects of pension backloading on mobility are, if anything, likely to be an overstatement of the true effects.

VIII. Summary and Conclusions.

This study has investigated the relationship between economic incentives and job mobility. The model used to estimate this relationship contains two important features which are not usually found in other studies of job mobility. For one, it uses a detailed description of the pension plan, as provided by the employers of the individuals in the sample, to construct accurate measures of the financial incentives against job mobility which pension plans are widely thought to provide. Also, the propensity to change jobs is characterized as depending, among other things, on the difference in compensation that an individual can obtain on his current job and on the best alternative job.

Two strong conclusions can be drawn from the results of this paper. The first of these has to do with estimation: the failure to separate compensation into its opportunity compensation and compensation premium components may severely bias the estimated effects of economic incentives in mobility equations. Specifically, estimated effects of pension capital losses may be sharply overestimated if possible premia in pension jobs are not taken into account. The second conclusion is more of a policy nature, namely, that the role of the typical backloading of pensions in restricting job mobility seems fairly small for most individuals. Although the estimates indicate that potential compensation gains in an alternative job have a strong effect in reducing mobility, the small size of the backloading relative to compensation premia in pension jobs means that only a small amount of the difference in mobility rates between individuals with and without pensions is attributable to the backloading.

The findings reported here have some implications for the ongoing debate as to the relative importance of tilting the compensation profile and using efficiency wages as devices for raising worker productivity. The examination of compensation on current and alternative jobs suggests that on pension covered jobs, compensation premia constitute a much more important fraction of the loss from job termination than does backloading of pension benefits. Simulations with the mobility equation confirm that it is not backloading, but wage premia that accounts for the large difference in mobility between pension covered and noncovered jobs. In so far as the compensation tilt and efficiency wage models incorporate penalties from turnover and worker response to these penalties, these findings provide greater support for the efficiency wage view.

Footnotes

1. The expression "backloading of benefits" is sometimes used to refer to a weighting scheme whereby the pension formula explicitly gives greater weight to later than to earlier years of employment. In the context of this paper, backloading refers to the positive slope of the accrual profile that results even when all years of work receive equal weight in the pension benefits formula.

2. Efficiency wages refer to wages paid in excess of the competitive wage by profit maximizing firms in order to increase productivity. For a discussion, see Krueger and Summers (1986). For a discussion of related work by earlier generations of labor economists, see Segal (1986).

3. Agricultural and self-employment are screened on the basis of the individual's longest job, since information on these characteristics is not always available for the 1978 job.

4. If the individual was unemployed in 1983, he was considered a mover if previous job began after 1978, and a stayer if his previous job began before 1978. This effectively measures mobility as the taking of a new job, and not as the separation (which may or may not be temporary) from an old job.

5. The general growth rate of nominal wages used in the calculations is the 30 year average from 1953-83, and the discount factor is taken to be equal to the general growth rate of nominal wages. For the procedure followed if the individual's pension record was missing or defective, see Section IV.

6. These figures are based on all movers in the sample, regardless of whether compensation amounts are observed.

7. Since the same variables are included in the two compensation equations, the vectors X_2 and X_3 are identical empirically. We continue the notational distinction in order to maintain the correspondence between the X vectors and the β vectors.

8. By the construction of the data set, all 1978 jobs with pensions are observed. Hence, the pension coefficient in the 1978 job selection equation has an implied value of $+\infty$ and is not estimated.

9. Given the problem with the unobserved 1978 jobs in the SCF, including opportunity compensation would require equations to project compensation gain and opportunity compensation separately in equation (1'), and we lack sufficient good instruments to do so.

Table 1
Pensions and Job Mobility

	With Pension	Without Pension	Both
Mobility rate in percent	8.8%	55.9%	28.0%
Average age (in 1978)	39.7	39.3	39.5
Average experience (in 1978)	20.2	19.9	20.1
Average education	12.8	12.1	12.5
Percent in manufacturing in longest job	48.3	26.4	39.4
Percent white collar in longest job	8.2	10.1	9.0
Percent management and professional in longest job	30.8	31.7	31.2
Percent union in longest job	48.9	18.5	36.6
Percent in firms larger than 100 employees in longest job	87.3	49.3	71.7
Percent residing in an SMSA	64.7	51.1	59.1
Average years until expected retirement (in 1978)	22.3	24.0	23.0
Percent married	89.1	85.9	87.8
Percent with children under 18	61.6	67.0	63.8
Percent who own home	85.2	76.2	81.5
Percent with employed spouses	55.6	59.0	57.0
Percent black	6.0	9.3	7.4
Number of Observations	331	227	558

Table 2
Pensions and Compensation

	All Individuals with Observations for 1978 Job		Job Changers with Observations for Both Jobs	
	With Pensions	Without Pensions	With Pensions	Without Pensions
Log of average discounted hourly compensation in 1978 job to retirement (standard deviation)	2.67 (0.47)	2.10 (0.64)	2.70 (0.33)	2.13 (0.62)
Log of average discounted hourly compensation in alternative job to retirement (standard deviation)			2.51 (0.54)	2.29 (0.64)
Average age (in 1978)	39.3	39.2	38.7	38.4
Average experience (in 1978)	19.9	19.4	19.3	20.7
Average education	12.8	11.9	13.1	11.7
Percent in manufacturing in longest job	48.3	20.0	55.6	23.3
Percent white collar in longest job	8.0	7.3	0.0	6.7
Percent management and professional in longest job	29.4	36.4	38.9	33.3
Percent union in longest job	50.0	15.5	72.2	10.0
Percent in firms larger than 100 employees in longest job	86.4	40.0	77.8	40.0
Percent residing in an SMSA	63.3	51.8	72.2	60.0
Average years until expected retirement (in 1978)	22.6	24.4	23.9	26.1
Percent married	89.5	84.5	77.8	90.0
Percent with children under 18	63.3	70.9	72.2	80.0
Percent who own home	83.9	74.5	44.4	76.7
Percent with employed spouses	57.3	59.1	44.4	63.3
Percent black	6.3	10.9	0.0	3.3
Number of Observations	286	110	18	30

Table 3
Decomposition of Pension/Non-Pension
Compensation Differential

	Compensation Level	Components of Differential	Percent of Total Differential
Mean Log of Compensation of Individuals with Pensions	2.671		
		0.027	4.8%
Mean Log of Compensation of Individuals with Pensions, Excluding Pension Tilt	2.644		
		0.112	19.7
Mean Log of Compensation of Individuals with Pensions, Excluding Value of Pensions	2.532		
		0.429	75.5
Mean Log of Compensation of Individuals without Pensions	2.103		
Total		0.568	100.0%

Table 4
Model Estimates

	Reduced Form Mobility Equation	Full Model: Mobility Equation	Full Model: Compensation Eqns. 1978 Job	Full Model: Compensation Eqns. Alternative Job	Full Model: 1978 Job Selection Equation
Constant	-0.253 (0.17)	0.431 (0.16)	1.248 (4.63)	0.786 (0.72)	-1.164 (0.46)
Years of Experience in 1978	0.056 (2.61) [0.02]	0.044 (1.29) [0.02]	-0.004 (0.93)	0.000 (0.02)	-0.028 (0.95)
Years of Education in 1978	0.030 (0.96) [0.01]	-0.034 (0.48) [-0.01]	0.087 (8.07)	0.122 (4.16)	-0.005 (0.11)
* Manufacturing in Longest Job	0.166 (1.04) [0.06]	0.279 (0.86) [0.12]	0.074 (1.42)	0.055 (0.39)	-0.034 (0.14)
* White Collar in Longest Job	0.316 (1.22) [0.11]	-0.398 (0.78) [-0.17]	0.064 (0.78)	0.454 (1.85)	-0.500 (1.39)
* Management/Professional in Longest Job	-0.028 (0.15) [-0.01]	0.145 (0.47) [0.06]	0.226 (3.59)	0.135 (0.97)	0.185 (0.64)
* Union in Longest Job	0.023 (0.12) [0.01]	-0.342 (0.98) [-0.15]	0.110 (2.10)	0.304 (1.70)	-0.087 (0.32)
Years Until Expected Retirement	0.027 (1.48) [0.01]	0.016 (0.45) [0.01]	-0.006 (1.35)	0.001 (0.07)	0.001 (0.03)
* SMSA in 1983	0.098 (0.67) [0.03]	-0.175 (0.64) [-0.08]	0.109 (2.32)	0.228 (1.86)	0.170 (0.80)
* Race (Black) in 1983	-0.488 (1.70) [0.16]	0.180 (0.23) [0.08]	-0.186 (2.12)	-0.513 (1.37)	0.287 (0.74)
* Firm Size > 100 in Longest Job	0.095 (0.64) [0.03]		0.120 (2.14)	-0.019 (0.19)	-0.279 (1.27)
* Pension Coverage in 1978 Job	-1.610 (10.03) [0.54]		0.392 (5.02)	-0.362 (1.59)	

Table 4 (continued)

	Reduced Form Mobility Equation	Full Model: Mobility Equation	Full Model: Compensation Eqns. 1978 Job	Full Model: Compensation Eqns. Alternative Job	Full Model: 1978 Job Selection Equation
Compensation Gain (α)		1.923 (2.92) [0.83]			
Age in 1978	0.027 (0.88) [0.01]	-0.020 (0.56) [-0.01]			0.037 (0.70)
* Married in 1983	0.084 (0.32) [0.03]	0.192 (0.64) [0.08]			0.010 (0.02)
* Children Under 18 in 1983	0.205 (1.18) [0.07]	0.119 (0.62) [0.05]			0.028 (0.10)
* Home Ownership in 1983	-0.668 (4.01) [-0.23]	-0.773 (3.90) [-0.33]			0.600 (2.34)
* Wife Employed in 1983	0.038 (0.24) [0.01]	-0.071 (0.44) [-0.03]			0.054 (0.22)
Standard Deviation of Error Term	1.00 --	1.00 --	0.411 (26.91)	0.565 (9.00)	1.00 --
		1.00 --	0.14 (1.01)	0.61 (3.14)	-0.95 --
Correlation Matrix (Σ^*)			1.00 --	0.45 (3.47)	0.03 (0.11)
				1.00 --	-0.34 (1.72)
					1.00 --
Log-likelihood	-237.11	-597.99			
Number of Observations	558	558			

Note: Figures in parentheses are asymptotic absolute t-statistics. Figures in brackets are derivatives at the means, where appropriate. Variables with asterisks are binary variables.

Table 5
Model Validation Simulations

	sample mobility rate	simulated mobility rate	number of observations
Full Sample	28.0%	29.8%	558
Individuals with Pensions	8.8	10.1	331
Individuals without Pensions	55.9	58.4	227
Union Members	20.1	21.2	204
Not Union Members	32.5	34.9	354

Table 6
Simulations of Changes in Compensation

	original mobility rate	post-change mobility rate	number of observations
Effects of Converting Defined Benefit Plans to Defined Contribution Plans of Equal Value			
All Individuals with Pensions	10.3%	11.6%	558
46-50 Year Olds	9.8	12.3	54
Union Members	11.8	13.3	143
Not Union Members	8.9	9.8	143
Effects of Dropping Pension Compensation Entirely	10.3	15.7	286
Effects of Higher Compensation in Pension Jobs, As Measured by Compensation Equation Estimates	10.3	52.3	286

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