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Estimating the Relationship between Employer-Provided Health Insurance, Worker Mobility, and Wages

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This document reports the results of research and analysis undertaken by the U.S. Census Bureau staff. It has undergone a Census Bureau review more limited in scope than that given to official Census Bureau publications. [This document is released to inform interested parties of ongoing research and to encourage discussion of work in progress.] This research is a part of the U.S. Census Bureau's Longitudinal Employer-Household Dynamics Program (LEHD), which is partially supported by the National Science Foundation Grant SES-9978093 to Cornell University (Cornell Institute for Social and Economic Research), the National Institute on Aging, and the Alfred P. Sloan Foundation. The views expressed herein are attributable only to the author(s) and do not represent the views of the U.S. Census Bureau, its program sponsors or data providers. Some or all of the data used in this paper are confidential data from the LEHD Program. The U.S. Census Bureau is preparing to support external researchers' use of these data; please contact U.S. Census Bureau, LEHD Program, Demographic Surveys Division, FOB 3, Room 2138, 4700 Silver Hill Rd., Suitland, MD 20233, USA.

Estimating the Relationship between Employer-Provided Health Insurance, Worker Mobility, and Wages*

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Abstract

In this paper, a joint model of wages, hazard of a job ending, and probability of holding employer-provided health insurance is estimated, taking account of unobservable person and job characteristics. A unique data source, the 1990 and 1996 SIPP Panels linked to SSA administrative job histories, enables the identification of random person and job effects and the correlation of these effects across the three equations. The explicit modeling of this correlation produces consistent estimates of the effect of tenure on wages and the effect of health insurance on mobility. Substantial levels of job-lock and significant annual returns to seniority are found. Increasing the job-specific probability of obtaining employer-provided health insurance from 60% to 63%, or increasing the job-specific hourly wage rate by $.80, are both associated with an equivalent decrease in the hazard of the job ending. However, the dollar value of the wage benefit is substantially higher.

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1 Introduction

The relationship between compensation and length of time on the job has been modeled theoretically and tested empirically many times. Economists have recognized that firms may design compensation packages with the explicit intention of influencing decisions by workers concerning job duration and have sought to quantify the effects of particular types of compensation. Both theoretically and empirically, researchers have acknowledged the necessity of distinguishing between types of people and types of jobs that appear observationally equivalent but in fact have unobservable characteristics that influence both compensation and job duration.

In this field, the relationship between tenure and wages has been one of the most extensively debated topics. Wages are observed to rise with job tenure above and beyond what can be attributed to accumulation of general labor market experience and quit probabilities are observed to decline with job tenure. Explanations for these empirical observations have included the development of firm-specific human capital, learning about job match quality, search, and individual heterogeneity. In evaluating these theories and determining whether tenure has a "true" effect on wages, controlling for individual and job quit propensities becomes essential.

A more recent debate has focused on the relationship between employer-provided health insurance and job duration. Health insurance has been termed "non-portable" because it is a job-specific benefit that is lost when a job ends. This non-portability feature, combined with the high incidence of people in the United States whose sole source of health insurance is job-related, has given rise to the hypothesis that some workers may be "job-locked." The core idea behind the job-lock theory is that workers may remain in job matches that would otherwise have been dissolved due to the possibility that health insurance may not be available at a new job. Workers with pre-existing conditions or families with large medical expenses have been thought to be the most vulnerable to job-lock. However estimating a job-lock effect is complicated by the same person and job heterogeneity which influences wages. An observed decrease in mobility rates for people at jobs with health insurance may be due to the lack of portability of their health insurance, the high quality of their jobs, or their personal preferences for mobility.

This paper seeks to combine aspects of both these literatures by treating wages, job tenure, and employer-provided health insurance as three outcomes that result from the interaction of worker and firm choices. Using a statistical model proposed by Lillard (1999), this paper will estimate the relationship between these three variables in a way that treats each of them as endogenous and determined by both observable and unobservable person and job characteristics. Unobservable characteristics will be modeled using random person and job effects that will be correlated across equations, the feature of the model that prevents bias. The results from this statistical model will answer important questions about the relationship of various components of the compensation package to job tenure. Does holding a health insurance policy from an employer make an employee less likely to quit, even after controlling for person and job type and the fact that health insurance represents higher compensation? Are career high-wage individuals at high-wage jobs more or less mobile than career low-wage workers at low-wage jobs? Does longer job tenure increase wages in a causal way or is the correlation between tenure and wages the result of sorting heterogeneous workers into heterogeneous jobs? Finally, how do wages and health insurance compare in their effects on mobility?

Using the 1990 and 1996 Survey of Income and Program Participation (SIPP) Panels linked to administrative job earnings and histories from the Social Security Administration, I am able to construct detailed monthly information about my three outcome variables of interest: wages, employer-provided health insurance, and job tenure. My results show substantial levels of job-lock and returns to tenure of 3.6% during the second through fourth years of a job and 1.5%-2% per year thereafter. High-wage workers are observed to be more mobile than low-wage workers while high-wage jobs are observed to last longer than low-wage jobs. A 5% log wage premium at a job reduces the hazard of the job ending by 2.3%. A rise in the probability of obtaining health insurance at a job from 60% to 63% reduces the hazard of that job ending by an equivalent amount even though this benefit costs the employer on average only 16% as much as the wage increase. These results point to the influential role of benefits as part of the compensation package and highlight the need for studies on the firm side to understand how profit-maximizing firms make choices about turnover and compensation.

2 Literature Review

2.1 Wages and Tenure

The literature on wage determination offers many theories for how tenure and wages may be related. One class of theories predicts tenure should have no effect after controlling for individual and match heterogeneity. If the work force is composed of "movers" and "stayers," mobility may decline with tenure because of a selection effect. Wages then rise faster for "stayers" because they are more likely to receive training by the firm. Search theory suggests that workers look for jobs that are good matches, where good matches are defined as high productivity pairings of workers and firms. These good matches pay more and hence raise the opportunity cost of leaving the job. The mobility of these workers is reduced because of the decreasing probability of receiving a better offer (Jovanovic (1979), Burdett (1978)). Workers in bad matches with low earnings switch jobs in order to find better jobs and raise earnings. Hence the observation of a positive correlation between wages and tenure is simply a manifestation of match heterogeneity. On the other hand, human capital models predict that tenure affects wages because firm-specific human capital is accumulated and rewarded (Mincer (1974)). Search models also predict an effect of tenure if it takes time to determine the quality of the match (Jovanovic, (1979)).

Traditionally the relationship between wages and tenure has been estimated by including tenure as an explanatory variable in a wage equation. However as researchers began to view tenure as endogenous, they developed new estimation methods to prevent biased results. Abraham and Farber (1987) use expected completed job duration to instrument for seniority and find a return of .25%-.5% per year. Altonji and Shakotko (1987) also propose an instrumental variables approach, using deviations of the tenure variables around their means for a given job match as the instrument for tenure. They find a similarly small return to tenure of .6% per year. All these authors compare their results to a traditional return to seniority of 1% per year and conclude that person and job heterogeneity plays a role in the wage determination process. Topel (1991) also employs techniques to account for person and job match heterogeneity but finds a return of 3% per year to seniority and concludes that job-specific human capital is an important component of wages. Abowd et. al (1999) argue that person and firm heterogeneity should be directly modeled in the wage equation and that seniority should be treated as a firm-specific time-varying effect. They estimate a 1% return to a year of seniority. Finally, a study by Topel and Ward (1992) gives further credence to the belief that wages and tenure are both endogenous outcomes by showing that the relationship between wages and tenure work in both directions. They estimate that a 10% increase in wages reduces the probability of changing jobs by 20%, while at the same time, job changes account for one third of total wage growth for male workers during the first ten years in the labor market.

Results of this type lead to the consideration of an approach proposed by Lillard (1999). He estimates both a wage equation and a tenure equation, each including random person and job effects. By allowing these effects to be correlated across the two equations, he obtains a consistent estimate of the effect of tenure on wages. He reports a return to job tenure of 5% for the first year, .6% for the second year, and .36% each year thereafter. Dostie (2001) estimates a similar model using French linked employer-employee data and finds no return to seniority.

2.2 Employer-Provided Health Insurance and Quit Decisions

A parallel literature has investigated the relationship between employer-provided health insurance and quit decisions. Until the mid-1980s, health insurance was a completely non-portable job benefit. COBRA legislation in 1985 made it possible to retain the old insurance for up to eighteen months if the worker paid 102% of the average cost to the old employer of continuing to provide the insurance (Madrian (1994)). However, for workers who faced pre-existing conditions clauses at new jobs that made them or their family members ineligible for new insurance, COBRA was only a temporary and expensive solution to the portability problem. In 1996, Congress passed HIPPAA, a new law that, among other reforms, limited the amount of time workers could be denied coverage for pre-existing conditions. These limits were based on how long the worker had previously held health insurance.

The literature on job-lock has attempted to measure the decline in mobility directly resulting from prohibitively high costs of losing employer-provided health insurance. This decrease in mobility is potentially efficiency decreasing. Unsuccessful job matches are perpetuated when their dissolution would have benefited the worker and the firm. In measuring job-lock, the thought experiment becomes, "How much would the probability of leaving a job decline, holding all job and worker characteristics constant, if the compensation package changed from not providing health insurance to providing health insurance." This total change in the quit probability would be due to a compensation component and a job-lock component. Hence the difficulty in actually performing the measurement is three-fold: holding the type of person constant, holding the type of job constant, and differentiating the compensation effect from the job-lock effect. Problems arise because jobs that offer health insurance are likely to be "good" jobs and to employ "good" workers. Unobserved job and worker heterogeneity with regard to quit rates is correlated with the presence of employer-provided health insurance and unless one controls for this heterogeneity, the coefficient on the health insurance indicator will be biased. Insured workers will quit less often because of the quality of their jobs, their personal preferences for mobility, their relatively higher compensation in the form of insurance benefits, and the non-portability of their health insurance. Thus the goal of the literature has been to find a method for dealing with the endogeneity of the health insurance variable and to separately identify the effect of health insurance from the effect of the type of job, type of person, and compensation levels.

The most popular approach comes from Madrian (1994) and involves a difference-in-difference estimator. Using a probit model and data from the 1987 National Medical Expenditure Survey (NMES), Madrian estimates the probability that a worker who is employed at the beginning of the survey will end the job by the date of the second interview one year later. She includes indicator variables for holding employer-provided health insurance, and other, non-employer provided health insurance, as well as an interaction term between these two types of health insurance. She then calculates the probability of quitting a job for workers in four different groups: M_{11},those with both other and employer-provided insurance, M_{01}, those with only employer-provided insurance, M_{10}, those with only other insurance, and M_{00}, those with no insurance. Her difference-in-difference estimator is calculated as

$$(M_{11} - M_{01}) - (M_{10} - M_{00}).$$

Thus workers with employer-provided health insurance and presumably similar higher levels of compensation are compared to each other across a dimension thought to reduce job lock while controlling for the independent effect of this treatment effect. If job-lock exists then the difference-in-difference should be positive. Having other insurance should cause a greater change in mobility for those with employer-provided insurance than those with no insurance. She estimates that job lock reduces mobility by 31%. Others have obtained different estimates of job-lock using similar methods to Madrian but different data. For instance, Holtz-Eakin (1994) finds insignificant amounts of job-lock when he estimates a difference-in-difference model for married men with and without spousal insurance using the 1984 wave of the Panel Study of Income Dynamics (PSID).

The common critique of this approach is to suggest that unobservables about jobs and people are not differenced away and hence additional variables should be included in the probability model. For instance, Buchmueller and Valletta (1996) argue that workers with insurance through their spouses may have been offered insurance through their own employers and turned it down. These workers may thus have better jobs than those with no insurance and hence have lower mobility. This would lower the difference $(M_{10} - M_{00})$ in a way unrelated to the effect of insurance from another source. They propose additionally including pension coverage and tenure in the probit model to capture some of the firm and person unobserved heterogeneity. Using the 1984 Survey of Income and Program Participation (SIPP), they estimate that employer-provided insurance reduces turnover by 35-40% for women, but find insignificant effects for men.

Kapur (1998) offers further refinement of the difference-in-difference estimate using the 1987 NMES. In order to insure the comparability of the control and treatment groups, she uses only married men with employer-provided health insurance in her estimation. Those men without spousal health insurance form the experimental group and those with this additional form of health insurance are the control group. She then estimates a difference-in-difference model with the treatment variable being family sickness. Using this method, she finds no evidence of significant levels of job-lock.

Two papers offer estimation strategies other than the difference-in-difference model. Monheit and Cooper (1994) use the 1987 NMES to estimate reduced form wage and health insurance equations in order to predict compensation at a new job using data from voluntary job changers. Using the predicted likelihood of obtaining health insurance coverage at a new job, workers were classified into one of three categories: gaining insurance, losing insurance, or no change in insurance status. The predicted changes in wages and health insurance status were then included as explanatory variables in the probit quit model, along with other job characteristics (paid sick leave, percentage in the worker's industry covered by pension plans, union membership) and worker characteristics (insurance coverage, spousal employment status, family size, own or family member health problems). Since the health insurance change variable was estimated from individual and labor market characteristics, the authors maintain that it is not contaminated by unobservable job attributes. They find that those expected to lose coverage by changing jobs were between 3%-6% less likely to switch.

In the second paper, Gilleskie and Lutz (1999) propose testing for job-lock by including both offer and acceptance health insurance indicators in a job transition model. They offer two interpretations for the offered insurance variable.

First, it could serve as a proxy for "good" jobs and measure the impact of job heterogeneity on the mobility decision separately from the impact of actually holding the non-portable health insurance benefit. Alternatively, it could represent what the authors term an "option-value," meaning that the offer holds value to the individual because of the potential to hold health insurance in the future, regardless of current take-up choices. Thus even an offer has the potential to cause job-lock. Using the NLSY from 1989-1993, the authors estimate a dynamic multinomial logit function which represents the likelihood of transitioning from the current job to each of the three possible future states. Without controlling for other benefits or health insurance offers, there is a 31% drop in job changes for married men when the individual has employer-provided insurance. The inclusion of an insurance offer variable reduces this to 12% and the additional inclusion of other benefit availability makes both the holds insurance and offered insurance variables insignificant. To test the robustness of their results, Gilleskie and Lutz estimate a joint probability model of initial tenure, employment status, marital status, the offer of employer provided health insurance, the holding of employer provided health insurance, the holding of health insurance from another source, and the employment transition decision and model unobserved individual heterogeneity as a person-specific random effect. This model again produces no evidence of job-lock for married men.

This literature on job-lock has arisen mainly in response to concerns about how to control for heterogeneity and how to account for the compensation effect of health insurance in lowering quit probabilities. Although many new types of controls have been used in the literature to solve these problems, some concerns still remain. The inclusion of tenure as an exogenous explanatory variable in order to control for person heterogeneity is problematic because tenure is the result of a sequential set of separation decisions, each of which is correlated with the provision of health insurance by the employer. Benefit variables such as pension coverage, included to control for job heterogeneity, may themselves be a source of job-lock and hence may control for more than just positive job characteristics. Since all data sets previously used had only one observation per job, distinguishing between job type and the particular effects of health insurance is difficult. Gilleskie and Lutz, for example, are only able to identify the health insurance equation separately from the job transition equation by using body mass index, number of jobs held (itself endogenous), and health limitations. Compensation effects are also possibly not well accounted for in the models previously discussed. The difference-in-difference model assumes that the compensation associated with employer-provided health insurance is constant across groups. However there is almost certainly a great deal of heterogeneity in the type and cost of offered insurance. Problems may arise when comparing employer-insured workers with and without health insurance through another person if the employer-provided policies are, on average, very different across the two groups. Those who chose to "double-insure" may have done so because their employer-provided policy had less generous benefits or more expensive premiums. Hence they are more mobile because they in essence receive less compensation. Another concern is that those with dual insurance sources may not tend to insure other family members as often as those with only employer-provided insurance. In addition to the direct effects of dependent family members on mobility, a family health insurance plan is worth substantially more than a single coverage plan and hence again represents higher compensation. These possible differences in compensation may artificially inflate the mobility of those with dual insurance relative to those with only employer-provided insurance. A final component of compensation which is also a concern is wages. Wages are universally included as exogenous variables in quit decision models although they are most likely jointly determined with tenure decisions.

3 Model

3.1 Estimating Health Insurance and Tenure Effects

My model will answer questions central to both the tenure and job-lock literatures and will uniquely contribute to each by exploring the relationship between compensation and tenure. Following Lillard, I will view quit decisions, employer-provided health insurance, and wages as three outcomes of a joint process and hence treat each one as endogenous. Since the nature of my data will allow the estimation of a job duration model, I will use a hazard model formulation to estimate the probability of separating conditional on the observed past job history. This essentially allows the estimation of a series of separation decisions over time. In estimating this system of equations, I will explicitly allow for both person and job heterogeneity by including random individual and job effects, with correlation between these random effects across equations. This correlation will account for the fact that workers with individual and job specific propensities for low turnover may have similar propensities for holding employer-provided health insurance and receiving high wages. Identifying job heterogeneity is made possible by the presence of multiple wage and health insurance outcomes for each job. This method will contribute to the job-lock literature by controlling for job heterogeneity in a new way, taking account of the effect of wages on job tenure in a way that does not assume that wages are exogenous, and modeling the actual tenure decision and not the probability of every individual quitting after the same arbitrary amount of time. The tenure literature will be advanced by the inclusion of other types of compensation in the tenure equation as well as the estimation of Lillard's model using another data source.

The effect of tenure on wages will be estimated using the coefficients on a tenure spline in the wage equation and the estimated effect will be consistent because of the heterogeneity controls. Evidence for the search or human capital theories of wage growth will come from the correlations of the random person and job effects across the tenure and wage equations. Testing for job-lock will not be quite as straight forward. I include an indicator variable for whether a person has employer-provided health insurance at a job in the hazard model for job duration and argue that the coefficient on this variable is consistent because of the random person and job effects. However this coefficient still contains the compensation effects of holding employer-provided health insurance. Thus to obtain an accurate measure of job-lock, I use two methods of controlling for the monetary value of this benefit. First, I calculate a difference-in-difference effect, following Madrian, where workers with employer-provided health insurance make up the experimental group, workers without this insurance make up the control group, and having health insurance through another person serves as the treatment. This method allows me to assess how my results compare to those of the literature and determine whether my controls for person and job heterogeneity reduce measured job-lock.

My second, and unique, method of testing for job-lock will utilize the estimated correlation between the job specific random effects from each of the equations. These correlations allow me to compare the impact of above average wages at a job on the hazard of ending the job with the impact of an above average probability of having health insurance and assess the relative magnitudes of these two effects. If health insurance is worth more to workers than an equivalent dollar amount of wages and we assume workers are being paid their marginal products, then this is evidence of rigidities in the labor market. Workers value their jobs more than the employer values their output and hence workers will be hesitant to leave for a more productive match if the compensation at the alternative job does not include the same mix of health insurance and wages.

Finally, in order to test the validity of the previously outlined concerns, I perform several specification checks where I control more specifically for what type of health insurance workers obtain from their employers, family or single coverage, and how much of the cost of health insurance the employer pays.

3.2 Econometric Model

Following Lillard's (1999) two equation tenure and wage model, I propose the following three equation model:

$$
\begin{aligned}
HIEMP_{ijt} &= 0 \text{ if } HIEMP^* = \beta_{H1}X_i + \beta_{H2}X_{ij} + \beta_{H3}X_{ijt} + \gamma_i + \varepsilon_{ij} + \xi_{ijt} < 0 \\
&= 1 \text{ if } HIEMP^* = \beta_{H1}X_i + \beta_{H2}X_{ij} + \beta_{H3}X_{ijt} + \gamma_i + \varepsilon_{ij} + \xi_{ijt} \geq 0
\end{aligned} \tag{1}
$$

$$
\begin{aligned}
h(t_{ij}) &= e^{\gamma T(t_{ij}) + \alpha HIEMP_{ijt} + \beta_{HZ1}X_i + \beta_{HZ2}X_{ij} + \beta_{HZ3}X_{ijt} + \delta_i + \lambda_1 \varepsilon_{ij} + \lambda_2 \psi_{ij}} \\
S(t_{ij}) &= \exp\{-\int_0^{t_{ij}} h(\tau_{ij}) d\tau\} \\
f(t_{ij}) &= h(t_{ij}) S(t_{ij})
\end{aligned} \tag{2}
$$

$$
\ln(w_{ijt}) = \beta_{W1}X_i + \beta_{W2}X_{ij} + \beta_{W3}X_{ijt} + \theta_i + \psi_{ij} + \eta_{ijt} \tag{3}
$$

where i is the person subscript, j is the job subscript, and t is the month subscript. The variables in the model are defined as follows:

$$
\begin{array}{rcl}
t_{ij} & = & \text{duration of job } j \text{ for person } i\text{, time-varying across jobs} \\
HIEMP_{ijt} & = & \text{employer provided health insurance,} \\
 & & \text{time-varying across and within jobs} \\
\ln(w_{ijt}) & = & \text{natural log of hourly wage, time-varying across and within jobs} \\
(\delta_i, \gamma_i, \theta_i) & = & \text{individual heterogeneity terms} \\
(\varepsilon_{ij}, \psi_{ij}) & = & \text{job heterogeneity terms} \\
X_i & = & \text{person characteristics, time invariant across and within jobs -} \\
 & & \text{all equations - white, Hispanic, male, schooling} \\
 & & \text{health insurance equation - health condition} \\
X_{ij} & = & \text{job characteristics - time-varying across jobs} \\
 & & \text{all equations - union status, industry, job type} \\
X_{ijt} & = & \text{person and job characteristics, time-varying within jobs -} \\
 & & \text{all equations - marriage, number of kids} \\
 & & \text{job duration equation - any health insurance coverage,} \\
 & & \text{coverage from other person, interaction of coverage from} \\
 & & \text{other person and coverage from employer} \\
 & & \text{health insurance equation - hours worked per week, age begin job} \\
 & & \text{wage equation - general labor force experience, tenure, time} \\
T(\cdot) & = & \text{linear splines in age, job tenure, calendar time,} \\
 & & \text{general labor force experience which form baseline hazard} \\
\xi_{ijt} & = & \text{time-varying component of the probit error, iid across months} \\
\eta_{ijt} & = & \text{time-varying component of the wage error, iid across months}
\end{array}
$$

The problem in the estimation process occurs because the health insurance variable in the tenure equation (HIEMP in equation 2) is endogenous and determined by person and job heterogeneity as shown in equation 1. If person heterogeneity in equation 1, δ_i, is correlated with person heterogeneity in equation 2, γ_i, then HIEMP is correlated with δ_i and α will be inconsistent. Thus joint estimation of the three equations, allowing for cross-equation correlation of the heterogeneity terms, is necessary.

This system of equations is simultaneous in the heterogeneity parameters and is triangular in structure. HIEMP enters the job duration equation and tenure enters the wage equation. However wages are restricted to affecting tenure and health insurance through the correlation across the job and individual heterogeneity terms. Tenure affects health insurance in a similar way.

In controlling for worker and job heterogeneity, I will model the individual heterogeneity terms, δ_i, γ_i , and θ_i, as random effects that are jointly normally distributed with an unrestricted variance-covariance matrix.

$$\begin{aligned}(\delta_i, \gamma_i, \theta_i) &\sim N(0, \Sigma_{\delta\gamma\theta}) \\ \Sigma_{\delta\gamma\theta} &= \begin{bmatrix} \sigma_\delta^2 & \sigma_{\delta\gamma} & \sigma_{\delta\theta} \\ \sigma_{\gamma\delta} & \sigma_\gamma^2 & \sigma_{\gamma\theta} \\ \sigma_{\theta\delta} & \sigma_{\theta\gamma} & \sigma_\theta^2 \end{bmatrix}\end{aligned}$$

Likewise the job heterogeneity terms, $\varepsilon_{ij}$and ψ_{ij}, are distributed as

$$\begin{aligned}(\varepsilon_{ij}, \psi_{ij}) &\sim N(0, \Sigma_{\varepsilon\psi}) \\ \Sigma_{\varepsilon\psi} &= \begin{bmatrix} \sigma_\varepsilon^2 & \sigma_{\varepsilon\psi} \\ \sigma_{\psi\varepsilon} & \sigma_\psi^2 \end{bmatrix}\end{aligned}$$

The time-varying health insurance and wage residuals, ξ_{ijt} and η_{ijt}, are normally distributed and are independent. The variance of ξ_{ijt} is not identified and so is normed to one.

$$\begin{aligned}\eta_{ijt} &\sim N(0, \sigma_\eta^2) \\ \xi_{ijt} &\sim N(0, 1)\end{aligned}$$

Given these distributional assumptions, the variances and correlations of the random effects can be estimated. Identification of the variances of the random person effects, δ_i, γ_i, and θ_i, is possible due to the presence of multiple jobs per person. Likewise the identification of variances of the random job effects, ε_{ij} and ψ_{ij}, is possible due to multiple wage observations for a given job and the presence of people who switch health insurance status within a job. Since job duration is observed only once per job, it is not possible to estimate the variance of a random job effect in the hazard equation. However the random job effects from the wage and health insurance equations can be included directly in the hazard model. Since the random effects have zero means, the coefficients on these effects, λ_1 and λ_2, will be estimates of the correlation between job specific hazard rates and propensities to have high wages and health insurance (Lillard (1999)). For instance a negative λ_1 implies that a job with a higher than average propensity to have health insurance will have a lower than average hazard of ending.

The model has a standard hierarchical structure with several levels of nested effects. To understand the estimation technique, consider first the likelihood function for estimating only equations 1 and 2.

$$\begin{gathered}P(\text{Observed HIEMP and Job History} \mid \beta_{HZ}, \beta_H, \sigma_\varepsilon^2, \Sigma_{\delta\gamma}) = \\ \int_\gamma \int_\delta \int_\varepsilon f(t_{ij} \mid \beta_{HZ}, \delta, \varepsilon) f(H_{ijt} \mid \beta_H, \gamma, \varepsilon) f(\varepsilon \mid \sigma_\varepsilon^2) f(\delta, \gamma \mid \Sigma_{\delta\gamma}) d\varepsilon d\delta d\gamma \qquad (4)\end{gathered}$$

The joint likelihood of the observed HIEMP and job tenure history is the product of a set of conditional probabilites, each representing a level of the heirarchy.

The first level is the probability of the observed job durations and employer-provided health insurance coverage, conditional on person and job heterogeneity. The next level is the probability of the job effect and the third level is the joint probability of the person effects. The coefficients on the observable characteristics β_{HZ} and β_H, are assumed to have point-mass prior densities which is equivalent to leaving the distribution of these effects out of the hierarchy and treating these as fixed unknown constants (Searle, Casella, McCulloch (1992)). The variance components σ_ε^2 and $\Sigma_{\delta\gamma}$ are also assumed to have flat, noninformative priors. The likelihood function for individual i with $j = 1$ to J_i jobs, each lasting $t = 1$ to T_{ij} time periods is

$$L_i = \int\limits_{\gamma}\int\limits_{\delta} f\left(\gamma, \delta \mid \Sigma_{\gamma\delta}\right) \times \tag{5}$$

$$\left\{ \begin{array}{c} \int\limits_{\varepsilon_{iJ}} \left\{ \prod\limits_{t=1}^{T_{iJ}} [1-\Phi(\frac{\beta_H X+\varepsilon_{iJ}+\gamma_i}{\sigma_\xi})]^{1-d_{iJt}} [\Phi(\frac{\beta_H X+\varepsilon_{ij}+\gamma_i}{\sigma_\xi})]^{d_{iJt}} \right\} \times \\ \left[h(t_{iJ} \mid \beta_{HZ})\right]^{D_{iJ}} S(t_{iJ} \mid \beta_{HZ}) \left[\frac{\phi(\varepsilon_{iJ})}{\sigma_{\epsilon_{iJ}}}\right] d\varepsilon_{iJ} \times \\ \prod\limits_{j=1}^{J_i-1} \int\limits_{\varepsilon_{ij}} \left\{ \prod\limits_{t=1}^{T_{ij}} [1-\Phi(\frac{\beta_H X+\varepsilon_{ij}+\gamma_i}{\sigma_\xi})]^{1-d_{ijt}} [\Phi(\frac{\beta_H X+\varepsilon_{ij}+\gamma_i}{\sigma_\xi})]^{d_{ijt}} \right\} \times \\ h(t_{ij} \mid \beta_{HZ}) S(t_{ij} \mid \beta_{HZ}) \left[\frac{\phi(\varepsilon_{ij})}{\sigma_{\epsilon_{ij}}}\right] d\varepsilon_{ij} \end{array} \right\} d\gamma d\delta$$

$$\begin{aligned} d_{ijt} &= 1 \text{ if individual i has HIEMP at job j at time t} \\ &= 0 \text{ otherwise} \\ D_{iJ} &= 1 \text{ if last job ends} \\ &= 0 \text{ if last job is censored} \\ \beta_H X &= \beta_{H1} X_i + \beta_{H2} X_{ij} + \beta_{H3} X_{ijt} \\ \beta_{HZ} X &= \alpha HIEMP_{ijt} + \beta_{HZ1} X_i + \beta_{HZ2} X_{ij} + \beta_{HZ3} X_{ijt} \end{aligned}$$

The parameters of the covariance-variance matrix, as well as the coefficients on the observable person and job characteristics for the hazard and health insurance equations, can be estimated by maximizing this likelihood function. The estimation is accomplished by first using numerical integration to integrate out ε, δ, and γ and then choosing β_{HZ}, β_H, σ^2_ε and $\Sigma_{\delta\gamma}$ to maximize the marginal likelihood function.

It is important to note that this likelihood is separable for individuals. All effects are at the level of the person or are nested within the person level. Thus instead of an employer effect, the likelihood function contains a job, or person-employer match, effect. This random job effect models the influence on a specific individual's outcomes of working for a specific employer. Thus by definition, workers who share an employer have different job effects[1].

The full estimation of equations 1, 2, and 3 adds the wage history to the likelihood. The probability of a worker's job, health insurance, and wage history is the product of the previous likelihood, now made conditional on wages, and the marginal probability of wages. This can be written as

$$P(\text{Observed Wage, HIEMP and Job History } \mid \beta_{HZ}, \beta_H, \beta_W, \sigma^2_\varepsilon, \Sigma_{\delta\gamma|W}, \sigma^2_\theta, \sigma^2_\psi, \sigma^2_\eta) =$$

[1]Although this may not be the most theoretically appealing approach, it is the only method that is currently feasible. Any effect shared by two individuals would complicate the integration to the point of making the computation intractable using currently available software.

$$\int_{\gamma}\int_{\delta}\left\{\int_{\varepsilon} f(t_{ij} \mid \beta_{HZ}, \delta, \varepsilon, W) f(H_{ijt} \mid \beta_H, \gamma, \varepsilon, W) f(\varepsilon \mid \sigma^2_{\varepsilon|W}, W) d\varepsilon\right\} \times \quad (6)$$

$$f(\delta, \gamma \mid \Sigma_{\delta\gamma|W}, W) d\delta d\gamma f(W \mid \beta_W, \sigma^2_{\theta}, \sigma^2_{\psi}, \sigma^2_{\eta})$$

The full likelihood becomes

$$L_i = (2\pi)^{-T_i/2} \left|\Sigma_{\theta+\psi+\eta,\theta+\psi+\eta}\right|^{-\frac{1}{2}} \exp\left\{-\frac{1}{2}\left(\overline{W}_i - \beta_w X\right) \Sigma^{-1}_{\theta+\psi+\eta,\theta+\psi+\eta} \left(\overline{W}_i - \beta_w \overline{X}\right)\right\}$$

$$\int_{\delta}\int_{\gamma} \phi(\gamma, \delta \mid \overline{W}_i) / \left[\Sigma_{\gamma\delta} \mid \overline{W}_i\right]^{-1}$$

$$\left\{\begin{array}{c} \int_{\varepsilon_{iJ}} \left[P(HIEMP_J)\left[h(t_{iJ} \mid \beta_{HZ})\right]^{D_{iJ}} S(t_{iJ} \mid \beta_{HZ})\right] \frac{\phi(\varepsilon_{iJ}|\overline{W}_i)}{\sigma_{\epsilon_{iJ}|\overline{W}_i}} d\varepsilon_{iJ} \times \\ \prod_{j=1}^{J_i-1} \int_{\varepsilon_{iJ}} \left[P\left(HIEMP_j\right) h(t_{ij} \mid \beta_{HZ}) S(t_{ij} \mid \beta_{HZ})\right] \frac{\phi(\varepsilon_{ij}|\overline{W}_i)}{\sigma_{\epsilon_{ij}|\overline{W}_i}} d\varepsilon_{ij} \end{array}\right\} d\gamma d\delta$$

(7)

$$\begin{aligned} \overline{W}_i &= \text{vector of wages over all months and jobs for person i} \\ \beta_w \overline{X} &= \beta_{W1}\overline{X}_i + \beta_{W2}\overline{X}_{ij} + \beta_{W3}\overline{X}_{ijt} \text{ vectors of X values} \\ P(HIEMP_j) &= \prod_{t=1}^{T_{ij}} [1 - \Phi(\frac{\beta_H X H_{ijt} + \varepsilon_{ij} + \gamma_i}{\sigma_\xi})]^{1-d_{ijt}} [\Phi(\frac{\beta_H X H_{ijt} + \varepsilon_{ij} + \gamma_i}{\sigma_\xi})]^{d_{ijt}} \\ D_{iJ} &= 1 \text{ if last job ends} \\ &= 0 \text{ if last job is censored} \end{aligned}$$

In estimating this more complicated likelihood function, the hazard and health insurance random effects, δ, γ, and ε are numerically integrated out as before, but probability distributions made conditional on wages are used. The resulting likelihood is then maximized directly by choosing the β's and variance components[2]. Section 4 describes the data used and Section 5 contains results.

4 Data Description

To estimate this model I use the 1990 and 1996 panels of the Survey of Income and Program Participation (SIPP). The SIPP is a longitudinal data set which interviewed respondents between eight (1990 panel) and twelve (1996 panel)

[2]Note that unlike in the case of the hazard and health insurance models, the random effects of the wage model, θ, ψ, and η, are not integrated our and $\overline{W_i} - \beta_W X$ is the quantity of interest. With sufficiently many numerical integration points, these two approaches are equivalent, but directly entering the random effects in the case of continuous outcome variables is computationally easier.

times and collected monthly data for the preceding four months. In the first interview (or wave) basic demographic information was collected as well as the date of labor force entry, years in the labor force, and start dates of on-going jobs. Then, during each four-month reference period, information about the number of children, marital status, job status, wages, hours, job characteristics, and health insurance information was collected. This interview pattern produced monthly data for 32 and 48 months respectively, with the panels ending in mid-1992 and early 2000. The use of two SIPP panels will be advantageous for two reasons. First, all previous studies have been done using data from the late 1980s and early 1990s. Using the 1990 SIPP will allow for comparisons with these earlier studies. Second, given the amount of change in the health insurance market during the 1990s and the multiple changes in laws regarding the portability of health insurance, the 1996 SIPP panel will allow the assessment of the importance of job-lock in the current labor market.

These Census surveys were both linked via Social Security Number to a confidential data set provided by the Social Security Administration (SSA) to the Census Bureau for the purpose of improving the Census SIPP product. This data set contained uncapped annual earnings broken down by employer over the time period 1978-2000. This information came from W-2 forms, shared by the Internal Revenue Service (IRS) with SSA for the purposes of administering SSA programs. Using this data, it is possible to tell how many employers a SIPP respondent had each year and how much the respondent earned from each employer. This data was crucial is constructing accurate job tenure histories, as will be discussed below.

Each of my three outcome variables involved a time series of responses. The wage series for each job was constructed using self-reported hourly wage rates when the respondent reported being paid by the hour and monthly earnings divided by weeks worked and usual weekly hours when he or she reported being paid a salary. No imputed wage values were used. Although earnings data were available from the Social Security administrative data, I chose not to use these values because they were annual and there was no means to allocate the earnings across months or create a wage rate, crucial information necessary in order to fully exploit the rich job tenure data collected in the SIPP[3].

Jobs reported in the SIPP were given longitudinal identifiers for the purpose of linking jobs for a given respondent across waves and calculating job tenure. Start dates for jobs in progress at the beginning of the survey were also collected to enable duration to be calculated for these jobs. However in the 1990 SIPP panel there was substantial miscoding of job identifiers. Between 30-40% of the jobs were coded such that jobs either falsely linked over time or erroneously failed to link. The later case affected approximately 10% of the jobs while the former affected 30% of the jobs. Thus the SIPP too often linked jobs that were in fact different, biasing job tenure upwards.

I attempted to solve this problem using two crucial sources of information:

[3]A separate paper compares the earnings measures collected by the SIPP to those reported in the SSA administrative data and finds high correlation between them. See Abowd and Stinson (2002).

name of the employer and SSA administrative counts of total number of employers. My process involved two steps. First, I used statistical name matching software to link jobs across interviews using the name of the employer, and to create a set of unique job identifiers for each individual. Using these new job identifiers, I then compared job totals for individuals over the course of the SIPP to job totals from the same time period in the SSA administrative data. This comparison allowed me to check the accuracy of the statistical name matches and find cases where misspellings of employer names or use of abbrevations had prevented the software from linking jobs. Since the SIPP allowed reports on only two jobs and did not cover all months of the year in the first and last years of each panel, I expected the job count for a given individual to be higher in the SSA data than in the revised SIPP data. For respondents where this was not the case, I flagged the job histories as likely to contain too many jobs and undertook additional editing, including a second pass with the matching software and hand-editing of remaining discrepencies. The resulting total count of all jobs held by SIPP respondents was 24% higher than the original total count.

In 1996, the Census Bureau instituted Computer Assisted Personal Interviews (CAPI) for SIPP data collection and as a result the job data in this panel were much cleaner. There was essentially no false linking over waves and a much lower incidence of false non-linking. Using a similar procedure to 1990, I reduced the overall job count by approximately 4% by linking jobs with similar names. Many of these non-links resulted from people missing interviews because of difficulties in tracking jobs across breaks in survey responses[4].

Health insurance information collected in the SIPP included questions about insurance status, and the source of insurance. Possible private sources were another family member, an employer, union, previous employer, the military, or some other private source. Government health insurance programs such as Medicaid or Medicare were coded separately. Reports of employer-provided health insurance were not associated with reports about a particular employer, and hence it was necessary to assign this benefit to a particular job when there were multiple jobs. To do this I assumed that those holding multiple jobs were likely to have a full and a part time job and to obtain their health insurance from their "main" employer, defined by where they worked the most hours or earned the most money. Thus, I selected a main employer based on hours, followed by earnings if hours were equal. Care was taken in this process to prevent small fluctuations in hours from causing frequent changes in the health insurance source.

To make my sample representative and comparable to other research, I restricted myself to original sample members who held at least one job with nonzero earnings and who were not active duty members of the military. Jobs

[4]The revision of the longitudinal job identifiers was accomplished through the use of non-publicly available SIPP data, such as the name of the employer, and the use of SSA administrative data, also not publicly available. Hence these accurate tenure calculations have also been unavailable to non-Census researchers up to this point. However, the LEHD program at the Census Bureau is preparing to release these revised job identifiers in May 2003 in order to aid those studying tenure issues using the SIPP.

for these people were included when the job was begun at age 18 or older and ended or was censored by age 60; job duration was greater than one day; the employer was not a family business; an hourly wage of at least $.1 was reported at least once. Weighted summary statistics for the 1996 and 1990 Panels are presented in Tables 1, 2, 3, and 4 respectively. My sample is fairly standard. SIPP respondents in the 1996 panel were 84% white, 49% male, and 62% were married in March 1996. On average, respondents held 1.98 jobs over the time period of the panel and 60% of people had employer-provided health insurance. Jobs lasted an average of four years, although the SIPP contains information about only 16 months of a job on average. This is due to the fact that many jobs were on-going when the SIPP began and many are missing wage information for several months during the panel due to non-response. The average job paid a real hourly wage (1999 dollars) of $15.01 in March 1996 and was held by a worker who worked 39 hours per week and who began the job at age 32 with 13 years of experience. The 1990 SIPP panel is quite similar to the 1996 panel with the only major differences being that 63% of people have employer provided health insurance, the average number of jobs held is 1.7, and the average hourly wage is $14.26.

Since variation in employer provided health insurance within job is an important component of my model, I investigated the causes of these types of changes in order to correct potentially false changes. Originally 17% of jobs contained at least one switch in employer provided health insurance status. These changes were compared to changes in other variables to determine whether the variation represented real coverage changes or survey response error. For instance, the SIPP allowed proxy respondents, and hence in some waves a person responded for him- or herself while in other waves, a family member responded for him or her. In many instances where a health insurance switch took place, there was also a change in the person responding. In these cases, I recoded the health insurance variables to always contain the values reported in the waves without proxies.

Another difficulty occurred when the health insurance change took place in either the first or last three months of the job. While employers commonly impose a waiting period for insurance at the beginning of a job, this kind of a change is unlikely to be truly exogenous variation. The worker began the job with the expectation that he or she would quickly obtain insurance through the employer and perceived this to be part of the compensation. Including this kind of a change might bias the effect of health insurance on job duration because jobs which did not last longer than three months were less likely to have health insurance regardless of whether it would have been eventually offered. Likewise, the loss of health insurance in the last three months of a job was likely the result of the worker's decision to leave the job and hence either a preliminary switch to another kind of health insurance or a difficulty on the part of the SIPP survey instrument in accurately coding the month of the health insurance change. Again, inclusion of this type of change would have biased the effect of health insurance. In this case it was the decision to separate which drove the health insurance change and not vice versa. As a result of these concerns, workers who gained health insurance from an employer during the first three months of a job or lost it during the last three months were coded as having employer-provided health insurance during the entire duration of the job. After recoding false health insurance changes, only 12.57% of jobs contain at least one switch in employer provided health insurance.

The major remaining concern was whether changes in health insurance status within jobs were exogenous. Since insurance changes are possibly correlated with "life events" such as marriage, divorce, addition of children to a family, or spousal employment changes, the hazard of separating may be changed in a way unrelated to, but indistinguishable from, the direct effect of the health insurance change. Figures 1 and 2 investigate this problem by categorizing changes in employer-provided health insurance within jobs according to whether a worker gained or lost this insurance and concurrently, whether they gained or lost health insurance coverage of any other kind. As previously described for the 1996 panel and shown in Figure 1, workers changed the status of their employer-provided health insurance during 12.57% of jobs. These changes were almost equally divided between those who gained and lost employer-provided health insurance. Among employer-provided health insurance gainers, 36% previously had health insurance coverage through another person, 22% had a private policy in their own name, and 44% were not insured at all[5]. Of those who had health insurance through another person, 70% kept this insurance after gaining coverage from their employers, while 30% dropped it. This is perhaps a surprising result and one of the major differences between panels. In 1990, of those who had been previously insured by another person, 79% dropped this insurance when they gained insurance from their employer (Figure 2). This result is consistent with the overall rise in dual coverage between 1990 (3.4%) and 1996 (5.1%).

The majority of those in the 1996 panel who lost health insurance through their employer, remained insured, either through another person (37%), through another type of policy in their own name (43%), or both (2% overlap). This again differs from 1990 where fewer people gained private health insurance in their own name and larger numbers lost health insurance coverage completely. In 1996 of those who were prevented from becoming uninsured by having a policy through another person, 47% had held this coverage previous to losing health insurance from their employer. This compares to only 25% who previously held coverage through another person in 1990 (Figure 2).

Of the types of changes described above, those most likely to have been exogenous changes in health insurance are those who changed general insurance status (insured to uninsured or vice versa) and those who dropped or gained private policies in their own names. In each of these cases there do not appear to have been switches in spousal provision of health insurance, and hence these changes are more likely to have resulted from employer decisions regarding provision of health insurance benefits. It is possible that the worker changed his or her behavior in a way which changed his or her eligibility, by changing hours

[5]These percentages sum to more than 100% because some people had two sources of insurance.

worked per week, for example. However for those who either became insured or dropped private coverage, only between 1% and 3% increased their hours by enough to move them from full-time to part-time. Those losing coverage completely or gaining private insurance decreased their hours by enough to move to part-time work between 4% and 6% of the time.

Changes in employer-provided health insurance during the tenure of a job that are also associated with changes in health insurance through another person are less likely to be exogenous. Workers who gain employer-provided health insurance and retain coverage through another person may have chosen to double insure, due to some change in family medical conditions, the addition of children, or simply a change from part to full-time status which caused the offer of health insurance at the job. However the changes that can be measured appear to be small. Only 1.25% gained children and 2% gained hours consistent with moving to full-time work. Those who gained employer-provided health insurance and gave up coverage through another person are perhaps most likely to represent endogenous changes. These workers could have experienced divorce, spousal job loss, changes in the number of children in the family, or changes in hours worked. In fact 3.24% of these workers had spouses who ended jobs, 2.2% of them had a change in their number of children, .7% of them divorced or separated, and 3% of them gained hours consistent with moving to full-time status. All these changes together still affect less than 10% of the workers in this category.

Of those who lost employer-provided health insurance, between 4% and 6% lost hours consistent with moving from full to part-time work. Of those who switched from insuring themselves to having insurance through another person, 3% of these married, 9% had a spouse gain a job, and 1.6% gained a child in the family. Of those who were dual insurance holders and then dropped their own coverage, 1.4% gained a child in the family and 4% lost hours.

Taken together these changes provide some evidence that within job changes in employer-provided health insurance are not always correlated with "life-events" and so may be useful in identifying job heterogeneity. A worker who began a job without health insurance, possibly because it was only offered after an extensive waiting period, or because the worker had other insurance with which he or she was satisfied, or because the firm did not offer the benefit, and then after some time period switches to employer-provided health insurance, will have an underlying job-specific probability of having health insurance, identified by the months during the job with and without insurance. The inclusion of this job-specific health insurance effect in the hazard equation will account for correlation between the propensity for a given job to end and the quality of that job, i.e. health insurance provision. The coefficient on the main health insurance indicator in the hazard model can then be interpreted as the change in the conditional probability of the job ending, holding all else constant, when the worker begins or ends employer-provided health insurance coverage.

5 Results

5.1 Results for Joint Health Insurance and Hazard Specifications

I begin by estimating the hazard equation (equ. 2) alone and then jointly with the health insurance equation (equ.1). The health insurance probit model is estimated with controls for race, gender, ethnicity, education level, existence of chronic health condition, number of kids, marital status, kids and marital status interacted with gender, indicator for weekly hours greater than 20, age at beginning of job, union status, industry and job type. The hazard equation is estimated with piecewise linear splines in age, calendar time, labor force experience and job tenure as well as controls for race, ethnicity, education level, gender, marital status, number of kids, interaction of marital status and number of kids with gender, own health insurance from non-employer source, union status, industry, and job type.[6] The primary explanatory variables of interest in the hazard equation are health insurance from the employer, health insurance from another person, and both. All three of these variables have been interacted with gender and marital status. The reported values for the health insurance variables are hazard ratios which are exponentiated hazard model coefficients. These ratios represent the probability of a worker with a given type of health insurance, X, leaving a job relative to the probability for an otherwise identical worker whose characteristics become the baseline and are summarized by βt. Mathematically, these ratios are described by

$$\frac{\exp\left(\beta t + \alpha X\right)}{\exp(\beta t)} = \frac{\exp(\alpha X)}{1}$$

Ratios less than one result when the type of insurance, X, has a negative effect on the hazard of ending a job while ratios greater than one result when X has a positive effect.

Hazard ratios for the health insurance variables are presented by gender/marital status group in Tables 5 and 6. . In Table 5 the baseline person is a married male with no health insurance. Column 1 presents results from estimating equation 2 alone with no random effects. Having health insurance from one's employer produces a hazard that is 34.7% that of the hazard for the baseline person, a 65.3% reduction. By comparison, having both employer-provided health insurance and insurance through another person gives a hazard that is only 48.4% of the baseline hazard, a 51.6% reduction. Thus the percentage increase in the hazard caused by moving from only employer-provided insurance to dual coverage is 28.9% [BOTH-HIEMPLOYER)/Both]. Those workers whose sole source of health insurance is another person are also less likely to end their jobs relative to an uninsured worker. The hazard ratio represents

[6] The excluded case is a worker who is non-white, female, single, age 32 years old, with no kids, no high school degree, 14 years of labor market experience, and job tenure of 0 months in March of 1996 at a non-union, private sector, wholesale trade job.

a 28% decrease from the baseline hazard. The percentage change in the hazard caused by the shift from non-insured to insured through another person is -39.1% [OTHER-1/OTHER]. The difference-in-difference effect reported in Table 5 is 68% [28.9%-(-39.1%)]. The effect takes account of the fact that alone, other insurance is associated with reduced mobility, and so the total effect of dual insurance both overcomes this negative effect and additionally further increases the hazard.

This difference-in-difference estimate is similar to methods used by Madrian and others in the literature but is not directly comparable because it represents a proportional shift in the hazard rate (i.e. a change in the conditional probability of ending the job at any point in time), instead of a change in the probability of leaving after a set amount of time. The large magnitude of this effect is a result of the large negative effect of insurance through another person on those for whom this is their sole source of insurance (HIOTHER). The literature also find a similar negative effect. The concern arises that having health insurance through another person, most commonly a spouse, signifies a certain type of person. These workers have better jobs and lower individual propensities to leave these jobs and so are not directly comparable to uninsured workers. Thus the -39.1% difference between the uninsured and those insured through other people is overstated because the comparison groups differ along dimensions other than the treatment effect.

Columns 2-4 take account of this critique by adding heterogeneity terms and treating the health insurance variable as endogenous. Column 2 shows results from the same specification as in Column 1 but with the addition of a random person effect in the hazard model. Column 3 presents estimates of equations 1 and 2 with random person effects correlated across equations and a random job effect included in equation 1. Column 4 lists results when the random job effect from equation 1 is included in equation 2. The joint estimation strategy and inclusion of random effects have the expected effect on the hazard ratio for employer-provided health insurance. The inclusion of heterogeneity but the failure to account for correlation between δ_i and HIEMP biases the hazard ratio even further downward in Column 2. When the joint health insurance, hazard model is estimated in Column 3, the hazard rises, a trend that continues in Column 4. This result lends support to the idea that taking account of individual and job propensities to have health insurance and be mobile reduces the estimated direct effect of health insurance because downward bias is removed. The reported standard deviations and correlation coefficients provide direct evidence of the negative correlation between employer-provided health insurance and hazard rates. The correlation coefficient, $\rho_{\delta\gamma}$, is negative and significantly different from zero. This correlation is interpreted as predicting that workers with high individual propensities to have health insurance have low individual propensities to have jobs end. The coefficient from the hazard model on the health insurance random job effect, λ_1, is interpreted as the correlation between the job specific propensities for separation and holding health insurance. It is also negative although the correlation is not as strong as with the individual heterogeneity. Jobs which are more likely to provide health insurance have a lower hazard of ending.

In spite of the rise in the hazard ratio for employer provided health insurance, the difference-in-difference estimate remains fairly large and falls only very slightly in Columns 3 and 4. This is due to the fact that HIOTHER remains a significant negative predictor of mobility and the interaction term between HIOTHER and HIEMP remains fairly constant and positive. The inclusion of person and job heterogeneity does not change the effect of these variables and hence significant differences in hazard rates remain between those with only HIEMP and those with dual insurance coverage.

Table 6 presents hazard ratios for three other demographic groups: single men, married women, and single women. In all cases the hazard ratios follow the same pattern as for married men. The hazard for employer-provided health insurance rises in Columns 3 and 4 while the hazard ratio for health insurance through another person remains fairly constant. The hazard ratio for dual coverage moves in tandem with the ratio for HIEMP, indicating a positive and constant interaction term. The difference-in-difference effects remain positive. The level of these effects, however, differs substantially depending on the group. Single men see a much smaller rise in the hazard of a job ending due to the presence of a second source of coverage (16%). Married women have the largest levels of job lock. An additional source of health insurance increases their hazard by over 80%. Single women have lower levels than married workers but higher levels than single men.

Figure 3 shows relative hazard rates for four different groups of workers beginning jobs in March 1996 over the first 48 months of a job. The base person is a white, non-Hispanic, married man, age 32 years old with one child, a high school education and 14 years of labor force experience, beginning a non-union, private sector job in the wholesale trade industry. The hazard over time is generated using the tenure spline coefficients. The probability of quitting at time t conditional on the job having lasted until $t-1$ peaks at three months and then tapers off sharply until six months, when it levels out and decreases more gradually. The hazard for those workers who have employer-provided health insurance but are otherwise identical to the base person is substantially lower. The hazards for those with dual insurance and those with only insurance from another person lie in between the base and HIEMP workers. This figure gives a graphical presentation of the difference-in-difference result and shows the dramatic impact of health insurance benefits on the probability of a job ending.

Table 7 presents results from the same specifications as in Table 5 but for the 1990 SIPP Panel. Due to the introduction of new portability laws in July 1997 (HIPPAA) which restricted employers' ability to deny coverage because of pre-existing conditions, one would expect job-lock to be substantially higher in 1990. In fact this is true only for single men and women. The difference-in-difference estimates of job-lock are 52% and 65%, respectively, compared to 16.5% and 36% in 1996. For married men and women, the difference-in-difference estimate is very similar across panels, 61% and 86% in 1990 compared to 67% and 86% in 1996. While this model is not meant to be an explicit test of HIPPAA, it does provide some preliminary evidence that the effect of the new law has either been somewhat limited or slow to have an impact on mobility.

The results from Tables 5, 6 and 7 seem to point to significant interactions between family status, spousal employment status, and the effect of various types of health insurance. Some of these results may be influenced by heterogeneity in the types of employer-provided health insurance held by workers with and without alternate coverage sources. I consequently estimate two specification checks for whether the type of employer-provided insurance has an effect on the level of job-lock. In table 8 I divide workers with employer-provided health insurance into three categories: those who cover only themselves, those who cover one other family member, and those who cover multiple family members. Using this definition of health insurance benefits, those who cover only themselves experience a much larger increase in mobility when they have a second source of health insurance. The difference-in-difference estimate of job-lock is 62% for insure-self-only workers compared to only 34% for insure-multiple-family-members workers. However it is possible that this merely reflects the fact that workers with families are less mobile due to family reasons and hence having an optional source of health insurance does not increase the hazard by as much.

Table 9 investigates the issue of whether the amount of compensation associated with the health insurance benefit has any influence on the level of job lock. Here the results are quite striking. For those workers who pay none of the cost associated with their employer-provided health insurance, the addition of another source of coverage will increase mobility by 29%. For those who pay part of the cost of HIEMP, dual coverage increases their mobility (raises the hazard) by 41%. However for those who pay all of the costs, having additional coverage lowers the hazard relative to those who only have HIEMP, indicating that these workers were not job-locked.

In the case of a health insurance benefit that requires full-premium payment, the compensation component of the health insurance effect is greatly reduced. The financial component now includes only the opportunity to buy group health insurance that allows risk pooling and to have premuims deducted from pre-tax wages. This is reflected in the difference between the HIEMP hazard for the "pays all costs" group and the "pays no cost group." Workers who pay all their premiums experience only a 19% direct reduction in their hazard rate compared to a 56% reduction for those who pay nothing (Table 9, column 2). The difference between the hazard ratios of these two groups (37%) indicates that the compensation component accounts for at least 66% of the total decrease in the hazard ratio when the worker has a paid-in-full health insurance benefit.

5.2 Results for Joint Wage and Hazard Specifications

Table 10 presents results from the joint estimation of the hazard and wage equations. Coefficients from the tenure and general labor market experience splines are reported, as well as the coefficients for education categories. In the first column, the coefficients are from a simple linear wage regression. In the second column, random person and job effects are added to the wage regression. In the third column, the wage and hazard models are jointly estimated and correlation across the equations in the person heterogeneity components is allowed. In the final column, wage job heterogeneity is included in the hazard equation and a correlation coefficient on this effect is estimated. Comparing the tenure coefficients across these four specifications shows the effects of controlling for heterogeneity and endogeneity bias. Adding heterogeneity substantially decreases the return to seniority during the first year of the job. In column 1, an additional month of tenure after three to six months on the job implies a .006% increase in log wages, an annual rate of 7%. The effect during the sixth to twelfth months is similar. However by adding heterogeneity terms, these effects fall significantly to an annual rate of approximately 3.5% over the course of the third to twelfth months. In column 1, returns after the first year are substantially lower and represent annual rates of return to seniority of 1%-2%. These rates are also more stable across specifications. Perhaps most surprising is that there are few differences in the tenure spline coefficients between columns 2, 3 and 4. Making tenure endogenous does not seem to have a significant impact on the tenure coefficients.

Hazard ratios from the job duration part of the joint estimation in columns 3 and 4 are not reported but are fairly similar to the results in columns 3 and 4 of Table 5. One of the most informative parts of this joint model is the relationship between the random effects in each equation. The standard deviations of the random effects are reported in Table 10 as well as the parameters of the cross-equation relationships. The positive correlation between the person effects in the wage and hazard models ($\rho_{\delta\theta} > 0$), implies that high-wage workers are also likely to be "movers" in that they have high individual propensities to have jobs end. On the other hand, $\lambda_2 < 0$ implies that jobs with high wages ($\psi_{ij} > 0$) are also jobs which are less likely to end, i.e. have lower hazard rates. These results are consistent with search models where there are returns to changing jobs because match quality rises and workers who are observed to engage in search benefit in the form of higher wages. However, once a worker has found a good match in the form of a well-paying job, he or she becomes less mobile.

5.3 Results for Joint Wage, Hazard, and Health Insurance model

Results for the full three equation model are reported in Table 11 for both the 1996 and 1990 SIPP Panels. Standard deviations of all the random effects are reported as well as correlation coefficients, health insurance hazard ratios from the tenure equation, and tenure coefficients from the wage model. This specification allows the random effects in all three equations to be correlated and provides measures of seniority and health insurance effects which take this correlation into account. The signs of the correlation coefficients are again instructive about the matching process and the estimated biases likely to arise if heterogeneity is not explicitly modeled. Job heterogeneity from both the wage and health insurance equations is associated with lower probabilities of a job ending ($\lambda_1, \lambda_2 < 0$). However the correlation is much higher for wages than for health insurance. This result implies that highly compensated worker-job matches are less likely to end, possibly either because the match is very productive for both sides or because the opportunity cost of leaving the job is high for the worker. On the person heterogeneity side, the correlations between wages and mobility and health insurance and mobility have opposite signs. Workers likely to have health insurance are also less likely to have jobs end ($\rho_{\delta\gamma} < 0$), while high wage workers tend to be more mobile ($\rho_{\delta\theta} > 0$ in 1996 panel although not in 1990). Individual and job propensities for high wages also tend to be associated with individual and job propensities for health insurance ($\rho_{\theta\gamma}, \rho_{\psi\varepsilon} > 0$).

The result of controlling for these correlations is consistent with the hypothesis that heterogeneity biases seniority coefficients in wage models and health insurance coefficients in tenure models. Tenure effects average 4% during the first year and fall to approximately 2% after four years. These results are very similar to those shown in column 4 of Table 10, indicating that the inclusion of the health insurance equation did not substantially alter the results from the wage equation. On the other hand, the hazard ratio for employer-provided health insurance rises to approximately 60%, substantially higher than in Table 5, indicating that adding the wage equation had a large impact on the hazard model. In this model, the magnitude of job lock can be estimated in two ways. First, one can use the standard difference-in-difference technique discussed previously and compare those with dual insurance coverage to those with health insurance coverage only from an employer and then subtract the independent effect of insurance coverage through another person. This difference-in-difference estimator (last row of Table 11) predicts a 55%-61% drop in the hazard of a job ending when the employee holds work-related health insurance. This effect remains large because other insurance continues to reduce mobility when it is the only type of insurance held but increase mobility when it is a second source of insurance. The difference-in-difference estimate is lower in 1996 than in 1990, but only by 6%.

Another alternative to using the difference-in-difference method is to consider the implications of the random person and job effects. For example, consider a worker who earns \$16.50 per hour on average during the SIPP panel, a pay rate which is \$1.50 (.1 log wages) more per hour than the March 1996 sample average of \$15. For simplicity, assume the worker held only one job and was observationally equivalent to the "average" worker earning \$15 per hour. Using the variance estimates, σ_θ, σ_ψ, and σ_η, one can predict values for θ_i and ψ_{ij} which estimate how much of this "excess" wage is due to unobservable person and job characteristics[7]. In this example, $\theta_i = .0445$ log wages ($\approx$ \$.67) and $\psi_{ij} = .0549$ log wages ($\approx$ \$.82). The remaining amount is due to random time variation. Given this realized value for ψ_{ij}, one can predict an effect on the hazard of the job ending using λ_2.

[7] Searle, Casella, and McCulloch (1992) derive a Best Linear Unbaised Predictor (BLUP) for a linear model with a normally distributed, mean zero random effect as $BLUP(u) = \frac{cov(u,y)}{var(y)}\left(y - \mu_y\right)$.

$$\text{Hazard ratio} = \exp\{\lambda_2 * \psi_{ij}\} = .977$$

Thus a 5% increase in job-specific wages, results in a 2.3% reduction in the hazard.

A positive value of ε_{ij}, or an increase in the probability of having employer-provided health insurance due to unobservable job characteristics, similarly reduces the hazard of the job ending. This effect works through λ_1, the hazard coefficient on ε_{ij}, and through α, the hazard coefficient on employer-provided health insurance. Suppose in this example that $\varepsilon_{ij} = .085$ and that the individual had an initial probability of having health insurance of 60%. Due to a positive ε_{ij}, this probability increases to 63.24%. The hazard ratio would be

$$\begin{aligned}\text{Hazard Ratio} &= \exp\{\lambda_1 * \varepsilon_{ij} + \alpha * (\triangle\text{Probability of HIEMP})\} \\ &= \exp\{-.0894 * .085 + \ln(.612) * (.0324)\} = .977\end{aligned}$$

This represents a 2.3% reduction in the hazard rate. Thus in this example, ε_{ij} and ψ_{ij} both produce approximately the same effect on the conditional probability of a job ending. Increasing the job-specific probability of health insurance from 60% to 63.24% or raising wages at a job by 5% both reduce hazard by slightly more than 2%.

This effect of increasing the probability of health insurance depends on the reference point, i.e. the assumed initial probability of having employer-provided insurance. If the individual had been assumed to have a 20% probability of receiving employer-provided health insurance, this probability would only need to be raised to 22.95% to induce a 2.3% reduction in the hazard. Thus the less likely the employee is to have HIEMP initially, the less the probability must rise in order to achieve the same effect as the 5% wage premium.

Although no information is available in the SIPP about the cost to employers of providing health insurance policies, national averages provide some insight into the monetary value of health insurance benefits. Branscome and Brown (2001), using data from the 1998 Medical Expenditure Panel Survey, report that the average health insurance premium for an employer-provided single person coverage policy is \$2,174, of which the employer pays on average 82.4% (\$1,791). Thus a job which predicts a 3.24% increase in the probability of having single person employer-provided health insurance coverage also predicts a \$58 gain in expected compensation. Family coverage costs an employer \$4,208 on average and a 3.24% rise in the probability of having family coverage represents a \$136 expected gain. By comparison, a worker with a 5% job log wage premium compared to the sample average wage of \$15, earns an additional \$1,604 per year. Assuming a 45% total tax rate, this is equivalent to a real annual earnings increase of \$882 for the worker. Since health insurance benefits are not taxed, the expected gain in compensation associated with an increase in the probability of health insurance is all real increase. Thus reducing the hazard of a worker quitting by 2.3% by increasing the probability of providing that worker family health insurance benefits will cost a firm on average 16% of what it would cost to obtain the same reduction in mobility by increasing wages. This result is again consistent with the existence of job lock because the total compensation package if worth more to an employee than it costs the employer to provide it. Hence an alternate employer could not lure the employee to a new job by promising slightly higher compensation but all in the form of wages. Workers who have reason to believe they will be denied health insurance coverage at a new employer will not accept the offer even if the job pays higher wages.

6 Conclusions

My joint estimation of job duration, health insurance status, and wages explicitly takes into account many of the concerns arising in the estimation of the effect of employer-provided health insurance on the probability of a job ending and of the effect of tenure on wages. By controlling for person and job heterogeneity and allowing this heterogeneity to be correlated across equations, I am able to account for unobservable person and job characteristics which might otherwise bias the coefficients on tenure and health insurance. In spite of these controls, all of my specifications show some level of job lock (ranging from 30%-60%) using data from both the 1990 and 1996 SIPP panels. Controlling for person and job heterogeneity mitigates the direct negative impact of employer-provided health insurance but does not substantially change the effect of insurance from another source. Insurance through another person decreases mobility when it is held alone and increases mobility when it is an additional source. There is some evidence that the effect of employer-provided health insurance depends upon the type of insurance. For example, workers who pay all of their health insurance premiums without any employer contributions experience essentially no job-lock.

In the wage model, tenure is determined to be a significant predictor of wages. My estimates of 3.6% wage growth per year of tenure during the second through fourth years of a job and 1.5%-2% per year thereafter are slightly higher than the low estimates of Abraham and Farber and Altonji and Shakotko but lower than those of Topel. Heterogeneity seems to be the main cause for biased returns to seniority as the joint estimation of the tenure and wage equations did not reduce the tenure coefficients substantially more than had been caused by the inclusion of person and job heterogeneity.

The correlation coefficients from the joint estimation of the full three equation model provide support for the idea that wages, health benefits, and job tenure are three jointly determined outcomes. A job which pays a 5% log wage premium is 2.3% less likely to end all else equal. In comparison, a job which has a 63.25% probability of providing health insurance compared to an observationally equivalent job with a 60% probability, is also approximately 2.3% less likely to end. The monetary costs to the employer of these equivalent reductions in mobility are very different. The increase in the probability of family coverage health insurance represents a $136 gain in expected annual compensation while the 5% log wage premium represents a $882 increase in take-home pay. Health insurance appears to be worth more to workers than the equivalent amount of wages. These results point to the need for further studies of the wage/benefit trade-off using more detailed microdata on the cost of health insurance to the worker and the firm. Given the possibility that the health insurance market in the United States may move towards an individual plan-based system where firms subsidize but do not provide health insurance, employers will increasingly consider how much additional direct compensation is necessary in order to make a worker willing to give up health benefits and still remain relatively stable.

The relationship between the composition of benefit packages and tenure is also important when studying firms and the choices they make about what kind of workers to employ. Some firms may choose to pay average wages, provide no benefits, and experience high turnover rates. This type of business strategy would avoid the costs associated with job-lock. However, other kinds of firms may desire a more stable workforce and hence will have a cost incentive to increase the benefits/wages ratio. In making this decision, the firm will have to balance the efficiency loss due to job-lock with the gains from stability. By studying firm outcomes such as productivity per worker and sales per worker and how these relate to compensation packages and turnover rates, one could assess the efficiency gains and losses associated with high or low turnover and study how firms make these choices in order to maximize profits.

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Table 1: Summary Statistics for Individuals 1996 SIPP Panel*

Time-Invariant Person Char				Time Varying Person Char: Workers in 1996:3			
Variable	N	Mean	Std. Dev.	Variable	N	Mean	Std. Dev.
white (%)	35,060	84.31		married (%)	24,639	61.72	
years of education	35,060	13.03	2.701	kids	24,639	0.98	1.160
no high school degree (%)	35,060	15.85		kids < 18 years old	24,639	0.82	1.092
high school degree (%)	35,060	29.37		HI coverage (%)	24,639	84.69	
some college (%)	35,060	31.35		HI own name(%)	24,639	65.72	
college degree (%)	35,060	16.09		HI employer (%)	24,639	60.25	
graduate degree (%)	35,060	7.33		HI military (%)	24,639	0.47	
male (%)	35,060	49.31		HI previous employer (%)	24,639	1.56	
Hispanic (%)	35,060	9.86		HI private (%)	24,639	3.44	
number of jobs	35,060	1.98	1.324	HI other person's name (%)	24,639	24.06	
chronic,non-severe health condition (%)	35,060	14.12		HI both own and other (%)	24,639	5.08	
				Time Varying Person Char: Workers in 1999:3			
				Variable	N	Mean	Std. Dev.
				married (%)	19,937	62.41	
				kids	19,937	1.04	1.192
				kids < 18 years old	19,937	0.84	1.109
				HI coverage (%)	19,937	87.26	
				HI own name(%)	19,937	67.16	
				HI employer (%)	19,937	62.17	
				HI military (%)	19,937	0.52	
				HI previous employer (%)	19,937	1.55	
				HI private (%)	19,937	2.92	
				HI other person's name (%)	19,937	26.26	
				HI both own and other (%)	19,937	6.17	

*Includes original sample members of the 1996 SIPP Panel who were not active duty military at any point during the panel. In addition respondents must have held at least one job that 1) began after age 18, 2) either ended or was censored before age 60, 3) was on-going at some point during the period 1995:12 and 2000:2, 4) had total earnings > 0, 5) had at least one month with a non-imputed hourly wage >= $.1, 6) had job duration > 1 day, 7) a non-family employer. All statistics are weighted.

Table 2: Summary Statistics for Jobs 1996 SIPP Panel

Time-Invariant Job Characteristics				
Variable		N	Mean	Std. Dev.
Job Duration (months)		69,331	48.01	75.439
Number of months observed		69,331	15.74	14.816
Job Censored (%)		69,331	41.40	
Labor Force Exp begin job (months)		69,331	155.72	117.414
Age begin job (years)		69,331	31.09	10.010
Union/Covered by Union (%)		69,331	14.22	
Industry	Agriculture (%)	69,331	1.83	
	Mining (%)	69,331	0.42	
	Construction (%)	69,331	5.84	
	Manufacturing (%)	69,331	14.31	
	Transport., Comm.,Public Util. (%)	69,331	6.09	
	Wholesale Trade (%)	69,331	3.58	
	Retail Trade (%)	69,331	20.12	
	FIRE (%)	69,331	5.43	
	Services (%)	69,331	37.97	
	Public Admin (%)	69,331	4.41	
Job type	Private, For-Profit (%)	69,331	79.77	
	Private, Not-for-Profit (%)	69,331	6.96	
	Local Govt (%)	69,331	6.67	
	State Govt (%)	69,331	4.40	
	Federal Govt (%)	69,331	2.21	

Time-Varying Job Characteristics in 1996:3			
Variable	N	Mean	Std. Dev.
real hourly wage ($ 1999)	25,488	15.01	17.316
hours per week	25,488	39.36	12.288
Employer-provided HI (%)	25,488	58.82	

Time-Varying Job Characteristics in 1999:3			
Variable	N	Mean	Std. Dev.
real hourly wage ($ 1999)	20,980	14.34	19.993
hours per week	20,980	38.49	12.190
Employer-provided HI (%)	20,980	59.33	

Table 3: Summary Statistics for Individuals 1990 SIPP Panel*

Time-Invariant Person Char				Time Varying Person Char: Workers in 1990:3			
Variable	N	Mean	Std. Dev.	Variable	N	Mean	Std. Dev.
white (%)	18,634	85.27		married (%)	14,644	61.64	
years of education	18,634	13.06	2.699	kids	14,644	0.99	1.180
no high school degree (%)	18,634	16.57		kids < 18 years old	14,644	0.81	1.091
high school degree (%)	18,634	34.51		HI coverage (%)	14,644	87.23	
some college (%)	18,634	27.57		HI own name(%)	14,644	69.53	
college degree (%)	18,634	10.40		HI employer (%)	14,644	64.40	
graduate degree (%)	18,634	10.95		HI military (%)	14,644	0.50	
male (%)	18,634	50.02		HI previous employer (%)	14,644	1.13	
Hispanic (%)	18,634	7.68		HI private (%)	14,644	3.50	
number of jobs	18,634	1.78	1.133	HI other person's name (%)	14,644	21.21	
chronic,non-severe health condition (%)	18,634	8.65		HI both own and other (%)	14,644	3.51	
				Time Varying Person Char: Workers in 1992:3			
				Variable	N	Mean	Std. Dev.
				married (%)	14,661	62.01	
				kids	14,661	1.00	1.179
				kids < 18 years old	14,661	0.80	1.097
				HI coverage (%)	14,661	87.27	
				HI own name(%)	14,661	69.51	
				HI employer (%)	14,661	64.68	
				HI military (%)	14,661	0.58	
				HI previous employer (%)	14,661	1.11	
				HI private (%)	14,661	3.14	
				HI other person's name (%)	14,661	21.51	
				HI both own and other (%)	14,661	3.74	

*Includes original sample members of the 1990 SIPP Panel who were not active duty military at any point during the panel. In addition respondents must have held at least one job that 1) began after age 18, 2) either ended or was censored before age 60, 3) was on-going at some point during the period 1989:10 and 1992:8, 4) had total earnings > 0, 5) had at least one month with a non-imputed hourly wage >= $.1, 6) had job duration > 1 day, 7) a non-family employer. All statistics are weighted.

Table 4: Summary Statistics for Jobs 1990 SIPP Panel

Time-Invariant Job Characteristics				
Variable		N	Mean	Std. Dev.
Job Duration (months)		32,623	48.33	74.336
Number of months observed		32,623	14.74	11.550
Job Censored (%)		32,623	50.52	
Labor Force Exp begin job (months)		32,623	126.55	115.302
Age begin job (years)		32,623	30.59	9.974
Union/Covered by Union (%)		32,623	17.02	
Industry	Agriculture (%)	32,623	1.95	
	Mining (%)	32,623	0.47	
	Construction (%)	32,623	5.89	
	Manufacturing (%)	32,623	16.56	
	Transport., Comm.,Public Util. (%)	32,623	6.75	
	Wholesale Trade (%)	32,623	3.71	
	Retail Trade (%)	32,623	19.31	
	FIRE (%)	32,623	5.53	
	Services (%)	32,623	35.24	
	Public Admin (%)	32,623	4.58	
Job type	Private, For-Profit (%)	32,623	80.69	
	Private, Not-for-Profit (%)	32,623	5.13	
	Local Govt (%)	32,623	2.69	
	State Govt (%)	32,623	4.51	
	Federal Govt (%)	32,623	6.98	

Time-Varying Job Characteristics in 1990:3			
Variable	N	Mean	Std. Dev.
real hourly wage ($ 1999)	15,288	14.26	10.875
hours per week	15,288	38.55	11.624
Employer-provided HI (%)	15,288	62.29	

Time-Varying Job Characteristics in 1992:3			
Variable	N	Mean	Std. Dev.
real hourly wage ($ 1999)	15,262	14.07	12.275
hours per week	15,262	37.88	11.392
Employer-provided HI (%)	15,262	62.35	

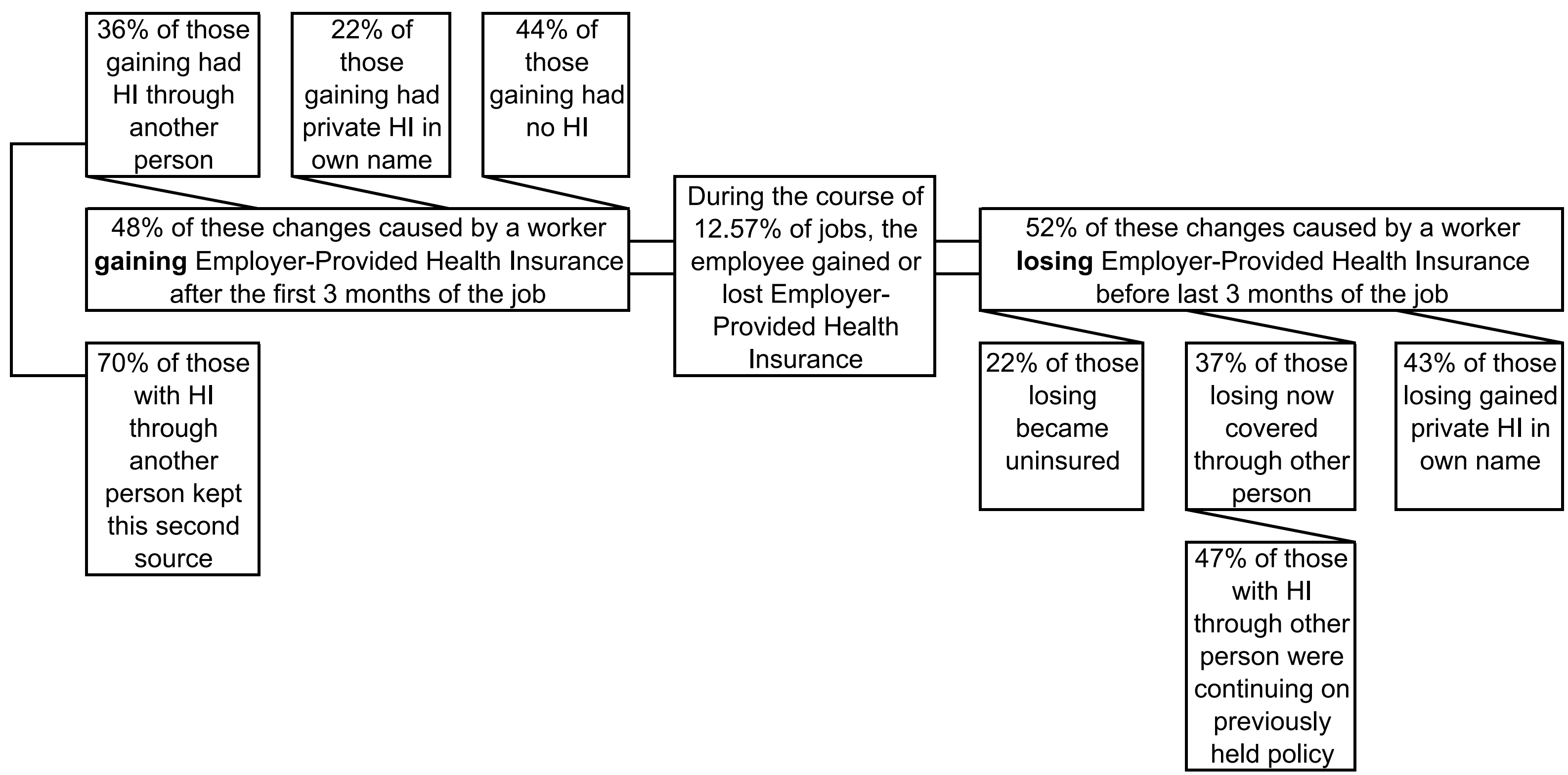

Figure 1: Breakdown of changes in Employer-Provided Health Insurance within Job: 1996 SIPP

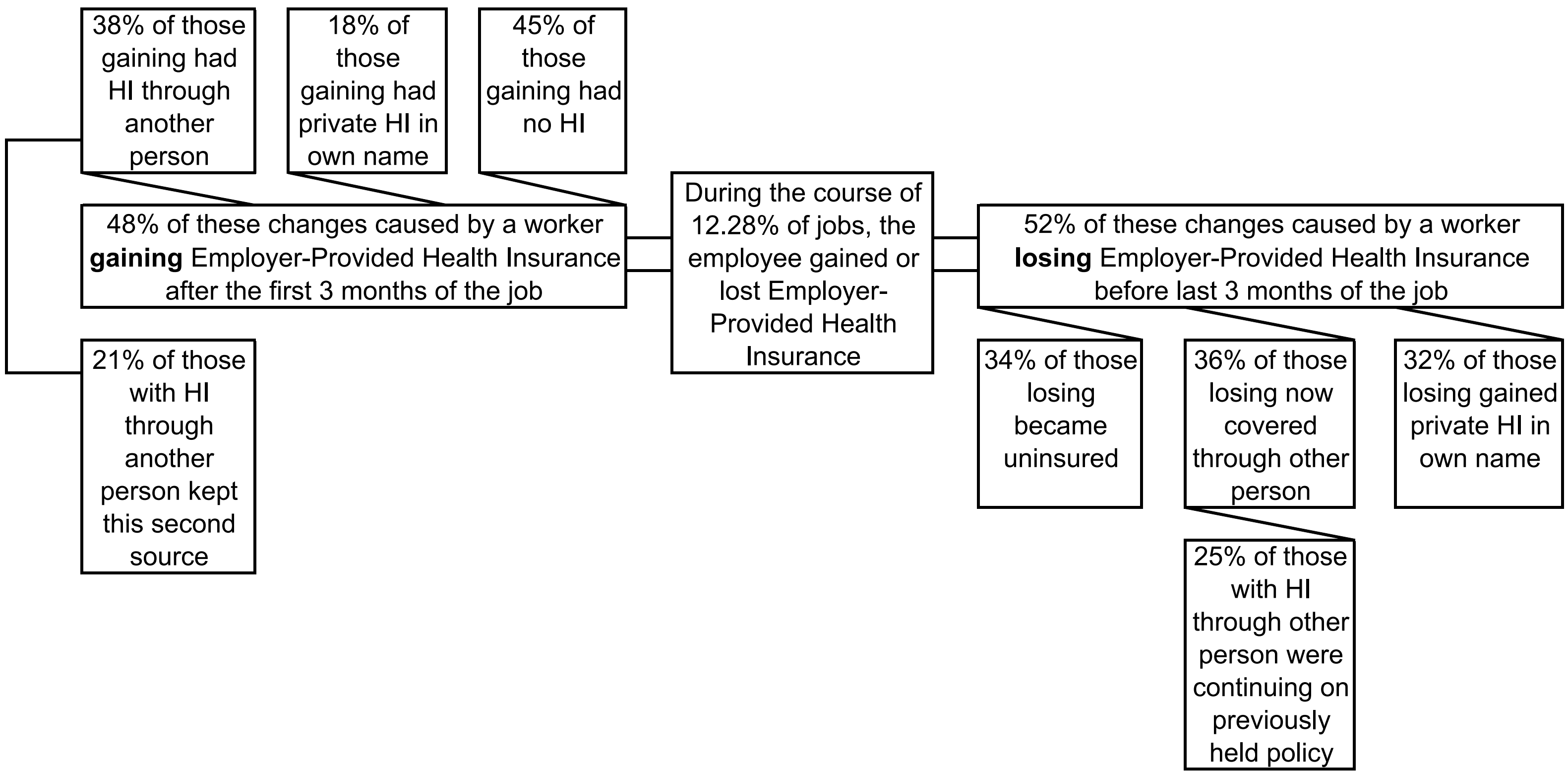

Figure 2: Breakdown of changes in Employer-Provided Health Insurance within Job: 1990 SIPP

Table 5: Joint Hazard, Health Insurance Specification: 1996 SIPP

	(1)	(2)	(3)	(4)
	Hazard only	Hazard only	Hazard,HI joint	Hazard, HI joint
	no heterogeneity	person heterogeneity	correlated person heterogeneity	correlated person and job heterogeneity
Hazard Ratios for Health Insurance Indicators: married, male baseline				
HIEMPLOYER	0.347	0.309	0.357	0.444
	(0.0250)***	(0.0282)***	(0.0253)***	(0.0274)***
HIOTHER	0.719	0.701	0.718	0.716
	(0.0270)***	(0.0320)***	(0.0276)***	(0.0284)***
BOTH	0.488	0.452	0.500	0.613
	(0.0499)***	(0.0537)***	(0.0868)***	(0.0517)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER				
	0.680	0.743	0.679	0.671
Heterogeneity Terms				
St.dev. of person effects in hazard eq. and health insurance eq.; correlation betw. effects				
σ_δ		0.5324 ***	0.1624 ***	0.2587 ***
σ_γ			4.2883 ***	4.2908 ***
$\rho_{\delta\gamma}$			-0.2182 ***	-0.2166 ***
St. dev. of job effect in health insurance eq.; hazard eq. coeff on job effect				
σ_ε			2.6890 ***	2.6945 ***
λ_1				-0.0771 ***

Note on Hazard Equation: In addition to reported controls, the hazard equation was estimated with piecewise linear splines in age, calendar time, labor force experience, and tenure. Controls for race, ethnicity, education level, gender, marital status, number of kids, interaction of marital status and number of kids with gender, own health insurance from non-employer source, union status, industry, and job type were also included. Gender and marital status were interacted with the health insurance variables so these hazard ratios are group-specific; hazard ratio-1= percentage change in the probability of leaving a job;

Note on Health Insur. Equation: probit model for holding employer-provided health insurance; included as controls were: race, gender, ethnicity, education level, existence of chronic health condition, number of kids, marital status, kids and marital status interacted with gender, indicator for weekly hours greater than 20, age at beginning of job, union status, industry and job type

Standard errors are in parantheses; * indicates significance at the 10% level, ** at the 5% level, and *** at the 1% level.

Table 6: Joint Hazard, Health Insurance Specification: 1996 SIPP

	(1)	(2)	(3)	(4)
	no heterogeneity	person heterogeneity	correlated person heterogeneity	correlated person heterogeneity, job heterogeneity
Hazard Ratios for Health Insurance Indicators: single, male baseline				
HIEMPLOYER	0.371	0.335	0.380	0.470
	(0.0253)***	(0.0280)***	(0.0257)***	(0.0277)***
HIOTHER	1.119	1.168	1.128	1.141
	(0.0237)***	(0.0301)***	(0.0245)***	(0.0259)***
BOTH	0.510	0.498	0.528	0.662
	(0.0910)***	(0.09437)***	(0.0918)***	(0.0934)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER				
	0.165	0.183	0.166	0.165

Hazard Ratios for Health Insurance Indicators: married, female baseline				
HIEMPLOYER	0.305	0.265	0.312	0.386
	(0.0321)***	(0.0352)***	(0.0324)***	(0.0345)***
HIOTHER	0.660	0.628	0.652	0.646
	(0.0223)***	(0.0267)***	(0.0227)***	(0.0236)***
BOTH	0.450	0.403	0.461	0.561
	(0.0434)***	(0.0465)***	(0.04387)***	(0.04548)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER				
	0.838		0.856	0.860

Hazard Ratios for Health Insurance Indicators: single, female baseline				
HIEMPLOYER	0.329	0.288	0.336	0.417
	(0.0257)***	(0.0280)***	(0.0260)***	(0.0281)***
HIOTHER	0.944	0.952	0.949	0.953
	(0.0236)**	(0.0296)***	(0.0243)**	(0.0257)**
BOTH	0.471	0.426	0.481	0.606
	(0.1022)***	(0.1095)***	(0.1034)***	(0.1057)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER				
	0.361	0.375	0.356	0.361

Note: In addition to reported controls, the hazard equation was estimated with piecewise linear splines in age, calendar time, labor force experience, and tenure and controls for union status, industry, and job type. Gender and marital status were interacted with the health insurance variables so these hazard ratios are group-specific; hazard ratio-1= percentage change in the probability of leaving a job; standard errors are in parantheses; * indicates significance at the 10% level, ** at the 5% level, and *** at the 1% level.

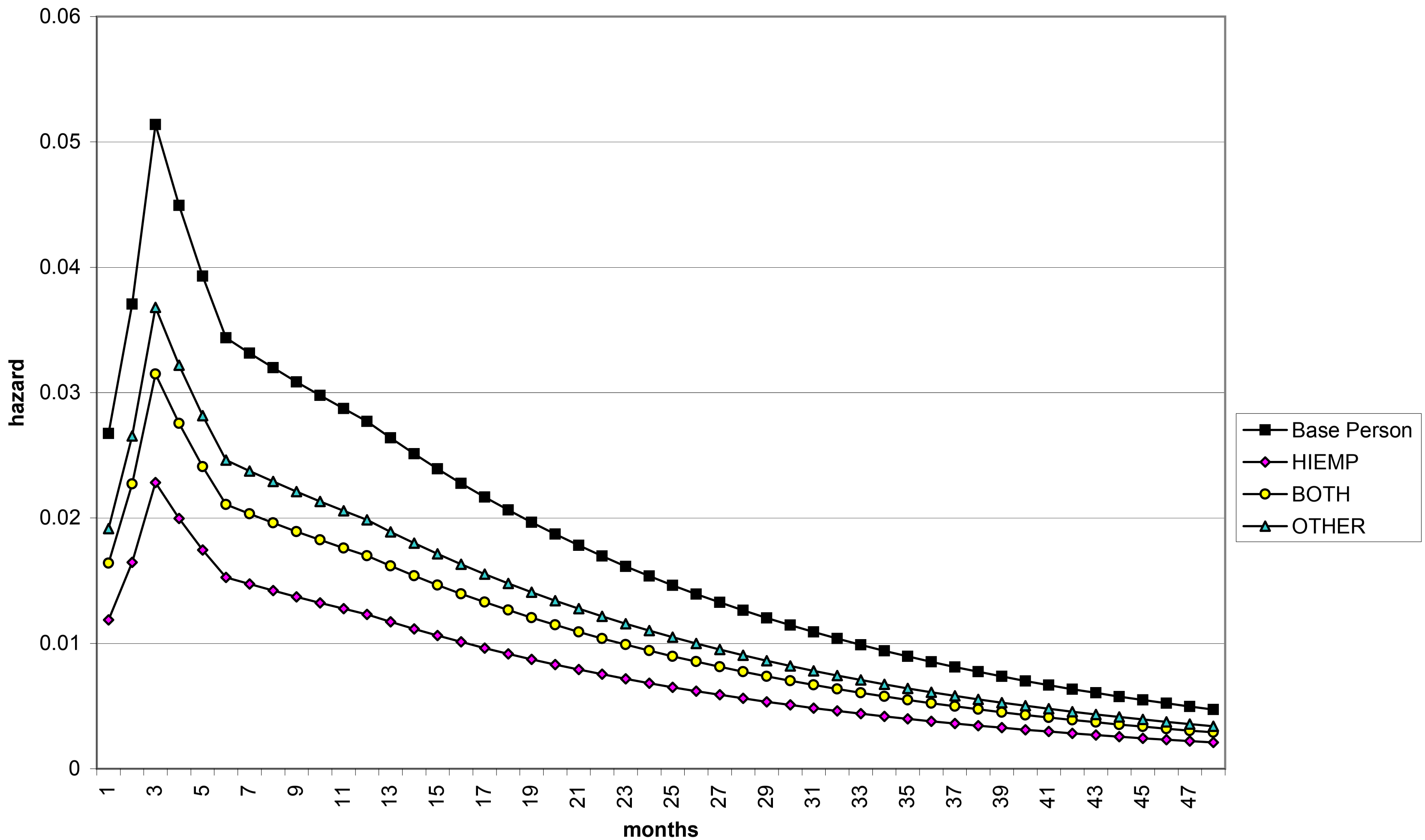

Figure 3: Hazard of Job Ending for Married Men 1996 SIPP

Table 7: Joint Hazard, Health Insurance Specification: 1990 SIPP

	(1)	(2)	(3)	(4)
	Hazard only	Hazard only	Hazard,HI joint	Hazard, HI joint
	no heterogeneity	person heterogeneity	correlated person heterogeneity	correlated person and job heterogeneity
Hazard Ratios for Health Insurance Indicators: married, male baseline				
HIEMPLOYER	0.380	0.368	0.403	0.472
	(0.0387)***	(0.0410)***	(0.0408)***	(0.0458)***
HIOTHER	0.774	0.772	0.764	0.765
	(0.0480)***	(0.0522)***	(0.0517)***	(0.0522)***
BOTH	0.555	0.538	0.581	0.680
	(0.0908)***	(0.0945)***	(0.0946)***	(0.0972)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER				
	0.607	0.611	0.617	0.613
Hazard Ratios for Health Insurance Indicators: single, male baseline				
HIEMPLOYER	0.391	0.377	0.409	0.477
	(0.0390)***	(0.04136)***	(0.0409)***	(0.0463)***
HIOTHER	0.967	0.985	0.975	0.977
	(0.0395)	(0.0453)	(0.0444)	(0.0450)
BOTH	0.731	0.749	0.790	0.947
	(0.2655)	(0.2703)	(0.2722)	(0.2732)
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER				
	0.499		0.509	0.520
Hazard Ratios for Health Insurance Indicators: married, female baseline				
HIEMPLOYER	0.342	0.326	0.356	0.415
	(0.0508)***	(0.0533)***	(0.0527)***	(0.0566)***
HIOTHER	0.699	0.689	0.684	0.682
	(0.0379)***	(0.04128)***	(0.0405)***	(0.0410)***
BOTH	0.563	0.547	0.592	0.689
	(0.0807)***	(0.0838)***	(0.0831)***	(0.0861)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER				
	0.823	0.856	0.860	0.864
Hazard Ratios for Health Insurance Indicators: single, female baseline				
HIEMPLOYER	0.337	0.322	0.351	0.413
	(0.0413)***	(0.0433)***	(0.0427)***	(0.0480)***
HIOTHER	0.970	0.986	0.978	0.979
	(0.0389)	(0.0435)	(0.0425)	(0.0433)
BOTH	0.880	0.886	0.951	1.126
	(0.1973)	(0.2026)	(0.2037)	(0.2060)
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER				
	0.648	0.651	0.654	0.655
Heterogeneity Terms				
σ_δ		0.3640 ***	0.3252 ***	0.3505 ***
σ_γ			5.7055***	5.6900***
$\rho_{\delta\gamma}$			-0.2179 ***	-0.2216 ***
σ_ε			3.6369***	3.6256***
λ_1				-0.0467 ***

Note: Both the hazard and health insurance equations were estimated with the same set of controls as in Table 5; hazard ratio-1= percentage change in the probability of leaving a job; standard errors are in parantheses; * indicates significance at the 10% level, ** at the 5% level, and *** at the 1% level.

Table 8: Joint Hazard, Health Insurance Specification: 1996 SIPP
Health Insurance Coverage Types: single, single+1, multiple

	(1)	(2)
	correlated person heterogeneity	correlated person and job heterogeneity
Hazard Ratios for Health Insurance Indicators: HIEMP covers worker only		
HIEMPLOYER	0.377	0.470
	(0.0176)***	(0.0204)***
HIOTHER	0.815	0.813
	(0.0132)***	(0.0138)***
BOTH	0.622	0.770
	(.0405)***	(.0420)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER		
	0.620	0.619
Hazard Ratios for Health Insurance Indicators: HIEMP covers worker + 1 family member		
HIEMPLOYER	0.347	0.432
	(0.0286)***	(0.0304)***
HIOTHER	0.815	0.813
	(0.0132)***	(0.0138)***
BOTH	0.438	0.540
	(.0769)***	(.0662)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER		
	0.436	0.430
Hazard Ratios for Health Insurance Indicators: HIEMP covers worker + >1 family members		
HIEMPLOYER	0.329	0.409
	(0.0238)***	(.0259)***
HIOTHER	0.815	0.813
	(0.0132)***	(0.0138)***
BOTH	0.372	0.455
	(.0647)***	(.0619)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER		
	0.341	0.331
Heterogeneity Terms		
σ_δ	0.1505 ***	0.2528 ***
σ_γ	4.2485***	4.2937***
$\rho_{\delta\gamma}$	-0.2183 ***	-0.2168 ***
σ_ε	2.6746***	2.6962***
λ_1		-0.0776 ***

Note: Both the hazard and health insurance equations were estimated with the same set of controls as in Table 5; hazard ratio-1= percentage change in the probability of leaving a job; standard errors are in parantheses; * indicates significance at the 10% level, ** at the 5% level, and *** at the 1% level.

Table 9: Joint Hazard, Health Insurance Specification: 1996 SIPP
Health Insurance Coverage Types: worker pays all, part, or none

	(1)	(2)
	correlated person heterogeneity	correlated person heterogeneity, job heterogeneity
Hazard Ratios for Health Insurance Indicators: Worker pays no HIEMP costs		
HIEMPLOYER	0.348	0.440
	(0.0224)***	(0.0246)***
HIOTHER	0.816	0.813
	(0.0131)***	(0.0138)***
BOTH	0.371	0.463
	(.0601)***	(.0566)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER		
	0.290	0.280
Hazard Ratios for Health Insurance Indicators: Worker pays part of HIEMP costs		
HIEMPLOYER	0.317	0.400
	(0.0174)***	(0.0200)***
HIOTHER	0.816	0.813
	(0.0131)***	(0.0138)***
BOTH	0.387	0.481
	(.0474)***	(.0640)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER		
	0.408	0.400
Hazard Ratios for Health Insurance Indicators: Worker pays all of HIEMP costs		
HIEMPLOYER	0.637	0.808
	(0.0420)***	(.0437)***
HIOTHER	0.816	0.813
	(0.0131)***	(0.0138)***
BOTH	0.344	0.422
	(.0603)***	(.0960)***
Difference-in-Difference: (BOTH-HIEMP)/BOTH - (HIOTHER-1)/HIOTHER		
	-0.628	-0.684
Heterogeneity Terms		
σ_δ	0.1522 ***	0.2641 ***
σ_γ	4.2917***	4.2908***
$\rho_{\delta\gamma}$	-0.2182 ***	-0.2165 ***
σ_ε	2.6903***	2.6949***
λ_1		-0.0844 ***

Note: Both the hazard and health insurance equations were estimated with the same set of controls as in Table 5; hazard ratio-1= percentage change in the probability of leaving a job; standard errors are in parantheses; * indicates significance at the 10% level, ** at the 5% level, and *** at the 1% level.

Table 10: Joint Wage, Hazard Specification: 1996 SIPP

	(1)	(2)	(3)	(4)
	Wage only	Wage only	Wage,Haz joint	Wage,Haz joint
	no heterog.	person and job heterog.	corr. person heterog.	corr. person and job heterog.
st.dev. wage person effect: σ_θ		0.3064 ***	0.3064 ***	0.3069 ***
st.dev. hazard person effect: σ_δ			0.5342 ***	0.5456 ***
corr. betw. person effects: $\rho_{\delta\theta}$			-0.0260 *	0.1647 ***
st.dev. wage job effect: σ_ψ		0.3375 ***	0.3374 ***	0.3375 ***
haz eq. coeff. on job effect: λ_2				-0.5299 ***
st.dev. wage residual: σ_η	0.4836 ***	0.2488 ***	0.2488 ***	0.2488 ***
tenure 0-3 months	-0.0040 **	-0.0111 ***	-0.0112 ***	-0.0117 ***
	(0.0017)	(0.0007)	(0.0007)	(0.0007)
tenure 3-6 months	0.0062 ***	0.0028 ***	0.0028 ***	0.0027 ***
	(0.0010)	(0.0004)	(0.0004)	(0.0004)
tenure 6-12 months	0.0059 ***	0.0031 ***	0.0031 ***	0.0032 ***
	(0.0003)	(0.0001)	(0.0001)	(0.0001)
tenure 12-48 months	0.0022 ***	0.0023 ***	0.0023 ***	0.0025 ***
	(0.00002)	(.00003)	(0.00003)	(0.00003)
tenure 48-120 months	0.0018 ***	0.0012 ***	0.0012 ***	0.0013 ***
	(0.000007)	(.00003)	(0.00003)	(0.00002)
tenure 120+ months	0.0008 ***	0.0006 ***	0.0006 ***	0.0007 ***
	(0.000003)	(.00003)	(0.00003)	(0.00003)
high school degree	0.1480 ***	0.1216 ***	0.1218 ***	0.1204 ***
	(0.0004)	(.0082)	(0.0082)	(0.0082)
some college	0.3113 ***	0.2752 ***	0.2753 ***	0.2744 ***
	(0.0004)	(.0079)	(0.0079)	(0.0079)
college degree	0.6087 ***	0.5809 ***	0.5810 ***	0.5806 ***
	(0.0004)	(0.0084)	(0.0084)	(0.0084)
graduate degree	0.8378 ***	0.8277 ***	0.8279 ***	0.8272 ***
	(0.0005)	(0.0097)	(0.0098)	(0.0098)
experience 0-12 months	-0.0055 ***	0.0004	0.0004	-0.0003
	(0.0011)	(0.0008)	(0.0008)	(0.0008)
experience 12-24 months	0.0066 ***	0.0049 ***	0.0049 ***	0.0047 ***
	(0.0004)	(0.0003)	(0.0003)	(0.0003)
experience 24-60 months	0.0010 ***	0.0024 ***	0.0024 ***	0.0023 ***
	(0.00004)	(0.0001)	(0.0001)	(0.0001)
experience 60-120 months	0.0012 ***	0.0018 ***	0.0018 ***	0.0018 ***
	(0.00001)	(0.00003)	(0.00003)	(0.00003)
experience 120-240 months	0.0010 ***	0.0010 ***	0.0010 ***	0.0010 ***
	(0.000003)	(0.00002)	(0.00002)	(0.00002)
experience 240+ months	-0.0001 ***	-0.0001 ***	-0.0001***	-0.0001***
	(0.000001)	(0.00002)	(0.00002)	(0.00002)

Note: Dependent variable is real monthly wages; In addition to reported controls, the wage equation was estimated with continuous calendar time and controls for gender, race, marital status, number of kids, interactions of gender, marital status, and number of kids, union status, industry, and job type. Wages are in 1999 dollar terms; Hazard equation was estimated with same set of controls as in Table 5. standard errors are in parantheses; * indicates significance at the 10% level, ** at the 5% level, and *** at the 1% level; excluded category=female, no high school degree, non-white, non-Hispanic, non-married, non-union,wholesale trade industry, private-for-profit employer.

Table 11: Joint Hazard, Health Insurance, and Wage Specification

	1990 SIPP	1996 SIPP
Heterogeneity Terms		
st.dev. hazard person effect: σ_δ	0.4291 ***	.5863***
st.dev. wage person effect: σ_θ	0.3096 ***	.3087***
st.dev. hi person effect: σ_γ	5.2214 ***	5.4395***
corr. hazard, wage person effects: $\rho_{\delta\theta}$	-0.2129 ***	.1764***
corr. hazard, hi person effects: $\rho_{\delta\gamma}$	-0.5287 ***	-.3242***
corr. wage, hi person effects: $\rho_{\theta\gamma}$	0.5400 ***	.3202***
st.dev. wage job effect: σ_ψ	0.2959 ***	.3432***
st.dev. hi job effect: σ_ε	4.9552 ***	4.8109***
corr. wage, hi job effects: $\rho_{\psi\varepsilon}$	0.2675 ***	.3468***
haz eq. coeff. on hi job effect: λ_1	-0.0627 ***	-.0894***
haz eq. coeff. on wage job effect: λ_2	-0.0980***	-.4259***
st.dev. residual: σ_η	0.2333 ***	.2487***
Wage Equation: Tenure Coefficients		
tenure 0-3 months	-0.0013	-.0121
	(0.0011)	(.0007)***
tenure 3-6 months	0.0051	.0028
	(0.0006)***	(.0004)***
tenure 6-12 months	0.0021	.0040
	(0.0002)***	(.0001)***
tenure 12-48 months	0.0022	.0031
	(0.00005)***	(.00003)***
tenure 48-120 months	0.0006	.0016
	(0.00005)***	(.00003)***
tenure 120+ months	0.0004	.0009
	(0.00004)***	(.00002)***
Hazard Ratios for Health Insurance Variables		
HIEMPLOYER	0.641	0.612
	(.0400)***	(.0226)***
HIOTHER	0.817	0.790
	(.0246)***	(.0162)***
BOTH	1.047	0.855
	(.0662)***	(.0357)***
Difference-in-Difference	0.611	0.549

Note: Joint hazard, health insurance, wage equations were estimated with controls for male, white, Hispanic, marital status, number of kids, interactions of kids and marital status with male, education levels, union status, and industry; hazard equation contained piecewise linear splines in age, tenure, labor force experience, and calendar time and a control for own HI from private source; wage equation contained spline in tenure coefficients and general labor market experience; health insurance equation contained age at beginning of job, health condition indicator, and hours worked per week; standard errors are in parantheses; * indicates significance at the 10% level, ** at the 5% level, and *** at the 1% level;