Labor Market Rigidities and the Employment Behavior of Older Workers[1]

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Abstract

The labor market is often asserted to be characterized by rigidities that make it difficult for older workers to carry out their desired trajectories from work to retirement. In this paper we address the following question: what is the association between the age composition of employment in an establishment and the propensity of older workers to separate from the establishment? In the absence of a direct measure of labor market rigidity, we use the share of older workers in an establishment's workforce as a proxy for the "older-worker-friendliness" of an establishment. We argue that establishments with a relatively large share of older workers, other things equal, are less likely to use technology or employment practices that result in labor market rigidities. As a result, older workers are more likely to be able to carry out their desired trajectory from work to retirement without separating from the firm. Our analysis uses longitudinal data on individuals from the U.S. Survey of Income and Program Participation merged with data on their employers from the Longitudinal Employer-Household Dynamics files. We use a difference-in-difference approach to analysis of the association between the age composition of employment in an establishment and the rate at which workers of different ages separate from the establishment. We find strong evidence that an older age structure of the work force at the establishment-level is associated with a lower separation propensity of its older workers, relative to the separation propensity of its younger workers. This finding is robust to many specification checks. These results provide indirect but suggestive evidence of the importance of labor market rigidities.

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1. Introduction

The labor market is often asserted to be characterized by rigidities that make it difficult for older workers to carry out their desired trajectories from work to retirement. The rigidities that are cited include lack of opportunity for part-time and flexible-hours work at many establishments; low wages and lack of fringe benefits in the part-time employment opportunities that are available; and lack of training and promotion opportunities for older workers both at their career employers and at potential new employers (Hurd, 1996). It is important to assess the extent of rigidities in the labor market and their impact on older workers, because workers who cannot carry out their optimal labor supply trajectory suffer a welfare loss. The economy loses the production and earnings of older workers who would like to work but cannot find a job with the desired hours and conditions and choose retirement instead. In addition, the government loses tax revenue, and the workers switch from being contributors to being claimants for Social Security. The approaching retirement of the baby boom generation and overall population aging amplify the importance of this issue. These demographic factors have raised concerns about whether labor supply will remain sufficient to meet employers' needs and whether Social Security and Medicare will remain solvent.

This paper provides new insights on the labor market for older workers by using rich longitudinal survey data on individuals matched to employment and earnings data on the establishments that employ them. The individual data are from the Survey of Program Participation (SIPP) and the employer data are from the Longitudinal Employer-Household Dynamics (LEHD) files (Abowd, Haltiwanger, and Lane, 2004). We address the following question: what is the association between the age composition of employment in an establishment and the propensity of older workers to separate from the establishment? In the absence of a direct measure of labor market rigidity, we use the share of older workers in an establishment's workforce as a proxy for the "older-worker-friendliness" of an establishment. We argue that establishments with a relatively large share of older workers, other things equal, are less likely to use technology or employment practices that result in labor market rigidities. As a result, older workers are more likely to be able to carry out their desired trajectory from work to retirement without separating from the firm. Hence we predict that a greater share of older workers in an establishment's workforce should be associated with a lower propensity for older workers to separate from the establishment.

We use a difference-in-difference approach to analysis: compare the job exit behavior of older and younger workers in establishments with a relatively large share of older workers to the job exit behavior of older and younger workers in establishments with a relatively small share of older workers. The availability of detailed establishment-level data on the age composition of employment allows us to experiment with alternative definitions of large and small proportions of older workers employed by establishments. Taking the difference between the employment behavior of older and younger workers makes it possible to disentangle the effects of labor market rigidities that affect all workers from those that are specific to older workers. In order to ensure that the establishment's age composition is not merely picking up the effects of other factors, we control for the worker's demographic characteristics, pension and health insurance coverage, wage rate, wealth, health, industry, occupation, and location. In some specifications we control for the *industry*-level age structure, in order to distinguish between the effects of industry-specific and establishment-specific age structure.

We use data from the 1990 – 2001 SIPP panels merged with *establishment*-level data on the age distribution of employment, derived from the LEHD data. In this version of the paper we report results from a sub-sample of the LEHD records that can be linked to the SIPP. The full sample of the LEHD records that are potentially matchable to SIPP was recently made available to us, and will be used in the next version of this paper. We find strong evidence that an older age structure of the work force at the *establishment*-level is associated with a lower separation propensity of older workers. This finding is robust to many specification checks.

The paper is organized as follows. The next section discusses previous evidence of the existence of labor market rigidities. Section 3 describes a theoretical framework for the analysis. Description of the data and methodology are provided in Section 4. Section 5 presents estimation results and their interpretations. Section 6 concludes with implications of the estimates and plans for future research.

2. Background and Literature

If tastes for leisure or demand for time in home production increase gradually at older ages, then other things equal workers might prefer to gradually reduce hours of work or partially retire as they age, before completely withdrawing from the labor force. Many studies have documented the existence of partial retirement and "bridge jobs" as a type of labor market withdrawal process (Gustman and Steinmeier, 1984, Ruhm, 1990, Blau, 1994, Maestas, 2004), but the majority of workers retire by moving directly from full-time employment to complete retirement. Data from the Health and Retirement Survey (HRS) shows that in a sample of individuals aged 51-72 who were employed full-time (35 or more hours per week) year-round (36 or more weeks per year) on a long-tenure job (at least 5 years) in any of the first five waves of the survey, two-year transition rates were 17.2% to non-employment compared to 5.6% to part time on the same job, 4.0% to a full-time year-round job with a new employer, and 2.7% to part time with a new employer (authors' calculations).

Thus, it seems doubtful that worker preferences alone can explain the predominance of the abrupt retirement pattern, assuming smooth changes in preferences for leisure and randomness of shocks. Several pieces of evidence suggest that factors other than individual preferences and shocks are at least partly responsible for the typical pattern of abrupt retirement.

First, self-employed individuals are much more likely to retire gradually than are otherwise similar wage-salary employees. It has been argued that self-employment offers greater flexibility in hours to accommodate changing tastes for leisure, thus facilitating gradual retirement (Quinn, 1980). Karoly and Zissimopoulos (2004) report that workers age 45 and older represented 38% of the workforce in total, but made up 54% of the self-employed in 2002. About one third of older self-employed workers entered self-employment after age 50. Karoly and Zissimopoulos also find that while average hours worked per week was similar for self-employed workers and employees, 59% of the self-employed worked full time compared to 74% of wage and salary workers. This suggests greater flexibility in choosing hours of work in the self-employment sector. In the HRS sample described above, the two-year transition rate from a full-time year-round long-tenure job to part-time employment (on the same job or a new job) was 7% for individuals who were employees and 16% for individuals who were self-employed, further suggesting that wage-salary workers face hours constraints imposed by their employers.

Second, the predominance of abrupt transitions from full time employment to non-employment could in principle be explained by health shocks. There is no doubt that health plays a major role in the timing of retirement (Blau and Gilleskie, 2001, Bound, 1991), but the majority of workers who follow the typical pattern of moving from a career employer directly to retirement appear to be in good health. In the HRS sample described above, 13% of two-year transitions to non-employment were associated with a change in self-reported health from “good” (excellent, very good, or good) to “bad” (fair or poor), compared to 68% who reported good health both before and after the transition. In comparison, of individuals who remained in the same full-time year-round long-tenure job between waves, 7% reported a change in health from good to bad, compared to 82% who reported good health both before and after the transition. As for other shocks, Maestas (2004) finds no significant differences in pre-retirement resources, preferences, expectations and their post-retirement realizations and retirement satisfaction between groups of individuals who retire abruptly and those who follow other paths to retirement.

Finally, when asked directly in surveys, many older workers who are employed full time state that they could not reduce the number of hours they work at their current employer (Hurd, 1996, Abraham and Houseman, 2004). Abraham and Houseman (2004) find that even though the fraction of older working Americans who plan to reduce their work hours or change the type of work around retirement age is almost equal to the fraction that plan to retire fully, the former are only about half as likely as the latter to actually follow through on their plans.

Many factors could be responsible for making the labor market rigid. As defined by Hurd (1996, p.12), "labor market rigidities are employment practices and work-related financial arrangements that constrain or limit the volume of work with respect to hours per day, days per week, or weeks per year" with the current employer or when changing employers. "Rigidities also include situations in which the volume of work can be varied, but the change requires a disproportionate sacrifice in compensation, job satisfaction, mental or physical requirements, or location". It has been well documented that many workers face strong discontinuities in retirement incentives that result from government policy and labor market institutions. Social Security and Medicare have strictly defined age eligibility criteria that may affect the employment behavior of individuals who face a significant liquidity constraint. The Social Security Earnings Test places a large implicit tax on earnings above a certain threshold prior to the normal retirement age. This has been found to affect employment behavior at those ages (Burtless and Moffitt, 1985, Friedberg, 2000). The Employee Retirement and Income Security Act (ERISA) prohibits workers from receiving benefits from a Defined Benefit (DB) pension plan while working at the establishment that provides the benefits. In addition, most DB plans link benefits to earnings in the last few years on the job, reducing a worker's incentive to decrease work hours at the career employer. Because they are not portable across establishments, these pension plans may further impede workers from changing employers in search of desired work-hours flexibility. Also, older workers who are covered by an employer-provided health insurance plan and have a health problem that requires medical attention may be reluctant to change employers (Scott, Berger, and Garen, 1995).

However, these factors alone cannot fully account for the prevalence of abrupt retirement, because abrupt retirement is the most common pattern even for individuals who don't face liquidity constraints, are not covered by DB pension plans, and have retiree health insurance. This suggests that other sources of labor market rigidity may be important. On the demand side of the labor market, if there are fixed costs to establishments of hiring, training, and employing a worker, then establishments may prefer to hire and employ full-time rather than part-time workers. If production takes place in teams, then the absence of a team member could reduce team productivity. In this case establishments might require the presence of workers at specific times, reducing the flexibility of workers in scheduling their hours of work. If monitoring worker effort is costly, then establishments may backload compensation so as to provide incentives to workers to avoid shirking. This results in compensation that exceeds a worker's marginal product at older ages, so the establishment might also specify a terminal date for employment. This could be implemented by mandatory retirement, or, if this is illegal, by structuring the pension so as to provide strong incentives for older workers to leave the establishment (Lazear, 1979). Workers could face statistical discrimination in the labor market as a result of the application of group characteristics to all members of the group (Hellerstein, Neumark, and Troske, 1999). For example, the short expected duration of future employment of an older worker reduces the incentive of the establishment to train and promote older workers (Hutchens, 1988), despite the fact that some older workers may plan to remain employed for a long time. If human capital is establishment-specific, it creates a wedge between the worker's wage at the current establishment and at other establishments. A worker might have to take a substantial pay cut in order to change employers.

Some of these sources of labor demand rigidities are caused by features of the technology of production that may affect all workers, not just older workers. But if the hours-of-work preferences of older workers differ systematically from those of younger workers, then the existence of technology-induced rigidities will be manifested in the age structure of an establishment's work force: the more important are technology-induced rigidities, the lower is the share of older workers at a establishment. There is evidence that production technology differs substantially across establishments, even within narrowly defined industries (Doms, Dunne, and Troske, 1997). These differences are hypothesized to arise from variation across establishments in managerial ability, expectations of future price and technological change, and past investment decisions (Davis and Haltiwanger, 1991). Thus, while technology cannot be measured directly, with establishment-level data it may be possible to detect evidence of technology-based rigidities if such rigidities are manifested in differences in the age structure of the work force across establishments.

The discussion above suggests that data on individual workers matched to data on the age distribution of employment at the establishments that employ them can be used to test for the existence of technology-driven labor market rigidities that affect older workers differently from younger workers. The key observable implications of technology-driven labor market rigidity are that (1) there will be variation across establishments in the age composition of employment within industries, and (2) such variation will be associated with variation in hours worked and employment turnover of older workers relative to younger workers. Specifically, we expect that if labor market rigidities are important, then older workers employed at an establishment with a smaller share of older workers will be more likely to exit the establishment, compared to younger workers, than will older workers at an otherwise similar establishment with a greater share of older workers, again compared to their younger counterparts, other things equal.

3. An Illustrative Model

We illustrate the logic of our conceptualization of labor market rigidities and their impact on the employment behavior of older workers using a very simple prototype of a two-sector equilibrium model of the labor market. There are two types of firms that differ by the technology employed. Type A firms use a technology that does not have any features associated with labor market rigidity, while type B firms use a technology that has at least one such feature. We use the example of team production here. The type A technology is standard: $Q = F^A(L_A, K_A)$ where Q is output, L is hours of labor input, and K is capital input. We assume that the marginal product of labor (MPL) is a continuous function of L_A, and is (eventually) smoothly declining in L_A for a given value of K. Thus a type-A firm is indifferent to the number of hours worked by any particular worker. The type B production function is $Q = F^B(L_B*(min\{L_1, L_2, ..., L_N\})^\theta, K_\mathrm{B})$, where L_i is the number of hours worked by the i^{th} worker, there are N workers employed by the firm, and $L_B = \Sigma L_i$. In this technology, there is a productivity bonus of $\theta \geq 0$ for every hour in which all members of the "team" of N workers are present, (assuming, for example, all workers begin the workday at the same time). We take N to be a parameter of the technology: team size must be no smaller than N in order to realize any gains from team production, and (in this simple example) there is no additional gain to a team size greater than N (see Coles and Treble, 1996, for a similar approach). If $\theta = 0$, then the production function is of the standard non-team type, and there will not be any labor market rigidity (the constraint of hiring N workers in this case is not binding). If $\theta > 0$, then the labor input for a type B firm has a fixed coefficients component in which the MPL = 0 for that component unless all team members increase hours worked jointly. Hence if $\theta > 0$, then a type B firm has an incentive to require all workers to work the same number of hours. We assume that type-B firms respond to this incentive by requiring all workers to work the same number of hours, denoted L_{iB}. The type-B production function can then be rewritten as $Q = F^B(L_B L_{iB}{}^\theta, K_\mathrm{B}) = F^B(NL_{iB}L_{iB}{}^\theta, K_\mathrm{B}) = F^B(NL_{iB}{}^{1+\theta}, K_\mathrm{B})$, where total labor input L_B is by definition equal to NL_{iB}, the number of workers multiplied by hours per worker.

Taking the capital input, the price of output, and the hourly wage rate in sector A, W_A, as given, a type A firm chooses the total number of labor hours demanded, $L_A{}^D$, to maximize profit. Taking capital input, price, team size N, and the hourly wage rate in sector B, W_B, as given, a type B firm chooses the number of hours demanded per worker, L_{iB}, to maximize profit, with the resulting total number of labor hours demanded by a type B firm given by $L_B{}^D = NL_{iB}$. We assume homogeneous firms within sector.

Workers spend two periods in the labor market. The utility function of a young worker is

$U(C, T\text{-}L, \delta)$, where C is consumption, T is total available time, L is hours of work, $T\text{-}L$ is hours of leisure, and $\delta>0$ is a parameter such that the marginal utility of leisure is increasing in δ. δ varies across individuals in the population according to the continuous cumulative distribution function $G(\delta)$. There is no access to the capital market, so consumption is given by $C = WL + Z$, where Z is nonwage income, assumed for simplicity to be the same for all workers. The utility function of an old worker is $U(C, T\text{-}L, \delta(1+\eta))$, where $\eta > 0$ is a constant. Thus we assume that the marginal utility of leisure increases proportionately with age for all individuals, other things equal. All individuals are assumed to be employed in both periods. Workers are homogeneous in productivity both within and across periods, and there is no cost of changing jobs.

Workers choose the sector (type of firm, A or B), and in sector A the number of hours of work, to maximize utility, taking W_A, W_B, L_{iB}, and Z as given. In equilibrium, the marginal young worker must be indifferent between working in the two sectors. The marginal young worker is defined by the value of the preference parameter, δ_Y, such that $V^A(W_A, Z, \delta_Y) = V^B(W_B, Z, \delta_Y, L_{iB})$, where V^A and V^B are indirect utility functions. A similar condition must hold in equilibrium for the marginal old worker, who is defined by the value of δ_O such that $V^A(W_A, Z, \delta_O, \eta) = V^B(W_B, Z, \delta_O, \eta, L_{iB})$. Define $L_{AY}(W_A, Z, \delta)$ as the labor supply function of a young worker in sector A, and $L_{AO}(W_A, Z, \delta, \eta)$ as the corresponding function of an old worker. There are at least two qualitatively different types of equilibrium in this model. We focus on the type that is of most relevance for our purposes. Thus suppose that, in equilibrium, optimal hours of work of the *marginal* young worker in sector A, $L_A(W_A, Z, \delta_Y)$ is less than required hours demanded in sector B, L_{iB}. And assume the same inequality holds for the marginal old worker: $L_A(W_A, Z, \delta_O, \eta) < L_{iB}$. Then young workers with $\delta > \delta_Y$ (relatively strong preference for leisure) choose sector A and work relatively short hours, and those with $\delta \leq \delta_Y$ choose sector B and work the relatively long hours required by firms in sector *B*. A similar pattern holds for old workers too, although the specific distribution of hours in sector *A* will differ by age. These inequalities determine labor supply to each sector. Normalizing the number of workers of each age to one, total hours of labor supplied to sector *B* by young workers is $L_{BY}^{S} = L_{iB}G(\delta_{Y}(W_A, W_B))$ and by old workers is $L_{BO}^{S} = L_{iB}G(\delta_{O}(W_A, W_B))$, where the dependence of the reservation values δ_Y and δ_O on wages is made explicit. Total hours of labor supplied to sector *A* by young workers for given wage rates is:

$$L_{AY}^{S} = \int_{\delta_{Y}(W_A, W_B)}^{\delta_{Max}} L_{AY}(W_A, Z, \delta)\mathrm{d}G(\delta), \text{ and by old workers is } L_{AO}^{S} = \int_{\delta_{O}(W_A, W_B)}^{\delta_{Max}} L_{AO}(W_A, Z, \delta, \eta)\mathrm{d}G(\delta),$$

where δ_{Max} is the upper limit of the support of the distribution of δ.

The model is closed by the assumption of market clearing: the quantity of labor supplied equals the quantity of labor demanded in each sector: $L_{AY}^{S} + L_{AO}^{S} = L_{A}^{D}$, $L_{BY}^{S} + L_{BO}^{S} = L_{B}^{D}$. These two conditions determine the equilibrium values of W_A and W_B, which in turn determine the threshold values δ_Y and δ_O, and hours of work required per worker in sector *B*, L_{iB}.

The nature of the equilibrium depends on the values of the parameters. In one type of equilibrium, the wage in sector *B* is higher than the wage in sector *A*, and sector *A* has a greater share of older workers than in sector *B*. In this type of equilibrium, there is a compensating wage differential for the rigid hours in sector *B*, and older workers tend to prefer the flexible-hours sector (*A*) since they can work fewer hours in that sector. We parameterized the model with Cobb-Douglas production and utility functions and a uniform distribution for *G* and solved for the equilibrium numerically (an analytic solution does not exist). We verified that the type of equilibrium just described does in fact exist for many parameter configurations. This demonstrates that technologically-determined labor market rigidities can cause variation in the age structure of employment across firms, although other causes of such variation may exist as well. In contrast, if the technology in sector *B* is standard instead of team-based ($\theta = 0$), then in equilibrium the fraction of old workers is the same in both sectors, regardless of whether leisure preferences increase with age. Hence a test for the (non-) existence of labor market rigidity in this model is a test of the null hypothesis $\theta = 0$.

The empirical analysis in this paper examines the effect of the share of older workers in an establishment on the propensity of older workers to separate from the establishment. In order to develop a prediction for this effect, we add to the model a third period of life in which an individual has the option of retiring in addition to the choice of sector. Thus, individuals now pass through three stages, young, old, and elderly. The utility function of an elderly individual is $U(C, L, \delta(1+\gamma))$, where $\gamma>\eta>0$. Thus, we assume that elderly individuals experience another preference shift toward leisure. There is a reservation value of δ for elderly individuals, δ_e, such that for $\delta \leq \delta_e$ an elderly individual chooses sector *B* and works L_{iB} hours. There is a second reservation value of δ for elderly individuals, δ_e^*, such that for $\delta_e^* \geq \delta > \delta_e$ an elderly individual chooses sector A and works $L_{Ae}(W_A, Z, \delta, \gamma) < L_{iB}$, and for $\delta_e^* \leq \delta$ the individual works zero hours (retires). As preferences shift toward leisure, some individuals who worked in sector *A* while old will retire when elderly. And some individuals who worked in sector *B* while old will move to sector *A* in order to reduce hours of work, while others may move directly from sector *B* to retirement. Given the assumption that all individuals experience a shift in preferences toward leisure, there is no movement from *A* to *B* or from retirement to employment. We solved the three-period version of the model numerically and verified that the same type of equilibrium described above for the two-period model exists for the three-period version as well: sector *A* has a greater share of older workers than sector *B*. As before, this is a result of the technology-induced rigidity in hours of work in sector *B*.

The question of interest now is whether workers who chose sector B when old are more likely to separate from their employer when elderly than are workers who chose sector A when old. That is, how does the exit rate from sector B compare to the exit rate from sector A? Since firms are assumed to be identical within sectors, the only exits that occur in the model are movements from sector B to A, sector B to retirement, and sector A to retirement. There does not appear to be any general result on the rate of exit from sector B compared to the rate of exit from sector A. We parameterized the three period version of the model and solved for the equilibrium for many alternative combinations of parameters. In every case in which an equilibrium of the type of interest exists, the exit rate from B was in fact higher than the exit rate from A. Thus, while the model does not deliver a prediction for the effect of interest, it is certainly consistent with the idea that labor market rigidity can result in a negative association between the share of older workers in an establishment, as a proxy for the degree of technological-induced rigidity, and the separation rate from the establishment.

4. Methods

4.1 Empirical Specification

Our empirical specification can be viewed as an approximation to the employment decision rule of a worker. Life cycle models of the employment behavior of older workers imply that the employment decision in a given period depends on health, demographic characteristics, the wage offer, net worth, potential Social Security and pension benefits, and health insurance coverage (Rust and Phelan, 1997; Blau and Gilleskie, in press; Van der Klaauw and Wolpin, 2005). We augment this list with a measure of the age composition of employment at the individual's establishment. As noted above, taking the difference between the employment behavior of older and younger workers makes it possible to disentangle the effects of labor market rigidities that affect all workers from those that are specific to older workers only. A simple illustration of our empirical specification is

$$\Pr(E_{ijt} = 1 \mid E_{ijt-1} = 0) = F(X_{ijt}\beta + \alpha A_{it} + \gamma R_{ij} + \delta A_{it}{*}R_{ij} + I_{ijt}\eta)$$

where $E_{ijt} = 1$ if individual i employed at establishment j at the beginning of period t separates from the establishment during period t, and equals 0 otherwise; X is a vector of individual and establishment characteristics; $A_{it} = 1$ if the individual is classified as an older worker in period t; R_{ij} is the proportion of older workers in the work force of establishment j; and I is a vector of industry dummies. This is a hazard model of the risk of separation, and is estimated as a logit. In the next version of this paper, we will analyze other outcomes as well, including the type of separation (quit or layoff), the destination of the separation (new job, non-employment), and hours of work.

The coefficient of interest is δ: the difference between the effect of the proportion of older workers on the separation propensity of older and younger workers (these are not the same parameters as in section 3). The main effect of age on employment behavior is captured by α. The main effect of the age composition of the establishment's work force γ captures any effects of workforce age composition on employment behavior that are independent of the worker's own age. For example, establishments with relatively few older workers may tend to be younger, and establishment age may affect the separation propensity of all workers at the establishment. The interaction effect δ captures any differences in the effects of the establishment's age composition on older workers relative to younger workers. Controlling for pension and health insurance coverage, occupation, and the wage rate (all included in X) as well as industry, we interpret differential effects of an establishment's workforce age composition on older versus younger workers as an indication that labor market rigidities affect the employment decisions of older workers.

4.2 Data

We merge longitudinal data on individuals from the U.S. Survey of Income and Program Participation (SIPP), 1990 – 2001 panels, with longitudinal data on their employers from the Longitudinal Employer Household Dynamics (LEHD) files. The SIPP collects detailed information on employment, demographic characteristics, and receipt of income from public programs. Sample members are interviewed every four months for 2½ to 4 years. Each interview wave records employment information separately for each of the four months since the previous interview, so a monthly record of employment, hours of work, earnings, industry, occupation, class of worker, and health insurance coverage for each job can be constructed. The SIPP topical modules, administered once or more per panel, record information on annual income, assets, health, retirement accounts, pension coverage, and employment history prior to the sampling period. The SIPP collects employment data for up to two jobs held during a given month. If an individual holds two jobs in a given month, we analyze behavior only on the *main* job, which we define to be the one with greater work hours per week. If hours per week are equal, we select the job which has been in progress longer.

The LEHD Infrastructure File system is based on state Unemployment Insurance (UI) administrative files with data currently available from 31 states covering about 80% of U.S. employment for the years 1990-2004, although the period covered varies by state (Abowd, Haltiwanger, and Lane, 2004). Employers covered by UI file a quarterly report for each individual who received any covered earnings from the employer in the quarter. An "employer" in this context is a UI-tax-paying entity, roughly equivalent to an establishment. Coverage is about 96% of private non-farm wage-salary employment, with lower coverage of agricultural and government workers, and no coverage of the unincorporated self-employed. The UI records contain information on the quarterly earnings of each individual from each employer for which he has any covered earnings during the quarter; the individual's Social Security number; and an identification number for the business, the State Employer Identification Number (SEIN). In addition to the UI records, partner states also deliver an extract of the file reported to the Bureau of Labor Statistic's Quarterly Census of Employment and Wages, formerly known as ES-202. These data are then merged with the Census Personal Characteristics File, which contains date and place of birth, sex, and a crude measure of race/ethnicity. About 96% of workers in the LEHD data files have this basic demographic data merged in; for the remaining 4% it is imputed, as described in LEHD Program (2002). An extensive discussion of the construction and the content of these files is provided in Abowd *et al.* (2006).

The key to our empirical analysis is matching workers in the SIPP sample to their employer or employers in the LEHD data. The Census Bureau maintains a master LEHD file with confidential worker identifiers. We were provided with LEHD data for all of the workers in the 1990-2001 SIPP panels data who appeared in any LEHD record, after stripping the confidential worker identifiers from the file [for this version of the paper, we actually have LEHD records for only a sub-sample of our SIPP sample]. For a given SIPP sample member, this file contains a record for every available quarter for every establishment that employed the worker from 1990 (or later, if the LEHD records for the state in which the individual was employed begin after 1990) through 2004. The LEHD record for a given establishment in a given quarter contains a stable establishment identifier (the SEIN), an industry code, and earnings and basic demographic data on the SIPP worker *and on all other workers employed at the* *establishment in that quarter*. This enables us to construct measures of the age (and earnings) distribution at the establishment and the establishment's size, both in the given quarter, and averaged over all available quarters. The latter provides a more stable longer-run measure that is not subject to transitory quarter-to-quarter variation.

We match LEHD and SIPP records as follows. If the individual reports in the SIPP that he held only one job during a given calendar quarter, and if there is only one employer record in the LEHD for the individual for that quarter, we match the employer record in the LEHD to the job in the SIPP for that quarter. If the LEHD records two different employers for an individual in a given calendar quarter, and the two employers have different industry codes, we match by industry to the industry code for the main job in the SIPP[2]. If the same industry codes are reported for the two LEHD employers, we check whether either job was matched to an LEHD employer in an earlier quarter. If so, this identifies the job-employer correspondence in the current quarter as well, since the SEIN does not change over time.

The percentage of all SIPP person-months in our sample that can be matched to an LEHD record is 62%. Failure to match can occur for several reasons. First, the LEHD file system is based on UI records and thus contains data only for workers who were employed in the UI-covered sector as wage-salary employees. Second, only about 80% of the SIPP sample members have a Social Security number available. The Social Security number is the basis for the confidential worker identifier that makes a link to the LEHD possible. Third, not all states provide data to the LEHD. Fourth, many states joined the LEHD program after 1990, so there are no data for such states for the early part of the SIPP sample. Finally, for person-months in which

[2] The SIPP provides three digit 1990 Census industry codes, while the LEHD provides six digit codes based on the 1997 North American Industry Classification System (NAICS). A crosswalk available from the Census Bureau web page http://www.census.gov/hhes/www/ioindex/indcswk2k.pdf was used for matching.

an individual held two jobs in the same industry, and neither job was matched to an LEHD employer in an earlier quarter, a match is not possible.

Table 1 presents summary statistics for two samples used in our analysis. The larger sample described in the first column consists of all person-months of SIPP individuals aged 35-64 who were employed at the beginning of the month. The smaller sample described in the second column consists of those observations that were actually matched to an LEHD establishment. As can be seen from Table 1, the two samples are very similar in terms of sample means and standard deviations. The variable "separated" is an indicator for whether the individual left his or her job in the calendar month. The separation rate is about 10% smaller in the matched sample, probably due to the inability to match some cases in which an individual reported holding two jobs in a given month. Figure 1 depicts the raw monthly separation rates by single year of age for the full sample and the matched sample. The separation rate increases substantially around age 60, as expected given typical retirement patterns in the U.S. The age pattern is clearly noisier in the matched sample.

We use two alternative measures of the establishment-level age composition of employment: the fraction of workers aged 55-64 and the fraction aged 60-64. We use *establishment*-specific fractions of older workers averaged across all observed quarters for a given establishment[3]. As noted above, we also include in some specifications the *industry*-specific age distribution of employment. This is based on a worker's self-reported three-digit industry, and was computed using the 1990 Census Microdata file in order to obtain large enough samples for each three-digit industry. The mean *establishment*-level fraction of workers aged 55-64 in the matched sample is 0.103 and the corresponding figure for age 60-64 is 0.038,

[3] We estimated models using time-varying quarterly age composition as well, but we prefer the average establishment age structure because it is less volatile, especially for smaller establishments.

and their industry counterparts are very similar. The standard deviations reported in parentheses show that the magnitude of establishment-level variation in the age distribution is much larger than the magnitude of the industry–level variation. This suggests that most of the variation in the establishment-level age composition of employment is within industry. We verified this by using the LEHD data to regress the fraction aged 60-64 in an establishment on a full set of four digit industry dummies, a set of 10 establishment size dummies, and several other establishment characteristics available in the LEHD data (location [county], ownership type, and a multi-plant indicator). The R^2 for this regression was .064, indicating that the great majority of variation in the establishment-level age structure is within-industry.

5. Results

To illustrate the basic patterns of interest, we first estimated a logit hazard model of separation using a set of single-year age dummies, the fraction of 60-64 year old workers at the individual's *establishment*, and interactions of these variables, with no other control variables. Figure 2 depicts the pattern of the predicted monthly separation hazard rate for two different values of the establishment-level fraction of workers aged 60-64: half a standard deviation below the sample mean (0.018) and half a standard deviation above the mean (0.058). The results in Figure 2 suggest that the separation propensity is lower at most ages beginning in the mid 40s when the fraction of the establishment work force aged 60-64 is higher. Next, we added the following set of control variables to the model: gender, race, marital status, education, family income other than the worker's earnings, wealth, self-reported health and disability status, the hourly wage rate, two-digit industry dummies, occupation dummies, class of worker, size of the establishment, job tenure and work experience, pension plan characteristics, health insurance coverage, and region. This specification also controls for the *industry*-level fraction of workers aged 60-64 and its interactions with single-year age dummies, in addition to the establishment-level share aged 60-64 and age interactions. Figure 3 presents the average predicted separation propensity by age based on this specification, for the same two values of the *establishment*-level age composition variable as in Figure 2. The separation rate is predicted for each individual and then averaged at each age. In this specification, a lower separation rate at ages 59-61 in establishments with a larger fraction of older workers is noticeable, but the separation rate is also lower at some younger ages as well when the share aged 60-64 is higher. These results suggest an association between the share of older workers in an establishment and the separation propensity of older workers, but the "difference-in-difference" implied by Figures 2 and 3 is not especially sharp. We anticipate that the much larger samples available in the full LEHD files will make sharper inferences possible in the next version of the paper.

Table 2 provides estimates of the coefficients of interest in a more parsimonious specification, in which dummies for five year age groups are used instead of single year age dummies. The full set of additional controls described above was also included. First, we estimate the model with the *industry*-level fraction of workers aged 60-64 only (specification 1). Since technology differs across industries, we might expect to find that industry-level differences in the age composition of employment are associated with differences in employment behavior of older versus younger workers. This specification does not contain any establishment-level data, so it can be estimated on sample 1 from Table 1: the full sample of SIPP cases that could be potentially matched to the available LEHD extract. The coefficient estimates on the interaction between dummies for workers aged 55-59 and 60-64 (the most common age range of retirement) and the industry fraction aged 60-64 are negative, significantly different from zero, and much larger than the interactions for the younger age groups. This is exactly the pattern we hypothesized, although it is for the *industry*-specific age composition rather than the *establishment*-level age composition.

Next, we estimate exactly the same specification using sample 2: observations that have establishment level data. Comparing specifications 1 and 2 allows us to determine whether the effect of *industry*- level age-structure is sensitive to sample composition. The main results from column 1 are unaffected by the change in the estimation sample.

Specification 3 adds establishment level characteristics *other than* the age distribution, including establishment-average earnings, ownership type, a multi-plant dummy, and total employment (all averaged over all available quarterly observations for a given establishment). Comparing specifications 2 and 3 allows us to investigate whether establishment characteristics other than the age distribution affect the impact of the industry- level age distribution. As can be seen, the results in columns 2 and 3 are very similar.

Specification 4 replaces the *industry*-specific fraction of workers aged 60-64 and its interactions with their *establishment*-level counterparts. The estimated effects of the establishmen*t*-level age composition and age interactions are smaller than those of the industry-level age composition. This is a result of the much larger variance of the establishment- level age60-64 share (.041) compared to the industry-level share (.012), documented in Table 1. In order to provide a useful metric for comparing the effects of the industry and establishment age60-64 shares, consider the impact of a one standard deviation increase in each. In specification 3, a one standard deviation increase in the industry-specific fraction aged 60-64 is predicted to reduce the log odds of separation of a worker aged 60-64 by .09 (5.722*.012 – 13.443*.012). In specification 4, the corresponding increase in the establishment- level fraction aged 60-64 is predicted to reduce the log odds of separation by .11 (1.574*.041 – 4.152*.041).

Thus the impact of the establishment-level measure is slightly larger than the impact of the industry-level measure when they are compared appropriately. The estimates in column 4 are not precise enough to distinguish between the effects of the age60-64 share on older and younger workers; only the difference between ages 35-39 (the reference category) and 60-64 is significantly different from zero. Nevertheless, the pattern of the interaction coefficient estimates is consistent with our prediction: the effects of the establishment-level share of workers aged 60-64 is larger at older ages.

Next, we present estimates from a specification that includes both establishment and industry fractions of workers aged 60-64 and their interactions with age group dummies (specification 5). This is our preferred specification, and coefficient estimates for the full set of control variables for this specification are provided in the Appendix. The main finding here is that the effects of the establishment-level share aged 60-64 are very similar in specifications 4 and 5: controlling for the industry-level fraction aged 60-64 hardly matters. And the effects of the industry-level share aged 60-64 are very similar in specifications 3 and 5. Figure 4 depicts the predicted monthly separation rate by age for the same pair of fraction values used in previous simulations, based on our estimates from this specification. Comparing Figures 3 and 4 shows that the more parsimonious specification of age effects does not distort the age pattern, and Figure 4 illustrates the age structure pattern more clearly than Figure 3. The visual evidence in Figure 4 is clear: workers at older ages have a lower propensity to separate from employers with a greater share of older workers, relative to their younger counterparts. In order to verify this result statistically, we report the following difference-in-difference estimates: the effect of the fraction 60-64 on the log odds of the separation propensity of 60-64 year old workers relative to 55-59 year old workers is –1.46 (–3.76 +2.3), with a p-value of 0.47 for the difference. The corresponding difference-in-difference estimate is 0.048 (p-value of 0.99) using 50-54 year olds as the comparison group, –1.585 (p-value of 0.55) compared to 45-49 year olds, and –2.7 (p-value of 0.19) compared to 40-44 year old workers. These difference-in-difference estimates are mainly not significantly difference from zero, but we expect that this will change when we re-estimate the model using the full LEHD sample.

Finally, we re-estimate the model controlling for three-digit industry fixed effects (specification 6). The industry fixed effects control for all industry-level factors that could be associated with the separation propensity, including observed factors such as the industry-specific age structure used in specification 5, and other unobserved factors. As can be seen, the effects of the *establishment*-level age composition are quite robust.

We perform another specification test by re-estimating our model using a broader definition of 'older workers' – the fraction of workers aged 55-64. Table 3 reports estimates of the coefficients of interest corresponding to specifications 5 and 6 from Table 2. The coefficient estimates on the interaction terms are smaller in Table 3 than in Table 2, but the standard deviation of the establishment-level fraction aged 55-64 is larger (.075 vs. .041 for the age 60-64 share). A one standard deviation increase in the establishment-level share of workers aged 55-64 is predicted to reduce the log odds of separation by .05 (1.147*.075 – 1.872*.075), compared to .11 for the effect of a one standard deviation increase in the fraction aged 60-64 as discussed above. This suggests that the 10 year definition of 'older workers' is too broad. Nevertheless, the major pattern is still noticeable – older workers have a systematically lower probability of separating from establishments with a larger share of older workers, compared to older workers at establishments with a smaller share of older workers.

6. Conclusions

This study presents the first analysis, of which we are aware, of the association between the age structure of employment in an establishment and the propensity of older workers to separate from the establishment. The empirical results show that a larger share of older workers in an establishment is associated with a lower separation propensity of older workers, relative to their younger counterparts. This evidence is consistent with the hypothesis that labor market rigidities, as manifested in the age structure of employment, are an important determinant of employment decisions of older workers. However, this interpretation of the results is admittedly speculative: we have no direct measure of technology-induced labor market rigidities. We argue that the share of older workers at an establishment is a useful proxy for the flexibility of technology at the establishment. We estimated many different specifications in order to verify that the results are robust, and we find that they are. Nevertheless, the results presented here are best viewed as suggestive of the possible importance of labor market rigidities affecting older workers, but clearly not as definitive evidence.

In the next version of this paper, we will extend the analysis of worker separation propensities in several ways. First, we will re-do the analysis using the full sample of SIPP workers that can be matched to employers in the LEHD. The preliminary analysis described above was based on a relatively small sub-sample of LEHD data made available to us for testing purposes. We recently received the full LEHD data files containing a record for every SIPP worker who ever worked for an LEHD establishment during the period covered by the LEHD, along with records for every other worker employed by the establishment. This much larger sample will very likely provide sharper inferences. Second, we will extend the analysis to examine the association between the age structure of *turnover*, *hiring*, and *separations* at an establishment and the worker separation propensity. Third, we will disaggregate the analysis to examine quits and layoffs separately. Fourth, we will disaggregate the analysis to separately examine separations that lead to withdrawal from the labor force and separations that involve a change of employer. Fifth, we will examine other outcomes, including hours of work and wages. Finally, we will compare the effects of the establishment-average age structure (where the average is taken over all available quarters of data) to the effects of quarter-specific age structure.

To conclude, some additional shortcomings of our study are worth mentioning. The approach we use in this paper imposes relatively little structure on the data, but the estimates do not provide an easily interpretable measure of the magnitude of the impact of labor market rigidities on older workers. We reported above that a one standard deviation increase in the share of workers aged 60-64 would result in an 11% decline in the separation propensity of workers aged 60-64. There is no obvious way to interpret the magnitude of this effect in terms of its implications for economic well being. This estimate also doesn't allow us to distinguish between specific sources of labor market rigidities discussed above. Finally, an important point made by Hurd (1996) is that we do not observe the wage and compensation that workers would have had if they had done something different from what they were observed doing. For example, what would the worker have earned if he had reduced his hours of work on the same job instead of remaining at full time hours, or if he remained full time rather than retiring? Establishment-level data by themselves do not overcome this selection bias. Hence, an important area for future research is to specify and estimate structural models that help to address the problems described above, at the cost of additional assumptions. The quantitative analysis of specific sources of labor market rigidities and their effects on employment behavior could be of considerable value in evaluating different types of policy interventions aimed at older workers.

Table 1: Means and Standard Deviations of Selected Sample Characteristics

		Sample 1 (potential matches)	Sample 2 (actual matches)
Age, years		45.72 (7.62)	45.73 (7.58)
Five-year age group fractions	Age 35-39	0.26	0.25
	Age 40-44	0.25	0.25
	Age 45-49	0.19	0.19
	Age 50-54	0.15	0.15
	Age 55-59	0.10	0.11
	Age 60-64	0.05	0.05
Gender, fractions	Males	0.47	0.48
	Females	0.53	0.52
Race, fractions	White	0.82	0.84
	Black	0.06	0.05
	Other	0.12	0.11
Marital status, fractions	Single	0.31	0.30
	Married	0.69	0.70
Education, years		13.52 (3.38)	13.55 (3.31)
Monthly income other than the individual's earnings, $		$1,417 (1,882)	$1,425 (1,845)
Wealth, $ thousands		$103.69 (207.63)	$109.02 (230.76)
Health status, % in good health		93.3%	93.4%
Disabled, %		7.8%	7.6%
Initial experience, years		19.00 (12.09)	18.72 (12.07)
Tenure, months		105.08 (100.09)	106.60 (99.16)
Wage rate, $ per hour		$10.90 (10.00)	$11.31 (10.61)
Health insurance in own name, %		71.4%	75.1%
Employer provided health insurance, %		76.2%	79.3%
Pension plan coverage, %		42.4%	46.1%
Defined benefit pension plans, %		26.9%	28.6%
Industry-specific fraction of 60-64 year old workers		0.042 (0.012)	0.041 (0.011)
Industry-specific fraction of 55-64 year old workers		0.103 (0.026)	0.102 (0.024)
Establishment-specific fraction of 60-64 year old workers			0.038 (0.041)
Establishment-specific fraction of 55-64 year old workers			0.103 (0.075)
Separated		0.0148 (0.121)	0.0133 (0.114)
Number of individuals		12,688	7,581
Number of person-months		252,645	156,307

Note: Dollar amounts are deflated by the Consumer Price Index, base year 1982-84.

Table 2: Selected Coefficient Estimates from Logit Models of Monthly Job Separation, Fraction of Older Workers 60-64 Years Old
(standard errors in parentheses)

	1	2	3	4	5	6
Age40-44 α_1	0.209 (0.137)	0.124 (0.200)	0.147 (0.201)	0.057 (0.088)	0.165 (0.203)	0.043 (0.091)
Age45-49 α_2	0.249 (0.158)	0.286 (0.221)	0.279 (0.222)	0.156 (0.108)	0.312 (0.226)	0.127 (0.109)
Age50-54 α_3	0.269 (0.171)	0.216 (0.240)	0.232 (0.242)	0.312* (0.115)	0.285 (0.243)	0.282** (0.118)
Age55-59 α_4	0.523** (0.210)	0.887* (0.280)	0.860* (0.281)	0.289** (0.125)	0.890* (0.283)	0.268** 0.132
Age60-64 α_5	1.046* (0.235)	1.060* (0.353)	1.041* (0.357)	0.682* (0.147)	1.174* (0.362)	0.635* (0.152)
Industry-specific fraction aged 60-64 β_0	3.218 (2.446)	6.073*** (3.458)	5.722 (3.504)		5.309 (3.541)	
Age40-44**industry*-specific fraction β_1	-4.609 (3.130)	-2.545 (4.607)	-3.166 (4.619)		-2.839 (4.745)	
Age45-49* *industry*-specific fraction β_2	-5.412 (3.578)	-4.873 (5.050)	-4.870 (5.072)		-4.076 (5.204)	
Age50-54* *industry*-specific fraction β_3	-4.863 (3.862)	-0.635 (5.453)	-0.969 (5.510)		0.738 (5.694)	
Age55-59* *industry*-specific fraction β_4	-10.384** (4.661)	-16.944* (6.473)	-16.441** (6.500)		-15.278** (6.657)	
Age60-64* *industry*-specific fraction β_5	-12.754* (4.982)	-13.515*** (7.809)	-13.443*** (7.902)		-12.041 (8.081)	
Establishment-specific fraction aged 60-64 γ				1.574 (1.566)	1.290 (1.585)	0.842 (1.605)
Age40-44**establishment*-specific fraction δ_1				-1.239 (2.123)	-1.060 (2.184)	-0.971 (2.167)
Age45-49**establishment*-specific fraction δ_2				-2.403 (2.654)	-2.175 (2.729)	-2.001 (2.611)
Age50-54**establishment*-specific fraction δ_3				-3.572 (2.652)	-3.808 (2.788)	-2.857 (2.643)
Age55-59**establishment*-specific fraction δ_4				-3.130 (2.177)	-2.300 (2.168)	-3.146 (2.421)
Age60-64**establishment*-specific fraction δ_5				-4.152** (2.092)	-3.760*** (2.106)	-3.082 (2.171)
N(person-months)	252,645	156,307	156,307	156,307	156,307	153,475
N(individuals)	12,688	7,581	7,581	7,581	7,581	7,475

* significant at 1 % level
** significant at 5 % level
*** significant at 10 % level

Table 3: Selected coefficient estimates from logit models of monthly job separation, fraction of older workers 55-64 years old

(standard errors in parentheses)

	7	8
Age40-44 α_1	0.263 (0.249)	0.156 (0.109)
Age45-49 α_2	0.463*** (0.269)	0.180 (0.124)
Age50-54 α_3	0.503*** (0.289)	0.417* (0.132)
Age55-59 α_4	0.660** (0.328)	0.382** (0.152)
Age60-64 α_5	0.995** (0.431)	0.631* (0.175)
Industry-specific fraction aged 55-64 β_0	2.877 (1.799)	
Age40-44**industry*-specific fraction β_1	-1.190 (2.447)	
Age45-49* *industry*-specific fraction β_2	-3.079 (2.602)	
Age50-54* *industry*-specific fraction β_3	-0.612 (2.868)	
Age55-59* *industry*-specific fraction β_4	-2.949 (3.100)	
Age60-64* *industry*-specific fraction β_5	-3.264 (4.074)	
Establishment-specific fraction aged 55-64 γ	1.147 (0.711)	1.036 (0.713)
Age40-44**establishment*-specific fraction δ_1	-1.533 (1.114)	-1.722 (1.090)
Age45-49**establishment*-specific fraction δ_2	-0.930 (1.182)	-1.396 (1.144)
Age50-54**establishment*-specific fraction δ_3	-2.734** (1.179)	-2.5145** (1.143)
Age55-59**establishment*-specific fraction δ_4	-1.930** (0.966)	-2.268** (1.068)
Age60-64**establishment*-specific fraction δ_5	-1.872*** (1.134)	-1.586 (1.155)
N(person-months)	156,307	153,475
N(individuals)	7,581	7,475

* significant at 1 % level
** significant at 5 % level
*** significant at 10 % level

Figure 1
Raw Monthly Separation Rates by Single Year of Age

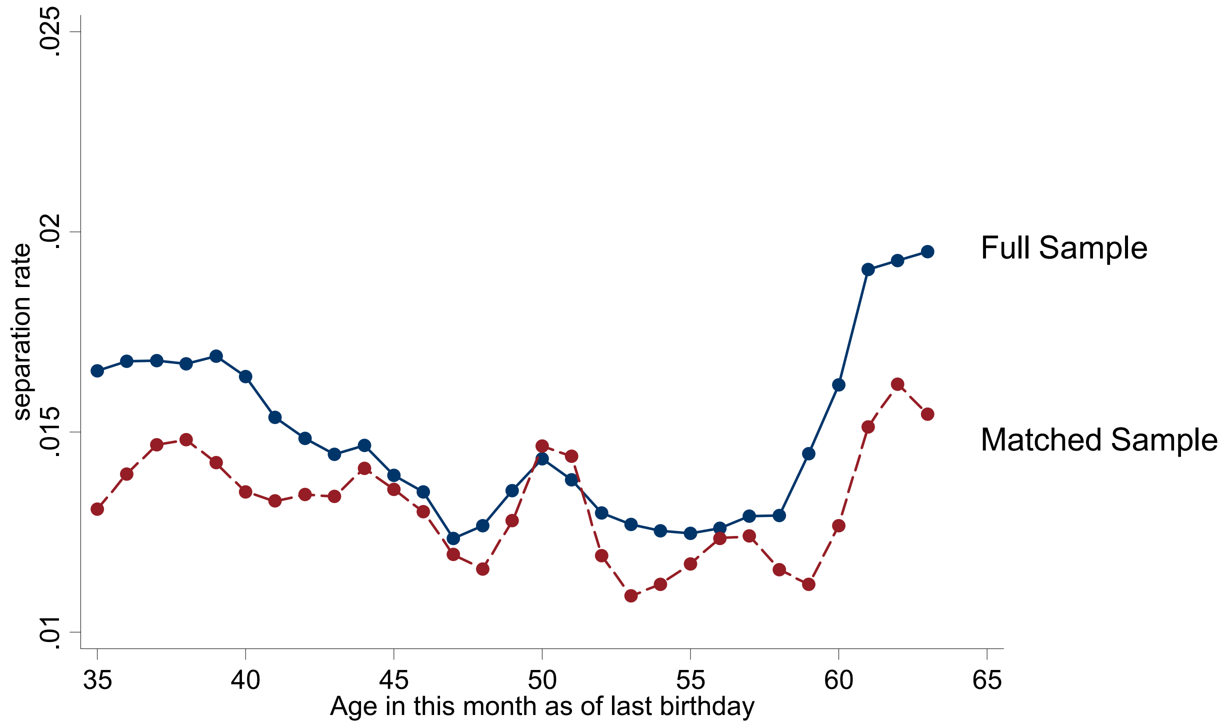


Figure 2
Predicted Monthly Separation Rate
by Single Year of Age and Establishment-Specific Fraction of 60-64 Year Old Workers
and no Other Controls

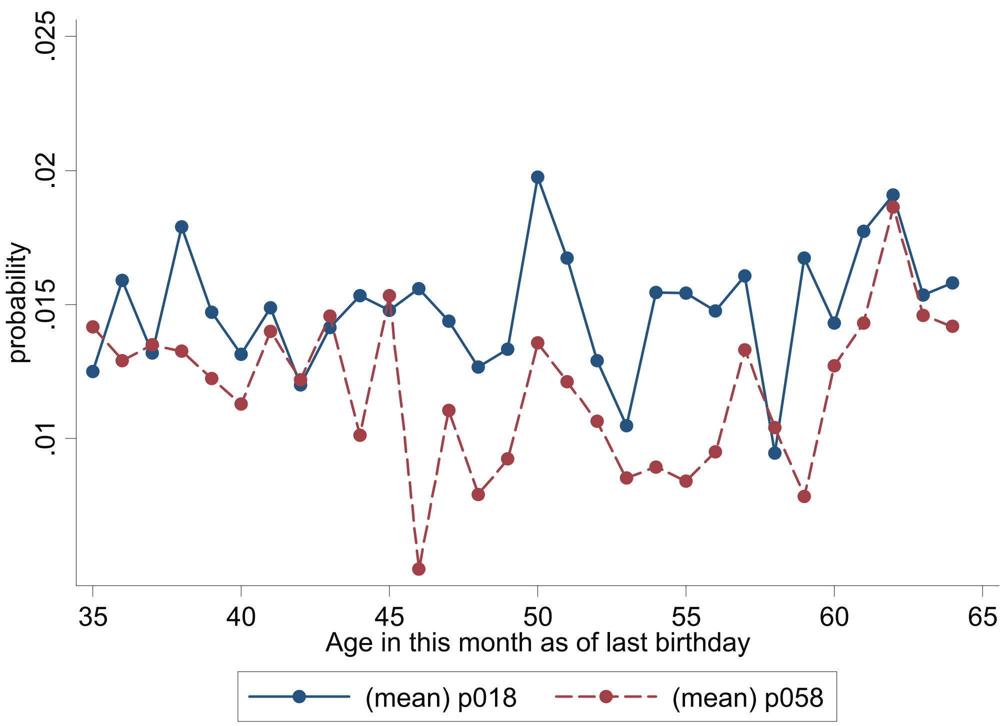


Figure 3
Predicted Monthly Separation Rate
by Single Year of Age and Establishment-Specific Fraction of 60-64 Year Old Workers
with the Full Set of Control Variables

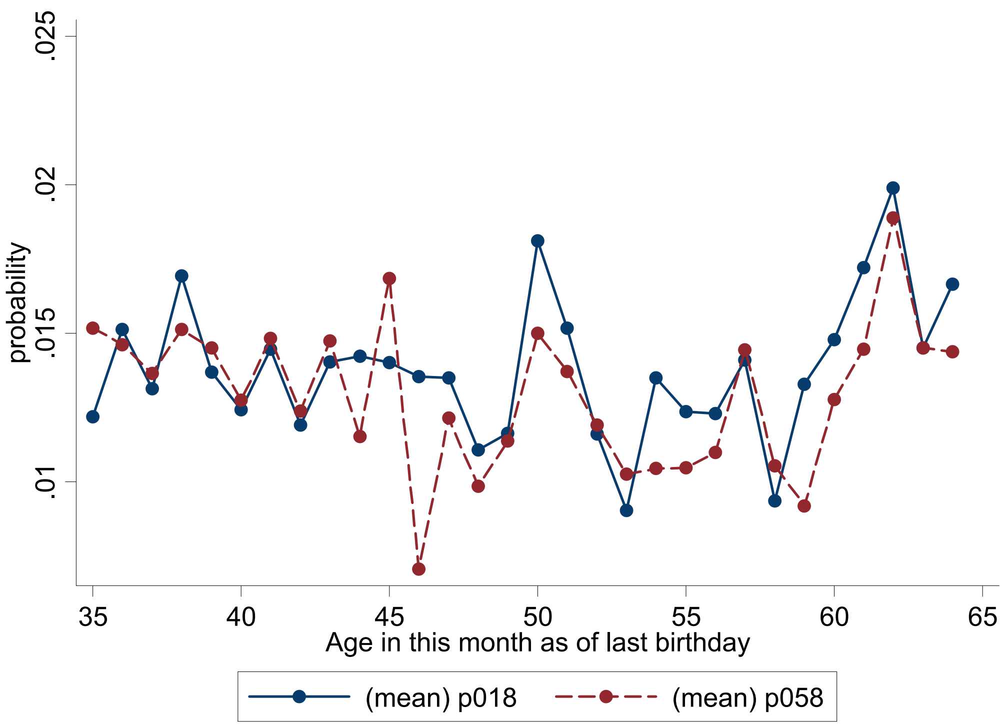
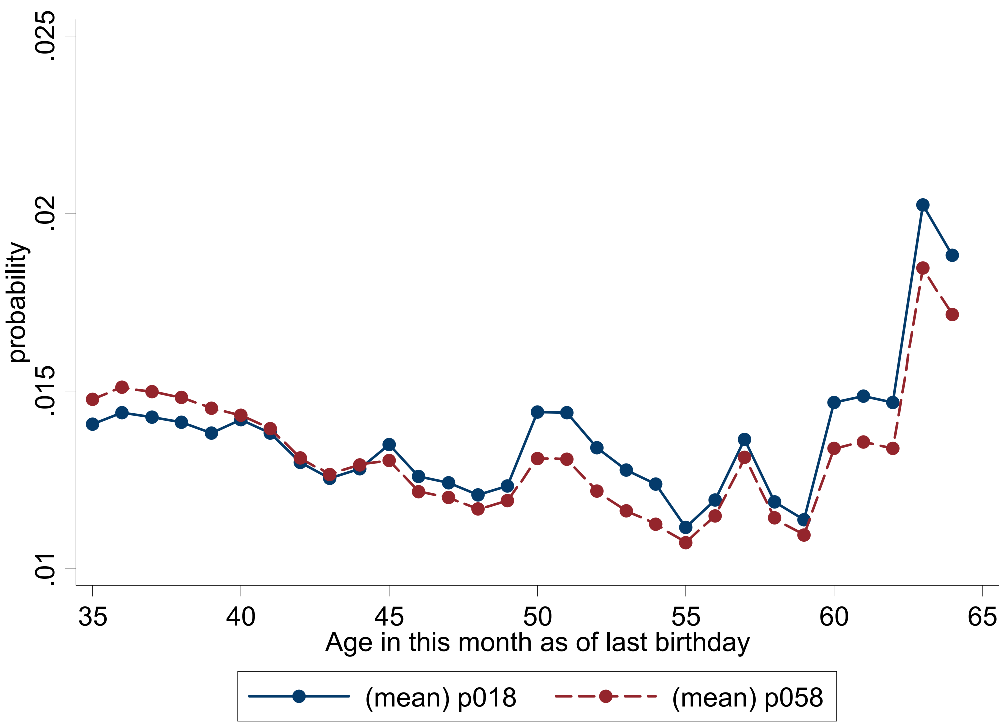


Figure 4
Predicted Monthly Separation Rate
by Age and Establishment-Specific Fraction of 60-64 Year Old Workers
with Five Year Age Group Dummies and the Full Set of Other Control Variables
(from specification 5 in Table 2)



References

Abowd, John M., Bryce E. Stephens, Lars Vilhuber, Fredrik Andersson, Kevin L. McKinney, Marc Roemer, and Simon Woodcock (2006). "The LEHD Infrastructure Files and the Creation of the Quarterly Workforce Indicators," LEHD Technical Paper 2006-01, January.

Abowd, John M., John Haltiwanger, and Julia Lane (2004). "Integrated Longitudinal Employee-Employer Data for the United States," LEHD Technical Paper 2004-02, May.

Abraham, Katherine G. and Susan N. Houseman (2004). "Work and Retirement Plans Among Older Americans," Upjohn Institute for Employment Research Staff Working paper No. 04-105.

Blau, David M. (1994). "Labor Force Dynamics of Older Men," *Econometrica* 62(1): 117-156

Blau, David M. and Donna B. Gilleskie (2001). "The Effect of Health on Employment Transitions of Older Men," in Solomon W. Polacheck (ed.) Worker Wellbeing in a Changing Labor Market, Research in Labor Economics, Volume 20, JAI (Elsevier Science, Amsterdam: 35-66.

Blau, David M. and Donna B. Gilleskie (in press). "Health Insurance and Retirement of Married Couples," *Journal of Applied Econometrics*.

Bound, John (1991). "Self-Reported and Objective Measures of Health in Retirement Models," *Journal of Human Resources* 26, Summer: 106-138.

Burtless, Gary and Moffitt, Robert A. (1985). "The Joint Choice of Retirement Age and Postretirement Hours of Work." *Journal of Labor Economics* 3:209-236.

Doms, Mark, Timothy Dunne, and Kenneth Troske. (1997). "Workers, Wages, and Technology," *Quarterly Journal of Economics* 112: 253-290.

Davis, Steve J. and John Haltiwanger (1999). "Gross Job Flows," in Orley C. Ashenfelter and David Card (eds.) Handbook of Labor Economics, Volume 3B, Amsterdam, Elsevier: 2711-2808.

Friedberg, Leora. (2000). "The Labor Supply Effects of the Social Security Earnings Test," *Review of Economics and Statistics* 82 (1), February: 48-63.

Gustman, Allan L. and Thomas L. Steinmeier (1984). "Partial Retirement and the Analysis of Retirement Behavior," *Industrial and Labor Relations Review* 37(3): 403-415.

Hellerstein, Judith K., David Neumark, and Kenneth R. Troske (1999). "Wages, Productivity, and Worker Characteristics: Evidence From Plant-Level Production Functions and Wage Equations," *Journal of Labor Economics* 17 (3): 409-446.

Hurd, Michael D. (1996). "The Effect of Labor Market Rigidities on the Labor Force Behavior of Older Workers," in David Wise (ed.) Advances in the Economics of Aging, University of Chicago Press for the NBER, Chicago.

Hutchens, Robert M. (1988). "Do Job Opportunities Decline with Age?" *Industrial and Labor Relations Review* 42 (1): 89-99.

Karoly, Lynn A. and Julie Zissimopoulos (2004). "Self-Employment among older U.S. Workers," *Monthly Labor Review*, July: 24-47.

Lazear, Edward (1979). "Why Is There Mandatory Retirement?" *Journal of Political Economy* 87(6): 1261-1284.

LEHD Program (2002). "The Longitudinal Employer-Household Dynamics Program: Employment Dynamics Estimates Project Versions 2.2 and 2.3," LEHD Technical Paper 2002 05 (rev.).

Maestas, Nicole (2004). "Back to Work: Expectations and Realizations of Work After Retirement," RAND Working Paper 196.

Quinn, Joseph (1980). "Labor Force Participation Patterns of Older Self-Employed Workers," *Social Security Bulletin*, April: 17-28.

Ruhm, Christopher J. (1990). "Bridge Jobs and Partial Retirement," *Journal of Labor Economics* 8(4): 482-501.

Rust, John and Christopher Phelan (1997). "How Social Security and Madicare Affect Retirement Behavior in a World of Incomplete Markets," *Econometrica* 65(4), July: 781-832.

Scott, Frank A., Mark C. Berger, and John E. Garen (1995). "Do Health Insurance and Pension Costs Reduce the Job Opportunities of Older Workers?" *Industrial and Labor Relations Review* 48 (4): 775-791.

Van der Klaauw, Wilbert and Kenneth I. Wolpin (2005). "Social Security and the Retirement and Savings and Retirement Behavior of Low Income Households," Working Paper, Department of Economics, University of North Carolina at Chapel Hill.

Appendix

Logit parameter estimates of the monthly job separation hazard

	Coefficient	Robust Std. Err.
age40_44	0.165	0.203
age45_49	0.312	0.226
age50_54	0.285	0.243
age55_59	0.890	0.283
age60_64	1.174	0.362
Industry fraction 60-64	5.309	3.541
age40_44 * industry fraction 60-64	-2.839	4.745
age45_49 * industry fraction 60-64	-4.076	5.204
age50_54 * industry fraction 60-64	0.738	5.694
age55_59 * industry fraction 60-64	-15.278	6.657
age60_64 * industry fraction 60-64	-12.041	8.081
Establishment fraction 60-64	1.290	1.585
age40_44 * establishment fraction 60-64	-1.060	2.184
age45_49 * establishment fraction 60-64	-2.175	2.729
age50_54 * establishment fraction 60-64	-3.808	2.788
age55_59 * establishment fraction 60-64	-2.300	2.168
age60_64 * establishment fraction 60-64	-3.760	2.106
Male	-0.101	0.057
Black	-0.241	0.107
American Indian	-0.037	0.161
Asian	-0.061	0.078
Married, Spouse Absent	0.388	0.168
Widowed	-0.216	0.178
Divorced	0.159	0.068
Separated	0.113	0.115
Never married	0.175	0.079
Education	0.019	0.009
Real income of other household members	-0.607	1.547
Total household wealth	-0.021	0.013
Indicator: Wealth imputed	-0.573	0.155
Real wage	0.002	0.001
Indicator: Wage imputed	1.596	0.194
Tenure	-0.007	0.001
Tenure squared	0.000	2.52e-06
First quarter of tenure	0.136	0.068
First year of tenure	0.270	0.085
Year 2-5 of tenure	0.012	0.076
Initial experience	-0.005	0.003
Indicator: Experience imputed	0.022	0.111
Pension plan indicator	-0.461	0.180
DB pension plan indicator	0.255	0.152
Employer contributions indicator	-0.181	0.150

Indicator: Pension information imputed	1.76	0.067
Establishment size <= 5 workers	-0.218	0.113
Establishment size 6-10 workers	-0.095	0.124
Establishment size 11-25 workers	-0.227	0.097
Establishment size 26-50 workers	-0.048	0.092
Establishment size 51-75 workers	-0.125	0.117
Establishment size 76-100 workers	-0.106	0.113
Establishment size 101-200 workers	0.010	0.092
Establishment size 201-500 workers	-0.027	0.081
Establishment size 500-1000 workers	-0.171	0.094
Local government establishment	-0.088	0.420
Private sector establishment	-0.107	0.413
Multi-plant dummy	0.043	0.063
Average earnings at establishment	-0.912	0.697
Average number of workers	-0.085	0.033
Disabled	0.300	0.071
Bad health	0.094	0.078
Indicator: Self-reported health imputed	-0.423	0.104
Health insurance, own name	-0.149	0.082
Health insurance, others name	0.116	0.068
Employer provided health insurance	-0.481	0.077
Midwest	0.155	0.396
South	0.030	0.359
West	-0.379	0.316
Metropolitan area	-0.043	0.078
Time trend	0.004	0.000
Constant	-99.353	13.004
Industry:		
Mining	0.086	0.391
Construction	0.310	0.207
Non-durables	0.197	0.207
Durables	0.313	0.202
Transportation	-0.067	0.241
Public utilities	0.406	0.229
Wholesale trade	0.194	0.208
Retail trade	-0.061	0.203
Finance	0.070	0.206
Repair services	0.430	0.205
Personal services	-0.008	0.220
Recreation services	0.179	0.246
Health services	-0.186	0.213
Educational services	0.039	0.241
Other services	0.221	0.206
Public administration	0.146	0.270
Occupation:		
Executives	0.001	0.086
Professionals	0.025	0.127
Technicians	-0.105	0.098
Sales	-0.002	0.080

Administrative support	-0.458	0.412
Private household	-0.073	0.208
Protective service	-0.135	0.104
Farming, forestry and fishing	0.334	0.191
Craft and repair	0.030	0.096
Machine operators	-0.085	0.115
Transportation and material moving	-0.036	0.142
Handlers, helpers, and laborers	-0.065	0.143
Class of worker:		
private non-profit	-0.357	0.118
federal government	-0.360	0.189
state government	-0.430	0.249
local government	-0.353	0.202
armed forces	1.135	0.625
family business	-0.881	1.270