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ABSTRACT

Labor Market Rigidities and the Employment Behavior of Older Workers*

The labor market is often asserted to be characterized by rigidities that make it difficult for older workers to carry out their desired trajectory from work to retirement. An important source of rigidity is restrictions on hours of work imposed by firms that use team production or face high fixed costs of employment. Such rigidities are difficult to measure directly. We develop a model of the labor market in which technological rigidity affects the age structure of a firm's work force in equilibrium. Firms using relatively flexible technology care only about total hours of labor input, but not hours of work per worker. Older workers with a desire for short or flexible hours of work are attracted to such firms. Firms using a more rigid technology involving team production impose a minimum hours constraint, and as a result tend to have a younger age structure. A testable hypothesis of the model is that the hazard of separation of older workers is lower in firms with an older age structure. We use matched worker-firm data to test this hypothesis, and find support for it. Specification tests and alternative proxies for labor market rigidity support our interpretation of the effect of firm age structure on the separation propensity. These results provide indirect but suggestive evidence of the importance of labor market rigidities.

JEL Classification: J26, J23

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1 INTRODUCTION

The majority of workers retire by moving directly from full-time year-round employment on a long term job to non-employment. Gradual retirement, partial retirement, bridge jobs, and other less abrupt transitions to retirement are common, but much less frequent than abrupt and complete exit from employment. Why is retirement typically so abrupt? If most individuals retire as a result of a health shock, then the prevalence of abrupt retirement would be understandable. Deterioration in health is in fact an important cause of retirement, but most changes in employment status at older ages are not associated with a decline in self-reported health (we document this below). In the absence of a health shock, it seems implausible that preferences for leisure would change abruptly at older ages. One indication that abrupt changes in preferences are unlikely to be a major cause of abrupt retirement is that self-employed workers, who have much more discretion over their hours of work than do wage-salary workers, are much more likely to retire gradually (also documented below).

An alternative explanation for the prevalence of abrupt retirement is labor market rigidity. The labor market has been 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 often cited include lack of opportunity for part-time and flexible-hours work; 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). Labor market rigidities may limit the employment options of workers of all ages, but older workers will be more affected by rigidities if they have a stronger desire for leisure or flexible work hours than younger workers.

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". Many factors could be responsible for making the labor market rigid. For example, older workers face discontinuities in retirement incentives as a result of government policy and labor market institutions. Social Security and Medicare have strictly defined age eligibility criteria that affect employment incentives, particularly for workers who are liquidity-constrained (Rust and Phelan, 1997). 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 (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 firm that provides the benefits, before the normal retirement age in the pension plan. In addition, most DB plans link benefits to earnings in the last few years on the job, making it costly for a worker to decrease work hours at the career employer. 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 as we document below abrupt retirement is the most common pattern even for individuals who don't appear to 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, some features of technology could induce firms to impose constraints on hours of work. If there are fixed costs of hiring, training, and employing a worker, then firms may impose a minimum hours constraint on their workers (Hamermesh, 1993). If production takes place in teams, then the absence of a team member could reduce team productivity. In this case firms might require the presence of workers at specific times, reducing the flexibility of workers in scheduling their hours of work. Other factors could result in reluctance of firms to hire older workers under the same terms as younger workers, but would not result in hours restrictions placed on older workers who "age in place." For example, 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). The short expected duration of future employment of an older worker reduces the incentive of a firm to train and promote older workers, despite the fact that some older workers may plan to remain employed for a long time (Hutchens, 1988).

Some of these sources of labor demand rigidities are caused by features of the technology of production that may affect all of a firm's workers, not just older workers. But if the preference for leisure increases with age, then the preferences of older workers differ systematically from those of younger workers. Thus the existence of technology-induced rigidities could be manifested in the age structure of a firm's work force: the more important are technology-induced rigidities, the lower is the share of older workers at a firm. There is evidence that production technology differs substantially across firms, even within narrowly defined industries (Doms, Dunne, and Troske, 1997). These differences are hypothesized to arise from variation across firms in managerial ability, expectations of future price and technological change, and past investment decisions (Davis and Haltiwanger, 1999). Thus, while it is difficult to measure technology directly, 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 firms.

It is important to study 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 contributors to claimants on 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. Most research on employment of older workers adopts a labor supply framework of analysis. It is important to broaden the study of employment at older ages to consider the labor demand side as well.

In this paper, we develop a simple model of the labor market in which, in equilibrium, older workers are disproportionately concentrated in firms with flexible technology, allowing older workers to reduce hours of work gradually. Firms with more rigid technology require longer hours of work, and pay a wage premium to attract workers. Workers at such firms cannot gradually reduce hours of work as they age. In this model, a firm’s age structure is an indicator of its technology: a relatively large share of older workers indicates flexible technology. Although technology itself is not observable, we develop a testable hypothesis from the model: the hazard of separation by an older worker is lower in firms with an older age structure than in firms with a younger age structure.

Empirically, we study the effect of the firm-level age composition of employment on the separation propensity of older workers. We use data on workers from the Survey of Program Participation (SIPP) matched to data on their employers from the Longitudinal Employer-Household Dynamics (LEHD) files (Abowd, Haltiwanger, and Lane, 2004). We use a difference-in-difference approach to analysis, comparing the job exit behavior of older and younger workers in firms with different shares of older workers. Comparing older and younger workers makes it possible to determine whether labor market rigidities affect older workers disproportionately, as we hypothesize. In order to ensure that the firm'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. We also control for firm characteristics, including the number of workers, average earnings, and number of plants. The empirical results show that an older age structure of the work force at the firm level is associated with a lower separation propensity of older workers. This finding supports our hypothesis, and is robust to many specification checks. We explore a number of alternative explanations for the finding, and find little evidence to support them.

The next section of the paper discusses evidence on labor market rigidities and the age structure of firms. Section 3 describes the theoretical framework for the analysis. Description of the data and methodology are provided in Section 4. Section 5 presents the basic estimation results, and section 6 discusses alternative estimates and interpretations. Section 7 concludes.

2 BACKGROUND AND LITERATURE

If tastes for leisure or demand for time in home production increase gradually at older ages, then workers might prefer to gradually reduce hours of work or partially retire as they age, rather than working a fixed number of hours and then abruptly and 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). However, the majority of workers retire by moving directly from full-time employment to complete retirement.

To illustrate, Table 1 shows employment transition rates computed from the Health and Retirement Study (HRS) for individuals aged 51-72 who were employed full-time year-round on a long-tenure job (at least five years) in any of the first five survey waves. The first row of the table shows that 17.7% of these individuals were not employed as of the next survey wave (two years later on average). In comparison, 4.0% were employed on a new year-round full-time job, 5.9% were employed part-time or part-year with the same employer, and 2.7% were employed part-time or part-year with a new employer. Thus of the total of 30.3% who changed employment status between survey waves, the majority (58.4%) made a complete exit from employment.[2]

As mentioned above, deterioration in health is clearly a major cause of retirement, but most changes in employment status at older ages are not associated with a decline in health. Table 1 shows wave-to-wave employment transition rates by the associated wave-to-wave change in self-reported health status. The exit rate from employment conditional on health declining from “good” to “bad” is twice as large as the exit rate conditional on remaining in good health. However, comparing the sample sizes in the last column, it is clear that most exits from employment are not associated with a decline in self-reported health. Thus 67.7% of exits from employment were by individuals whose health remained good, compared to only 13.7% whose

[2] Employment status is defined here by the five mutually exclusive and exhaustive categories shown in Table 1. The HRS is a biannual survey, but a detailed employment history is collected, so it is possible to compute transition rates for shorter intervals. Annual transition rates show a very similar pattern.

health declined from good to bad. Other measures of health show similar results (Blau and Gilleskie, 2001).

As noted above, self-employed individuals are much more likely to retire gradually than are otherwise similar wage-salary employees. 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. 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. The data in Table 1 show that 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 7.3% for wage-salary workers and 16.5% for the self-employed. This clearly suggests that wage-salary workers face a constraint on hours of work imposed by their employers.[3]

Finally, as noted above, workers who are liquidity-constrained, are covered by DB pensions, and do not have Employer-Provided Retiree Health Insurance (EPRHI) are more likely to face incentives to avoid gradual reduction in hours of work. The last three rows in Table 1 show that workers who are covered by DB pensions, do not have EPRHI, and are in the lower half of the distribution of net worth are in fact more likely to retire abruptly (70.8% of all changes in employment status) than are workers who are not covered by a DB pension, do have EPRHI, and are in the upper half of the distribution of net worth (55.6%). Nevertheless, even among workers who, by these criteria, are relatively unlikely to face institutional or liquidity constraints on hours of work, the majority retire abruptly.

[3] Workers with a strong preference for gradual retirement might be more likely to move to self-employment at older ages in order to indulge this preference. However, Karoly and Zissimopoulos (2004) report that only about one third of older self-employed workers entered self-employment after age 50.

Direct evidence on technological sources of labor market rigidity is scarce. When asked 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 (2005) report that 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, but the former are only about half as likely as the latter to actually follow through on their plans.

Hutchens and Grace-Martin (2006) surveyed 950 establishments with at least 20 employees and two white collar employees aged 55+, and posed questions about phased retirement policy. 82% of the establishments report that they have a phased retirement policy. Most of the policies described by the establishments were informal and discretionary, and fewer than half of the establishments with a phased retirement policy reported that any older employees had actually shifted from full time to part time work in the three years prior to the survey. The survey did not inquire about any conditions that might be associated with phased retirement, such as wage cuts and pension eligibility.

Several studies have examined the implications of technological change for older workers. For example, Aubert, Caroli, and Roger (2006) report that adoption of new technology by French firms tended to reduce hiring of older workers but did not affect the age pattern of exit. Friedberg (2003) found that older computer users retired later than older non-users on average, suggesting that computer technology is not a specific cause of labor market rigidity for older workers. Bartel and Sicherman (1993) report that an unanticipated acceleration in the rate of technological change in an industry is associated with a higher exit rate from employment by older workers. These studies are quite interesting, but do not directly address the issue of labor market rigidity.

3 A CONCEPTUAL FRAMEWORK

We illustrate the logic of our conceptualization of technological sources of labor market rigidity and their impact on the employment behavior of older workers using a two-sector general equilibrium model of the labor market.[4] The idea of the model is simple: there are two sectors of the labor market, differentiated by the technology employed. One sector has a "flexible" technology: firms care about the total number of labor hours employed, but are indifferent to the number of hours worked by individual workers. Workers in this sector can reduce hours of work as they age, if they so desire. The other sector has a "rigid" technology: firms in this sector care about hours of work per worker, and as a result they impose a minimum hours-of-work constraint. Workers employed in this sector cannot reduce their hours of work as they grow older unless they shift to an employer in the "flexible" sector or withdraw from the labor force. There is no direct cost of changing sectors, but in equilibrium firms in the rigid sector pays a higher wage than in the flexible sector in order to induce workers to work the number of hours demanded. Thus, leaving the rigid sector does entail an opportunity cost. There are many homogeneous firms in each sector, and the type of technology employed by a firm is fixed.

If the preference for leisure increases with age, some workers who preferred the high-wage rigid sector while young will shift to the lower-wage flexible sector when they are older.

[4] See Hutchens and Grace-Martin (2006) for a related partial equilibrium model, based on fixed costs of employment to employers. Hamermesh and Donald (2007) propose a model of fixed time costs of employment to workers as an explanation for abrupt retirement. Both of these models may help explain abrupt retirement, but neither model has implications for the age structure of employment at the firm level. Hence our approach is complementary with these alternative models.

Thus the flexible sector will have a higher share of older workers than the rigid sector. Workers who experience an increase in the preference for leisure can reduce hours of work without leaving their employer if they are in the flexible sector, but not if they are in the rigid sector. So the hazard rate of exit from a firm will be higher for firms with a younger age structure.

A formal model is presented in the Appendix, using a specific parameterization. There, we show that an equilibrium in which the share of older workers is greater in the flexible sector than in the rigid sector exists.[5] We also show that the separation hazard of older workers is higher in the rigid sector than in the flexible sector as long as preferences for leisure increase rapidly enough with age. We cannot test the first prediction, because we do not have data on technology, so we focus in the empirical section on testing the second prediction: older workers in firms with a larger share of older workers have a lower hazard rate of exit from the firm.

An important question is whether there are other mechanisms that, even in the absence of rigid technology, would result in an association between the age structure of a firm's workforce and the exit rate of its workers. If so, this would limit our ability to draw inferences about labor market rigidity based on the empirical association between age structure and turnover propensity. In a one-sector version of our model in which there is no rigid-technology sector, there is no association between a firm's age structure and the exit propensity of its workers. However, this does not rule out the possibility that other types of models could yield such an association. For example, suppose the age profile of wages is steeper in some sectors than in others. If older workers are concentrated in sectors in which the relative wage rate of older workers is high, this could lead to a lower exit rate of older workers from firms with a higher share of older workers. This suggests that it is important to control for a worker's wage rate in a model of separation. We

[5] Other types of equilibrium exist as well, depending on specific parameter values. The equilibrium described in the text is the one of interest for our purposes.

do this and we also control for average earnings at the worker's firm.

Alternatively, suppose workers prefer to work with co-workers of the same age group. Then an older worker who, by chance or design, finds himself in a firm with a large share of older workers might be less inclined to separate from the firm than an older worker in a firm with a smaller share of older workers. This would yield an association between age structure and separation propensity that has nothing to do with technology-based rigidity.[6] To deal with this and other possible sources of association between age structure and separation propensity that are unrelated to technology, we use an alternative proxy for a firm's technological flexibility: the share of female workers under the age of 30 in the firm's workforce. Women under the age of 30 are in their prime childbearing years, and are much more likely to occupy part-time and flexible-hours jobs than are other workers[7]. If older workers in firms with a larger share of young female workers are less likely to separate from the firm than older workers in firms with a smaller share of young female workers, this would be hard to explain by mechanisms other than technology-based hours inflexibility.

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,

[6] Leonard and Levine (2006) studied employee turnover in 800 workplaces owned and operated by a single firm. They focus on the effects of workplace diversity along the dimensions of age, race, and gender, and do not estimate the effect of the share of older workers on turnover. Their results indicate that a change in age diversity at a workplace had no effect on turnover.

[7] Tabulations from the March 2005 Current Population Survey show that 21% of working women under age 30 worked 20 or fewer hours per week, compared to 9% of other workers; and 18% of working women under age 30 worked 21-34 hours per week, compared to 11% of other workers.

the wage offer, net worth, potential Social Security and pension benefits, and health insurance coverage (Rust and Phelan, 1997; Blau and Gilleskie, 2006; Van der Klaauw and Wolpin, 2005).[8] We augment this list with a measure of the age composition of employment at the individual's firm. As noted above, taking the difference between the employment behavior of older and younger workers makes it possible to determine whether labor market rigidities disproportionately affect older workers. A simple illustration of our empirical specification is

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

where $S_{ijt} = 1$ if individual *i* employed at firm *j* at the beginning of period *t* separates from the firm during period *t*, and equals 0 otherwise; *X* is a vector of individual and firm characteristics; $A_{it} = 1$ if the individual is classified as an older worker in period *t*; and R_{ij} is the proportion of older workers in the work force of firm *j*. This is a hazard model of the risk of separation, and is estimated as a logit.

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. The main effect of age on employment behavior is captured by α. The main effect of the age composition of the firm'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, firms with relatively few older workers may tend to be younger, and firm age may affect the separation propensity of all workers at the firm. The interaction effect δ captures any differences in the effects of the firm's age composition on older workers relative to younger workers. Controlling for pension and health insurance coverage, occupation, industry, and the wage rate (all included in *X*), we interpret differential

[8] It is straightforward to include the wage rate, because the sample consists of workers. We cannot easily include pension and Social Security benefits, since these are observed only for individuals who begin to collect benefits during the sample period. Thus, we omit these benefits and re-interpret our specification as a quasi-reduced form.

effects of a firm's workforce age composition on older versus younger workers as an indication that labor market rigidities affect the employment decisions of older workers differentially.

4.2 DATA

We merge data on individuals from the U.S. Survey of Income and Program Participation (SIPP), 1990 – 2001 panels, with 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 unit of analysis is a person-month. We focus on workers aged 45-69. We exclude younger workers because their behavior is likely to be influenced by factors such as human capital investment, family formation, and so forth that are not relevant for older workers. Thus we compare the behavior of workers in the typical age range of retirement (late 50s to late 60s) to the behavior of mature workers who are not yet approaching typical retirement ages (45-mid 50s).

The LEHD Infrastructure File system is based on state Unemployment Insurance (UI) administrative files, with data available to us from 31 states covering about 80% of the U.S. work force 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. If a firm owns several establishments in a given state, all of these establishments would constitute a single employer. If a firm owns establishments in several states, its establishments in one state are a different employer in the LEHD data than its establishments in another state. This reflects the fact that UI is administered and largely financed by states. Thus an employer in this context is in general neither a firm nor an establishment. The data include the number of establishments per employer. UI covers 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 employer. These data are merged by the Census Bureau 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). The Social Security numbers are then replaced by a scrambled worker identification number, to protect confidentiality. Additional employer information such as industry, location, and ownership type is merged in from the Employer Characteristics Infrastructure Files. 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 provided us with an extract of the LEHD data, containing data for all the workers surveyed in the 1990 – 2001 SIPP panels who appeared in any LEHD record. For a given SIPP sample member, the LEHD file contains a record for *every* available quarter for *every* employer that paid any UI-covered earnings to 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 employer in a given quarter contains a stable firm identifier, the employer characteristics described above, and earnings and basic demographic data on the SIPP worker *and on all other workers who were paid any UI-covered earnings by the employer in that quarter*. Thus we have a census of the entire workforce of a given employer in a given quarter, which allows us to construct measures of the age distribution of the firm's workforce.

We match SIPP and LEHD records as follows. If an 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[9]. 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 employer identifier does not change over time.

[9] 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.

Table 2 presents summary statistics for two samples used in our analysis. The larger sample described in the first column contains SIPP individuals aged 45-69 who were employed at the beginning of a given month and who resided or reported working in one of the LEHD-covered states. The smaller sample described in the second column consists of those observations from the first column that were actually matched to an LEHD firm. The percentage of all SIPP person-months in our sample that is matched to an LEHD record is 52%. Failure to match occurs 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, 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 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.

As can be seen from Table 2, the two samples are very similar in terms of sample means and standard deviations. The variable "separated this month" is an indicator for whether the individual left his or her job in the calendar month. This is the main dependent variable in our analysis. The mean separation rate is about 20% smaller in the matched sample. This is likely due to the fact that it is more difficult to match short and unstable jobs, for all of the reasons discussed in the previous paragraph. Figure 1 depicts the monthly separation rate by single year of age for the samples of potential and actual matches. The separation rate increases noticeably beginning around age 57, and there are large spikes at ages 62 and 65, as expected given typical retirement patterns in the U.S.

The key explanatory variable in our analysis is the employer-specific fraction of workers aged 65-69. This is our preferred proxy for the flexibility of the technology used by the employer. We experimented with proxies based on other age ranges (60-64, 60-69) and found that the results were generally similar but less precise than with our preferred measure. We use the employer-specific fraction of older workers averaged across all observed quarters for a given employer. This provides a relatively stable measure of age structure that is not subject to transitory quarter-to-quarter variation. We also control for the industry-specific age distribution of employment. We compute the industry-specific fraction of older workers using the 1990 Census Microdata file, rather than the SIPP data, in order to obtain large enough samples for each three-digit industry. We merge the industry-specific age composition to the estimation sample based on a worker's self-reported three-digit industry in the SIPP. The mean employer-level fraction of workers aged 65-69 in the matched sample is 0.018, and the mean of the industry-level fraction is also 0.018. The standard deviation of the employer-level fraction is more than three times larger than the standard deviation of the industry-level fraction. This suggests that much of the variation in the employer-level age composition is within industry. We verified this by using the LEHD data to regress the fraction of workers aged 65-69 in a firm on a full set of four-digit Standard Industrial Classification (SIC) industry dummies, a set of 10 firm size dummies, and the other firm characteristics available in the LEHD. The R^2 for this regression is 0.074, indicating that the most of the variation in the employer-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 65-69 year old workers (abbreviated as share65-69 henceforth) at the individual's employer, 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 share65-69: half a standard deviation below the sample mean (0.003) and half a standard deviation above the mean (0.033). The separation rate is predicted for each individual and then averaged at each age. The results in Figure 2 suggest that the separation propensity of workers is lower at ages 62-65 when the share65-69 is higher. Differences at younger ages are small. This is the pattern predicted by our model.

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, job tenure and work experience, pension plan characteristics, health insurance coverage, size of the employer (number of workers), the demographic characteristics and earnings distribution of the employer's workforce, accession rate, ownership type, a multi-plant indicator, the employer's age, region and time. This specification also controls for the industry-level share65-69 and its interactions with single-year age dummies, in addition to the employer-level share65-69 and age interactions. Figure 3 presents the average predicted separation propensity by age based on this specification, for the same two values of the employer-level age composition variable as in Figure 2. In this specification, a lower separation rate at ages 62-65 in employers with a larger share65-69 remains noticeable even after controlling for many other factors that are likely to influence employment behavior. These results suggest an association between the share of older workers in a firm and the separation propensity of older workers.

Table 3 provides estimates of the main 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 omitted age category is 45-49). First, we estimate the model with the industry-level share65-69 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 contains all of the worker characteristics described above, but omits all employer characteristics, so it can be estimated on sample 1 from Table 2: the sample of SIPP person-months that could be potentially matched to the available LEHD extract. The coefficient estimates on the interaction between dummies for workers aged 55-59, 60-64, and 65-69 (the most common age ranges of retirement) and the industry fraction aged 65-69 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 employer-level age composition.

Next, we estimate exactly the same specification using sample 2: observations that were matched to the LEHD. 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.[10] Specification 3 adds all of the employer characteristics described above *other than* the age distribution (each averaged over all available quarterly observations for a given employer). Comparing specifications 2 and 3 allows

[10] The coefficient estimate on the age50-54 interaction becomes positive and significantly different from zero. This is not predicted by our theory, but is not inconsistent with the theory.

us to investigate whether employer 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 share65-69 and its interactions, with their employer-level counterparts. The estimated effects of the employer-level age composition and age interactions are smaller than those of the industry-level age composition. This is partly a result of the larger standard deviation of the firm-level share65-69 (0.030) compared to the industry-level share (0.009), documented in Table 1. In order to provide a useful metric for comparing the effects of the industry and firm age65-69 shares, consider the impact of a one standard deviation increase in each. In specification 3, a one standard deviation increase in the industry-specific share65-69 is predicted to reduce the log odds of separation of a worker aged 60-64 by 0.13 (0.221 * 0.009 – 13.097 * 0.009). In specification 4, the corresponding increase in the firm-level share65-69 is predicted to reduce the log odds of separation by 0.08 (1.402 * 0.030 – 3.998 * 0.030). Thus, the impact of the employer-level measure is smaller than the impact of the industry-level measure when they are compared appropriately.[11] The estimates in column 4 are not precise enough to distinguish between the effects of the share65-69 on each of the five-year age groups; only the difference between ages 45-49 (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 negative effect of the share65-69 is larger at older ages.

Next, we present estimates from a specification that includes both the employer and industry share65-69 and their interactions with age group dummies (specification 5). The coefficient estimates for the full set of control variables for this specification are provided in the

[11] The industry age structure will pick up the effects of any variables that vary within two-digit industry and are correlated with the age structure. This may explain why the industry age structure seems to have a large impact than the employer age structure.

Appendix Table. The main finding here is that the effects of the employer-level share65-69 are very similar in specifications 4 and 5. A one standard deviation increase in the employer-specific share65-69 is predicted to reduce the log odds of separation of a worker aged 60-64 by 0.06 (1.362 * 0.030 – 3.311 * 0.030). Controlling for the industry-level share65-69 hardly matters and the effects of the industry-level share65-69 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 the 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 substantially distort the age pattern. Figure 4 shows that workers aged 60-64 have a lower propensity to separate from employers with a greater share65-69, relative to their younger counterparts. The coefficient estimate on the interaction at ages 60-64 is significantly different from zero, although it is not significantly different from the interactions for the other age groups.

Finally, we re-estimated the model controlling for three-digit industry fixed effects (specification 6) instead of the two-digit industry fixed effects used in previous specifications. 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 employer-level age composition are quite robust.

6 ALTERNATIVE OUTCOMES AND SPECIFICATIONS

The longitudinal structure of the SIPP data allows us to examine the destination of job separations. We define a separation as leading to a change of employers if the respondent starts a new job within 30 days after separating from the previous employer. The residual category of no change of employer pools separations resulting in unemployment, withdrawal from the labor force, and separations of undetermined destination. Slightly more than 20% of monthly job separations are followed by a change of employers within 30 days of the separation (see Table 2). The first two columns in Table 4 presents selected estimates from a multinomial logit model in which the outcomes are (1) separate and change employers, (2) separate without starting a new job within 30 days, and (3) no separation (the reference category). The results are striking: a larger share65-69 reduces separations by older workers to non-employment, but has no impact on employer-to-employer separations. Many of the separations to non-employment are retirements, suggesting that a more flexible technology allows workers to retire gradually on the job rather than switching employers.

Another way to disaggregate separations is by the proximate cause: employer-initiated (laid off, fired, plant closed) versus worker-initiated (quit, retired). The second pair of columns in Table 4 contains estimates of a multinomial logit model in which the outcomes are (1) employer-initiated separation, (2) worker-initiated separation, and (3) no separation. A larger share65-69 reduces the likelihood of an employer-initiated separation by older workers, but has a much smaller effect on the likelihood of an older-worker-initiated separation. If voluntary separations are more likely to result in a change of employer within 30 days, then these results would be consistent with those in the first pair of columns. In contrast, the industry-specific share65-69 has a large negative impact on worker-initiated separations.

A potential problem with using the fraction of older workers as a proxy for technology flexibility is that the age structure of an employer's work force may be determined in part by the age structure of separations from the employer. This could result in spurious correlation between the age structure of employment and the separation propensity of older workers. We avoid transitory sources of spurious correlation by using the average age structure of the employer rather than age structure in the current quarter. Nevertheless, if there are persistent unobserved differences across employers, the long run average age structure of an employer's workforce could be influenced by worker turnover. As an alternative specification we use the fraction of female workers less than 30 years old in the employer's workforce as a proxy for technology flexibility. Firms with flexible technology will attract any type of worker who might value flexible hours of work and part-time schedules, in particular females of childbearing age[12]. Thus, we expect older workers to have a lower propensity to separate from firms with a larger share of young female workers. Table 5 reports estimates of the coefficients of interest corresponding to specifications 4, 5 and 6 from Table 3. The coefficient estimates on the interaction terms at ages 55-59, 60-64, and 65-69 are negative and most are significantly different from zero. Thus older workers have a systematically lower probability of separating from firms with a larger share of young female workers, compared to older workers at firms with a smaller share of young female workers. This suggests that our results using the fraction of older workers as a proxy for flexible technology are not driven by spurious correlation.[13]

Finally, Table 6 presents estimates of our main specification separately for males and females, separately for the periods 1990 – 1995 and 1996 – 2003 (more precisely, the 1990-1993 versus 1996-2001 SIPP panels), and using the quarter-specific employer-level age structure instead of the average age structure. Interestingly, it seems that the age structure of an employer's workforce has a much stronger effect on the separation propensity of older men

[12] The mean age of childbearing was 27.4 years in the U.S. in 2000.

[13] We also estimated a specification that included both the share65-69 and the share of younger female workers (including both the industry and employer-level measures of each share). The effect of the share65-69 was very similar in this specification to the effect shown in Table 3 (column 5), and the effect of the fraction of young women was very similar in this specification to the effect shown in Table 5. This suggests that the two measures may capture different dimensions of an employer's flexibility.

compared to older women. The pattern of coefficient estimates is similar across periods, but the age structure interaction effects are substantially larger in the more recent period. Finally, the age structure interaction effects using the time-varying measure of age structure are much smaller than the corresponding estimates using the employer-average age distribution. This is not surprising given the greater volatility in the quarter-specific age structure of employment.[14]

7 CONCLUSIONS

This study presents the first analysis, of which we are aware, of the association between the age structure of employment in a firm and the propensity of older workers to separate from the firm. The empirical results show a lower separation propensity of older workers, relative to their younger counterparts, in firms with a larger share of older workers. This evidence is consistent with the hypothesis that technology-driven labor market rigidities are manifested in the age structure of employment, and are an important determinant of employment decisions of older workers. Although we have no direct measure of technology-induced labor market rigidities, we argue that the share of older workers at a firm is a useful proxy for the flexibility of technology at the firm. We control for a rich set of worker and firm characteristics that affect separation decisions and that could be correlated with a firm's age structure. This reduces the likelihood that our results are driven by some alternative source of correlation between age structure and turnover behavior. We also show that the relationship holds when an alternative proxy for technological flexibility is used: the share of young women in a firm's workforce. Nevertheless, given the absence of a direct measure of technology, 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.

[14] About one quarter of the total variance in the quarter-specific share65-69 is transitory.

Labor market rigidity is one of several complementary explanations proposed for the prevalence of abrupt retirement. Our results, and evidence presented by Hurd (1996) and Hutchens and Grace-Martin (2006), suggest that labor market rigidity is a plausible explanation. Hamermesh and Donald (2007) present evidence that fixed time costs of employment faced by workers is another plausible explanation. Rust and Phelan (1997) and others have shown that Social Security and Medicare policy provide strong incentives for abrupt retirement by liquidity-constrained workers. The U.S. population will be aging rapidly in the next two decades, and it is generally believed that increasing the employment rate of older individuals will be a necessary part of the adjustment to this major demographic change. Thus it is important to explore all of the possible impediments to increased employment at older ages, including labor market rigidity.

To conclude, some additional limitations of our study are worth mentioning. The approach used here 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 share65-69 would result in a 6% to 8% decline in the log odds of job separation 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 demand-side labor market rigidities, such as team production versus fixed costs of employment. 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? Firm-level data by themselves do not overcome this selection bias. Hence, an important area for future research is to 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 increasing labor force participation at older ages.

APPENDIX: A Model of Labor Market Rigidity

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. For simplicity, we ignore non-labor inputs. The type *A* technology is standard: $Q_A = F^A(L_A)$ where Q is output and L is total hours of labor input. We assume that the marginal product of labor (MPL) is a continuous smoothly declining function of L_A. Thus a type-*A* firm is indifferent to the number of hours worked by any particular worker. The type *B* production function is $Q_B = F^B(L_B*(min\{L_1, L_2, ..., L_N\})^{\theta})$, 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 $\theta > 0$, 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 $Q = F^B(L_B L_{iB}^{\theta}) = F^B(NL_{iB}L_{iB}^{\theta}) = F^B(NL_{iB}^{1+\theta})$, where total labor input $L_B = NL_{iB}$, the number of workers multiplied by hours per worker.

Time is continuous and firms live forever. A given firm is endowed with either type *A* or type *B* technology, and cannot change its type. We assume a steady state environment with a stable population age distribution and no changes in technology. Taking 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 output 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}^D, to maximize profit, with the resulting total number of labor hours demanded by a type *B* firm given by $L_B^D = NL_{iB}^D$. We assume homogeneous firms within sector, with the number of firms per sector normalized to one.

Individuals enter the labor market at age $t = 0$ and live until age *T*. The utility function is $U(C, 1\text{-}L, \delta)$, where *C* is consumption, *1* is total available hours at a given instant, *L* is hours of work, *1-L* is hours of leisure, and $\delta>0$ is a parameter such that the marginal utility of leisure is increasing in δ. Individuals are heterogeneous in leisure preferences. Upon entering the labor market, an individual draws an initial value of δ, δ_0, from the continuous cumulative distribution function $G_0(\delta)$. Shocks to the value of δ arrive randomly according to a stochastic process. For simplicity, we assume that an individual can experience at most one preference shock in his lifetime, and all preference shocks are positive, i.e. increasing the preference for leisure. These assumptions are not essential, but they simplify the exposition considerably. Let $\lambda(t)$ denote the instantaneous hazard rate of the preference shock process, and let $F(x)$ denote the continuous CDF of the distribution of shocks, defined on $x \geq 0$. We assume that the size of the shock is independent of the age of arrival of the shock. The only restriction placed on $\lambda(t)$ is that it is non-decreasing with age. 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. Workers are homogeneous in productivity, both across workers at a given age, and over age for a given worker. There is no cost to a worker of changing sectors. Workers choose whether to work (L=0 or L>0), the sector (A or B) conditional on working, and in sector A the number of hours of work, to maximize utility, taking W_A, W_B, L_{iB}, and Z as given.

There are at least two qualitatively different types of equilibrium in this model. In one type of equilibrium, the hours of work required by firms in sector B, L_{iB}^D, is *less* than the optimal hours of work of the *marginal* worker. The *marginal* worker has a preference for leisure, δ, that leaves him indifferent between working in sector A or sector B, given W_A, W_B, L_{iB}, and Z. Let $d^*(W_A, W_B, L_{iB}, Z)$ denote the value of δ that makes an individual indifferent between the two sectors. In this case, workers with a relatively weak preference for leisure, $\delta < d^*$, choose the flexible sector (A), and workers with a stronger preference for leisure choose the rigid sector (B). In this type of equilibrium, the constraint imposed by technological rigidity is that it *limits* the number of hours worked, rather than requiring too many hours. This type of equilibrium is perfectly legitimate, but is not relevant for our purposes. Hence, we focus on the type of equilibrium of interest by assuming that in equilibrium, $L_A(W_A, Z, d^*) < L_{iB}^D$, where $L_A(\ )$ is the labor supply function of a worker in sector A.[15] In this case, the marginal worker prefers fewer hours of work than the hours required in sector B, hence the constraint imposed by technological rigidity is *excess* hours. In this type of equilibrium, individuals with a strong preference for leisure choose not to work, individuals with a weaker preference for leisure choose to work in Sector A, and individuals with the weakest preference for leisure choose sector B. It is then straightforward to show that the value of δ that makes an individual indifferent between working in sectors A and B, d^*, is defined by $V^A(W_A, Z, d^*) = V^B(W_B, Z, d^*, L_{iB})$, where V^A and V^B are

[15] We parameterized the model and solved it numerically, since an analytic solution does not exist. Both types of equilibrium were found to exist, for alternative parameter values.

indirect utility functions.[16] And there exists a reservation value of δ, $\bar{d} > d^*$, such that a worker is indifferent between sector A and non-employment if $V^A(W_A, Z, \bar{d}) = V^R(Z, \bar{d})$, where V^R represents the utility of retirement (R).

The assumption of a stable population age structure implies that d^* and $\bar{d}$ are constant over time. Thus an individual with $\delta_t < d^*$ chooses sector B at age t, an individual with $d^* \leq \delta_t < \bar{d}$ chooses sector A at age t, and an individual with $\bar{d} \leq \delta_t$ chooses retirement at age t. δ_t for a given worker is equal to the worker's initial draw, δ_0, if he has not received a preference shock by age t; and is equal to $\delta_0 + x$ if he has received a preference shock by age t, where $x \geq 0$ is the value of the shock drawn from distribution F. Normalizing the size of the population to one, total hours of labor supplied to sector B by young workers is

$$L_B^S = \int_0^T L_{iB} G_t(d^*)dt,$$

where G_t is the distribution of δ at age t, and the dependence of the reservation value d^* on wages is implicit. Total hours of labor supplied to sector A for given wage rates is

$$L_A^S = \int_0^T \int_{d^*}^{\bar{d}} L_A(W_A, Z, \delta) dG_t(\delta) dt.$$

The model is closed by the assumption of market clearing: the quantity of labor supplied equals the quantity of labor demanded in each sector: $L_A^S = L_A^D$, $L_B^S = L_B^D$. These two conditions determine the equilibrium values of W_A and W_B, which in turn determine the threshold values d^* and $\bar{d}$, and hours of work required per worker in sector B, L_{iB}.

To derive predictions in the most straightforward way, suppose that: (a) the distribution $G_0(\delta)$ is uniform on $(\delta_{min}, \delta_{max})$, with $\delta_{max} - \delta_{min} = 1$; (b) the distribution $F(x)$ is uniform on (0, 1); and (c) the age distribution is uniform on (0, T). Let t^* represent an arbitrary age that divides

[16] It is also straightforward to demonstrate that in this type of equilibrium, $W_B > W_A$.

older workers from younger workers. Finally, let $Q(t)$ denote the probability that an individual has *not* received a preference shock by age t:

$$Q(t) = exp\{-\int_0^u \lambda(u)\mathrm{d}u\}$$

Given the assumption that the magnitude of the preference shock, x, is independent of the initial value of δ and of the age at which the shock arrives, the probability that a worker chooses sector B at age t is given by

$$\Pr(B_t) = G_t(d^*) = \Pr(\delta_t < d^*) = \Pr(\delta_0 < d^*)(\Pr(\text{no shock by } t) + \Pr(\text{shock by } t,\ \delta_0 + x < d^*))$$

$$= G_0(d^*)(Q(t) + (1\text{-}Q(t))\int_{\delta_{min}}^{d^*} F(d^* - \delta_0)g_0(\delta_0)\mathrm{d}\delta_0 = (d^* - \delta_{\min})(Q(t) + (1\text{-}Q(t))½(d^* - \delta_{\min})^2),$$

where the last equality exploits the functional form assumptions described above. The share of all individuals in sector B, which is constant over time, is then given by

$$\Pr(B) = \int_0^T G_t(d^*)\mathrm{d}t = (d^* - \delta_{\min})(H(T) + (T - H(T))½(d^* - \delta_{\min})^2)/T,$$

where

$$H(T) = \int_0^T Q(t)\mathrm{d}t = \int_0^T exp\{-\int_0^t \lambda(u)\mathrm{d}u\}\mathrm{d}t$$

Using similar logic, the share of *older* individuals in sector B is

$$\Pr(B \mid t > t^*) = (d^* - \delta_{\min})([H(T) - H(t^*)][1 - ½(d^* - \delta_{\min})^2] + (T - t^*)½(d^* - \delta_{\min})^2))/T,$$

where

$$H(t^*) = \int_0^{t^*} Q(t)\mathrm{d}t = \int_0^{t^*} exp\{-\int_0^t \lambda(u)\mathrm{d}u\}\mathrm{d}t$$

The fraction of the labor force in sector B that is old is given by

$$\frac{\Pr(B \mid t > t^*)}{\Pr(B)} = \frac{[H(T) - H(t^*)][1 - ½(d^* - \delta_{\min})^2] + (T - t^*)(d^* - \delta_{\min})^2}{H(T) + (T - H(T))½(d^* - \delta_{\min})^2}$$

Similar derivations for sector A yield the fraction of the labor force in sector A that is old as

$$\frac{\Pr(A \mid t > t^*)}{\Pr(A)} = \frac{[H(T) - H(t^*)][1 - (d^* - \delta_{min})^2] + (T - t^*)(d^* - \delta_{min})^2}{H(T) + (T - H(T))(d^* - \delta_{min})^2}$$

Comparing these fractions, it is easy to show that the fraction of older workers in sector *A* is greater than the fraction of older workers in sector *B* if

$$\frac{H(t^*)}{H(T)} > \frac{t^*}{T}$$

To show that this condition holds, consider each side of the inequality as a function of t^*. From the expressions for $H(t^*)$ and $H(T)$ it is easy to see that both sides of the inequality are equal to zero for $t^*=0$ and both sides are equal to one for $t^* = T$. The right hand side is linear and increasing in t^*. The left hand side is concave and increasing in t^* under the assumption that $\lambda(t)$ is increasing in *t*. The only way for both sides of the expression to be equal at the endpoints when the left hand side is increasing and concave and the right hand side is increasing and linear is for the left hand side to be strictly greater than the right hand side everywhere except at the endpoints. This proves the claim that for $0 < t^* < T$,

$$\frac{\Pr(A \mid t > t^*)}{\Pr(A)} > \frac{\Pr(B \mid t > t^*)}{\Pr(B)}$$

The next question is whether the hazard rate for exit from sector *B* exceeds the hazard rate for exit from sector *A*. We argue in the text that this will be true because changes in leisure preferences can be accommodated in sector *A*, while in sector *B* the only way to accommodate a change in preferences is to leave the sector. We assumed above that an individual experiences at most one preference shock in his lifetime. This greatly simplified the proof that the share of older workers in the flexible sector (*A*) exceeds the share of older workers in the rigid sector (*B*). Under this assumption, the hazard rate for exit from sector *B* at age $t > t^*$, denoted $\gamma_B(t)$, is the probability of receiving a large enough preference shock at age *t*, conditional on not having received a shock prior to t. Under the functional form assumptions described above, this is given by

$$\gamma_B(t) = \lambda(t)(1 - \tfrac{1}{2}(d^* - \delta_{min}))T/(T-t_*)Q(t).$$

The hazard rate for exit from sector A, $\gamma_B(t)$, is equal to the weighted average of the hazard rate conditional on having started life in sector A and not having received a shock by t, and the hazard rate conditional on having started life in sector B and having already received a shock by t that resulted in a switch from B to A. Under the assumption of at most one shock in a lifetime, the latter hazard rate is zero. Under the functional form assumptions described above,

$$\gamma_A(t) = \frac{\lambda(t)Q(t)(\bar{d} - d^*)(1 - \tfrac{1}{2}(\bar{d} - d^*))T}{Q(t) + (1-Q(t))(d^* - \delta_{min})^2(T-t_*)}.$$

Then $\gamma_B(t) > \gamma_A(t)$ if

$$Q(t)(1 - \tfrac{1}{2}(d^* - \delta_{min})) + (1-Q(t))(1 - \tfrac{1}{2}(d^* - \delta_{min}))(d^* - \delta_{min})^2 \;>\; Q(t)(1 - \tfrac{1}{2}(\bar{d} - d^*))(\bar{d} - d^*)^2.$$

This cannot be shown to be true in general. It is more likely to hold if $Q(t)$ is small (the probability of a shock by age t^* is large); $(d^* - \delta_{min})$ is small (sector B is relatively small); and $(\bar{d} - d^*)$ is large (sector A is relatively large). We conjecture that as the number of shocks allowed per lifetime increases, the ambiguity will be resolved, but we have not been able to prove this.

Finally, if there is no rigid sector, then differences in the share of older workers across firms are determined solely by random variation in the arrival rate of preference shocks to the workers in different firms. These differences have no implications for the hazard rate of exit of older workers from employment. Thus, in a version of the model with no variation in technology, there is no reason to expect the share of older workers to be associated with the hazard rate of exit from a firm. In the limit, as firm size grows large, all firms would have the same age structure.

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Table 1: Employment status at wave *t*+1 of individuals who were employed full time, full year with at least five years of job tenure at wave *t*

	Not employed	FT-FY same job	FT-FY new job	PT or PY, same job	PT or PY, new job	Sample size
All	17.7	69.6	4.0	5.9	2.7	14,489
Health in wave t, wave t+1						
good, good	15.1	72.0	4.2	5.8	2.9	11,503
good, bad	30.1	58.2	2.7	7.3	1.6	1,168
bad, good	22.4	66.2	4.2	5.1	2.0	731
bad, bad	29.1	59.3	2.7	6.6	2.3	1,080
Class of worker						
Wage-salary	18.7	70.5	3.5	4.6	2.7	12,358
Self-employed	12.2	64.6	6.6	13.5	3.0	2,130
Pension, Retiree Health Insurance, and Wealth Status						
Does not have DB pension; has EPRHI, wealth > median	18.2	67.4	3.9	7.3	3.2	1,000
Has DB pension, does not have EPRHI, wealth <= median	16.7	76.4	2.7	2.6	1.6	1,009
Others	17.7	69.3	4.1	6.1	2.8	12,296

Source: Health and Retirement Study.

Notes: Sample: HRS cohort born 1931-1941 or married to someone in that birth cohort; age 51-72 at the date of survey; first six survey waves (1992-2002). Good health = self-reported excellent, very good, or good health; bad health = fair or poor health. Wealth is deflated by the CPI. Full time (FT) = 35+ hours per week. Part time (PT) = 1-34 hours per week. Full Year (FY) = 36+ weeks worked per year. Part Year (PY) = 1-35 weeks worked per year. Long tenure = 5+ years with employer. DB = Defined Benefit. EPRHI = Employer Provided Retiree Health Insurance. The wealth distribution is measured at wave *t*. The survey is bi-annual, so the average length of time between waves is two years.

Table 2: Means and Standard Deviations of Selected Sample Characteristics

(standard deviations in parentheses)

		SIPP sample of potential matches		Sample of actual SIPP/LEHD matches	
Age, (years)		52.65	(5.80)	52.57	(5.75)
Five-year age groups, (fractions)	**Age 45-49**	0.37		0.37	
	Age 50-54	0.29		0.29	
	Age 55-59	0.20		0.20	
	Age 60-64	0.11		0.11	
	Age 65-69	0.03		0.03	
Gender, (fractions)	**Males**	0.50		0.50	
	Females	0.50		0.50	
Race, (fractions)	**White**	0.87		0.89	
	Black	0.10		0.08	
	Other	0.04		0.03	
Marital status, (fractions)	**Single**	0.29		0.29	
	Married	0.71		0.71	
Education, (years)		13.45	(2.99)	13.52	(2.92)
Monthly income other than the individual's earnings, ($)		1404	(1740)	1402	(1698)
Wealth, ($ thousands)		111	(932)	123	(1270)
Wage rate, ($ per hour)		9.55	(7.93)	9.85	(7.98)
Initial experience, (years)		22.74	(14.76)	23.98	(14.13)
Tenure, (months)		141.45	(123.22)	143.84	(122.33)
Pension plan coverage, (fraction)		0.48		0.52	
Defined benefit pension plans, (fraction)		0.31		0.32	
Health status, (fraction in good health)		0.91		0.91	
Disabled, (fraction)		0.09		0.08	
Health insurance in own name, (fraction)		0.75		0.78	
Employer provided health insurance, (fraction of those with HI)		0.79		0.83	
***Industry*-specific fraction of 65-69 year old workers**		0.019	(0.010)	0.018	(0.009)
***Industry*-specific fraction of female workers less than 30 years old**		0.127	(0.075)	0.129	(0.075)
Average *employer*-specific fraction of 65-69 year old workers				0.018	(0.030)
Average *employer*-specific fraction of female workers less than 30 years old				0.106	(0.089)
Separated this month, (fraction)		0.012	(0.111)	0.010	(0.099)
Involuntary separations, (fraction of total separations)		0.39		0.37	
Separations leading to change of employer within 30 days, (fraction of total separations)		0.23		0.22	
Number of person-months		907,282		473,034	
Number of individuals		42,687		22,372	

Source: Survey of Income and Program Participation and Longitudinal Employer-Employee Dynamics Files.

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

Table 3: Selected Coefficient Estimates from Logit Models of Monthly Job Separation

(standard errors in parentheses)

	1	2	3	4	5	6
Worker's age group:						
Age50-54	0.068	-0.082	-0.088	0.088	-0.075	0.094
	(0.066)	(0.105)	(0.106)	(0.080)	(0.106)	(0.080)
Age55-59	**0.176**	**0.287**	**0.268**	0.186	**0.268**	0.199
	(0.098)	(0.153)	(0.153)	(0.136)	(0.154)	(0.136)
Age60-64	**0.648**	**0.662**	**0.657**	**0.483**	**0.659**	**0.497**
	(0.133)	(0.207)	(0.207)	(0.191)	(0.208)	(0.192)
Age65-69	**0.853**	**1.111**	**1.096**	**0.770**	**1.073**	**0.822**
	(0.179)	(0.277)	(0.278)	(0.252)	(0.279)	(0.254)
***Industry*-specific fraction aged 65-69**	1.301	-0.587	0.221		-0.455	
	(1.723)	(2.816)	(2.848)		(2.904)	
Age50-54 * *industry*-specific fraction	-1.251	**7.145**	**7.419**		**9.367**	
	(2.341)	(3.989)	(3.976)		(4.087)	
Age55-59 * *industry*-specific fraction	**-7.445**	**-7.908**	-6.078		-5.240	
	(2.550)	(4.180)	(4.155)		(4.308)	
Age60-64 * *industry*-specific fraction	**-19.206**	**-13.602**	**-13.097**		**-10.768**	
	(2.734)	(4.732)	(4.786)		(4.849)	
Age65-69 * *industry*-specific fraction	**-15.252**	**-16.943**	**-16.256**		**-15.451**	
	(3.575)	(5.742)	(5.755)		(5.816)	
***Employer*-specific fraction aged 65-69**				1.402	1.362	2.052
				(1.331)	(1.363)	(1.336)
Age50-54 * *employer*-specific fraction				-2.333	**-3.369**	-2.778
				(1.798)	(1.884)	(1.787)
Age55-59 * *employer*-specific fraction				-2.177	-1.690	-2.606
				(1.726)	(1.774)	(1.723)
Age60-64 * *employer*-specific fraction				**-3.998**	**-3.311**	**-4.474**
				(1.814)	(1.796)	(1.818)
Age65-69 * *employer*-specific fraction				-1.863	-1.496	**-2.476**
				(1.449)	(1.466)	(1.448)
N(person-months)	907,282	473,034	473,034	473,034	473,034	471,104
N(individuals)	42,687	22,372	22,372	22,372	22,372	22,296

Source: Survey of Income and Program Participation and Longitudinal Employer-Employee Dynamics Files.

Notes: All specifications include the additional variables described in the text. Coefficient estimates on the other variables are shown in the Appendix for the specification in column 5. The estimates in column 1 use the sample of SIPP observations potentially matchable to the LEHD. The other columns use the sample of SIPP observations that were actually matched to the LEHD. Columns 3-6 include employer characteristics other than the age distribution (firm-average earnings, ownership type, a multi-plant indicator, and total employment, each averaged over all available quarterly observations for a given firm). The specification in column 6 uses three digit industry dummies instead of two digit dummies. Coefficient estimates in bold are significantly different from zero at the 10% level.

Table 4: Selected Coefficient Estimates from Multinomial Logit models of monthly job separation, by destination and cause of separation

(standard errors in parentheses)

	Destination of separation		Reason for separation	
	New employer		**Employer-initiated**	
Worker's age group:				
Age50-54	-0.060	(0.204)	0.076	(0.174)
Age55-59	-0.151	(0.335)	0.134	(0.250)
Age60-64	-0.029	(0.490)	0.433	(0.356)
Age65-69	-0.956	(0.904)	0.107	(0.470)
***Industry*-specific fraction aged 65-69**	-8.793	(5.379)	**-8.100**	(4.794)
Age50-54 * *industry*-specific fraction	7.555	(7.617)	**15.451**	(6.364)
Age55-59 * *industry*-specific fraction	10.870	(9.400)	**14.649**	(6.278)
Age60-64 * *industry*-specific fraction	-7.779	(13.031)	1.322	(8.308)
Age65-69 * *industry*-specific fraction	8.867	(23.490)	9.045	(9.617)
***Employer*-specific fraction aged 65-69**	0.132	(2.991)	**3.698**	(1.700)
Age50-54 * *employer*-specific fraction	0.115	(3.661)	**-7.572**	(2.747)
Age55-59 * *employer*-specific fraction	-1.036	(3.822)	-2.732	(2.307)
Age60-64 * *employer*-specific fraction	0.297	(4.554)	**-4.819**	(2.284)
Age65-69 * *employer*-specific fraction	-2.587	(5.058)	**-3.688**	(2.230)
	Non-employment		**Worker-initiated**	
Worker's age group:				
Age50-54	-0.036	(0.128)	-0.166	(0.137)
Age55-59	**0.409**	(0.177)	**0.410**	(0.199)
Age60-64	**0.790**	(0.233)	**0.758**	(0.258)
Age65-69	**1.256**	(0.304)	**1.445**	(0.342)
***Industry*-specific fraction aged 65-69**	1.724	(3.523)	5.823	(3.774)
Age50-54 * *industry*-specific fraction	**9.274**	(4.881)	5.234	(5.305)
Age55-59 * *industry*-specific fraction	**-8.846**	(5.008)	**-20.548**	(6.056)
Age60-64 * *industry*-specific fraction	**-11.234**	(5.270)	**-18.820**	(6.006)
Age65-69 * *industry*-specific fraction	**-15.897**	(6.271)	**-28.292**	(7.186)
***Employer*-specific fraction aged 65-69**	1.893	(1.538)	-0.228	(2.098)
Age50-54 * *employer*-specific fraction	**-4.694**	(2.154)	-0.742	(2.653)
Age55-59 * *employer*-specific fraction	-2.276	(2.023)	-0.780	(2.750)
Age60-64 * *employer*-specific fraction	**-4.247**	(1.966)	-2.034	(2.623)
Age65-69 * *employer*-specific fraction	-1.916	(1.632)	0.003	(2.230)
N(person-months)	473,034		473,034	
N(individuals)	22,372		22,372	

Source: Survey of Income and Program Participation and Longitudinal Employer-Employee Dynamics Files.
Notes: All specifications include the additional variables described in the text. The specifications reported here correspond to the specification in column 5 of Table 3. The first pair of columns shows estimates from a multinomial logit model in which the outcomes are started a job with a new employer within 30 days of separation, remained non-employed 30 days after separation, and did not separate. The second pair of columns shows estimates from a multinomial logit model in which the outcomes are involuntary separation (laid off, fired, plant closed), voluntary separation (quit, retired), and no separation. Coefficient estimates in bold are significantly different from zero at the 10% level.

Table 5: Selected Coefficient Estimates from Logit Models of Monthly Job Separation Using the Fraction of Female Workers Less than 30 Years Old as a proxy for Flexibility

(standard errors in parentheses)

	1	**2**	**3**
Worker's age group:			
Age50-54	0.073	0.048	0.073
	(0.090)	(0.101)	(0.091)
Age55-59	**0.254**	**0.299**	**0.268**
	(0.141)	(0.150)	(0.142)
Age60-64	**0.580**	**0.555**	**0.599**
	(0.196)	(0.203)	(0.198)
Age65-69	**0.971**	**0.999**	**1.027**
	(0.260)	(0.271)	(0.262)
***Industry*-specific fraction females < 30**		0.699	
		(0.455)	
Age50-54 * *industry*-specific fraction		0.307	
		(0.601)	
Age55-59 * *industry*-specific fraction		-0.588	
		(0.660)	
Age60-64 * *industry*-specific fraction		0.369	
		(0.716)	
Age65-69 * *industry*-specific fraction		-0.184	
		(0.968)	
***Employer*-specific fraction females < 30**	**0.553**	0.441	**0.516**
	(0.292)	(0.330)	(0.307)
Age50-54 * *employer*-specific fraction	-0.143	-0.277	-0.145
	(0.382)	(0.494)	(0.382)
Age55-59 * *employer*-specific fraction	**-0.779**	-0.474	**-0.854**
	(0.429)	(0.543)	(0.435)
Age60-64 * *employer*-specific fraction	**-1.360**	**-1.550**	**-1.465**
	(0.467)	(0.588)	(0.468)
Age65-69 * *employer*-specific fraction	**-1.724**	**-1.629**	**-1.868**
	(0.650)	(0.802)	(0.657)
N(person-months)	473,034	473,034	471,104
N(individuals)	22,372	22,372	22,296

Source: Survey of Income and Program Participation and Longitudinal Employer-Employee Dynamics Files.

Notes: All specifications include the additional variables described in the text. The specifications reported here correspond to the specifications in columns 4-6 of Table 3. Coefficient estimates in bold are significantly different from zero at the 10% level.

Table 6: Selected Coefficient Estimates from Alternative Logit Models of Monthly Job Separation

(standard errors in parentheses)

	1	2	3	4	5
	Females	**Males**	**1990-1995**	**1996-2003**	**quarter-specific fraction**
Worker's age group:					
Age50-54	-0.115	-0.048	-0.267	0.032	-0.086
	(0.146)	(0.156)	(0.189)	(0.130)	(0.106)
Age55-59	0.239	0.206	0.008	**0.390**	**0.269**
	(0.218)	(0.223)	(0.255)	(0.194)	(0.154)
Age60-64	**0.611**	**0.578**	0.333	**0.764**	**0.661**
	(0.295)	(0.296)	(0.337)	(0.264)	(0.208)
Age65-69	**1.097**	**0.918**	**1.113**	**1.035**	**1.101**
	(0.388)	(0.408)	(0.457)	(0.358)	(0.282)
***Industry*-specific fraction aged 65-69**	-1.442	-0.963	-2.694	0.468	0.072
	(3.959)	(4.379)	(4.885)	(3.618)	(2.899)
Age50-54 * *industry*-specific fraction	**11.081**	8.485	**12.027**	7.680	**7.782**
	(5.492)	(6.339)	(6.913)	(5.138)	(4.096)
Age55-59 * *industry*-specific fraction	-1.224	-5.580	-3.889	-4.908	-5.225
	(6.006)	(6.443)	(7.048)	(5.474)	(4.217)
Age60-64 * *industry*-specific fraction	-1.427	**-16.061**	-10.716	-6.899	**-12.425**
	(6.825)	(6.970)	(7.218)	(6.460)	(4.817)
Age65-69 * *industry*-specific fraction	**-17.448**	-11.319	-11.864	**-15.379**	**-15.585**
	(7.994)	(8.286)	(10.418)	(7.265)	(5.998)
***Employer*-specific fraction aged 65-69**	-1.063	**4.074**	-0.789	3.081	0.419
	(1.594)	(2.115)	(1.937)	(1.903)	(0.965)
Age50-54 * *employer*-specific fraction	-0.123	**-8.422**	-1.020	**-5.354**	-1.019
	(2.188)	(3.585)	(2.907)	(2.488)	(1.516)
Age55-59 * *employer*-specific fraction	0.423	**-4.834**	-1.724	-2.612	-1.575
	(2.395)	(2.623)	(3.106)	(2.358)	(1.363)
Age60-64 * *employer*-specific fraction	-0.329	**-7.122**	-0.980	**-5.639**	-0.911
	(2.041)	(3.244)	(2.450)	(2.734)	(1.248)
Age65-69 * *employer*-specific fraction	1.160	-4.048	-1.026	-3.190	-0.483
	(1.679)	(2.538)	(2.620)	(1.981)	(1.228)
N(person-months)	236,815	236,219	176,692	296,342	455,359
N(individuals)	11,212	11,160	10,128	13,288	22,356

Source: Survey of Income and Program Participation and Longitudinal Employer-Employee Dynamics Files.

Notes: All specifications include the additional variables described in the text. The specifications reported here correspond to the specification in column 5 of Table 3. The specification in column 5 uses employer-quarter-specific measures of the fraction aged 65-69 instead of the employer-specific fraction averaged over all quarters. Coefficient estimates in bold are significantly different from zero at the 10% level.

Appendix Table: Logit Parameter Estimates of the Monthly Job Separation Hazard

(standard errors in parentheses)

	Coefficient	Robust St. Err.
Age50_54	-0.075	(0.106)
Age55_59	**0.268**	(0.154)
Age60_64	**0.659**	(0.208)
Age65_69	**1.073**	(0.279)
Industry-level fraction 65-69	-0.455	(2.904)
Age50_54 * industry fraction 65-69	**9.367**	(4.087)
Age55_59 * industry fraction 65-69	5.240	(4.308)
Age60_64 * industry fraction 65-69	**-10.768**	(4.849)
Age65_69 * industry fraction 65-69	**-15.451**	(5.816)
Employer-level fraction 65-69	1.362	(1.363)
Age50_54 * employer fraction 65-69	**-3.369**	(1.884)
Age55_59 * employer fraction 65-69	-1.690	(1.774)
Age60_64 * employer fraction 65-69	**-3.311**	(1.796)
Age65_69 * employer fraction 65-69	-1.496	(1.466)
Age	**-2.332**	(0.869)
Age squared	**0.042**	(0.016)
Age cubed	**0.0003**	(0.0001)
Male	-0.016	(0.041)
Black	**-0.157**	(0.064)
American Indian	0.099	(0.130)
Asian	-0.154	(0.098)
Married, Spouse Absent	0.060	(0.153)
Widowed	0.101	(0.073)
Divorced	**0.153**	(0.044)
Separated	0.063	(0.099)
Never married	**0.141**	(0.071)
Education	**0.011**	(0.007)
Real income of other household members	1.438	(1.049)
Total household wealth	-0.001	(0.002)
Indicator: Wealth imputed	**-0.421**	(0.099)
Real wage	**0.006**	(0.003)
Indicator: Wage imputed	**1.464**	(0.074)
Tenure	**-0.005**	(0.001)
Tenure squared	**0.000**	(0.000)
First quarter of tenure	**0.148**	(0.051)
First year of tenure	**0.174**	(0.059)
Year 2-5 of tenure	0.060	(0.052)
Initial experience	**-0.006**	(0.002)
Indicator: Experience imputed	-0.038	(0.066)
Pension plan indicator	**-0.266**	(0.102)
DB pension plan indicator	**0.183**	(0.082)
Employer contributions indicator	-0.029	(0.083)
Indicator: Pension information imputed	**1.975**	(0.047)
Disabled	**0.404**	(0.045)
Bad health	-0.012	(0.047)
Indicator: Self-reported health imputed	**-0.624**	(0.073)
Health insurance, own name	**-0.275**	(0.051)
Health insurance, others name	**0.101**	(0.046)
Employer provided health insurance	**-0.375**	(0.048)
Industry:		
Mining	0.393	(0.249)
Construction	0.237	(0.177)
Non-durables	0.202	(0.173)
Durables	0.157	(0.172)
Transportation	-0.072	(0.184)
Public utilities	**0.342**	(0.189)
Wholesale trade	0.154	(0.174)
Retail trade	0.004	(0.168)
Finance	0.022	(0.171)

Continued	Coefficient	Robust St. Err.
Repair services	0.228	(0.169)
Personal services	0.083	(0.181)
Recreation services	0.055	(0.189)
Health services	-0.070	(0.174)
Educational services	0.043	(0.182)
Other services	0.118	(0.172)
Public administration	0.061	(0.198)
Occupation:		
Executives	-0.062	(0.060)
Professionals	**-0.244**	(0.101)
Technicians	-0.030	(0.065)
Sales	0.025	(0.055)
Administrative support	**-0.748**	(0.311)
Private household	0.071	(0.125)
Protective service	**-0.205**	(0.068)
Farming, forestry and fishing	0.134	(0.153)
Craft and repair	0.016	(0.065)
Machine operators	0.003	(0.074)
Transportation and material moving	0.033	(0.083)
Handlers, helpers, and laborers	0.052	(0.095)
Class of worker:		
Private non-profit	**-0.142**	(0.070)
Federal government	**-0.213**	(0.107)
State government	-0.013	(0.116)
Local government	-0.140	(0.114)
Armed forces	**-1.641**	(0.443)
Family business	**-2.212**	(0.731)
Other employer characteristics:		
Firm size <= 5 workers	**-0.405**	(0.087)
Firm size 6-10 workers	**-0.224**	(0.084)
Firm size 11-25 workers	**-0.140**	(0.070)
Firm size 26-50 workers	0.012	(0.067)
Firm size 51-75 workers	0.016	(0.077)
Firm size 76-100 workers	-0.067	(0.083)
Firm size 101-200 workers	0.005	(0.059)
Firm size 201-500 workers	0.008	(0.053)
Firm size 500-1000 workers	**-0.126**	(0.058)
Average number of workers	**-0.040**	(0.019)
Fraction of females in the firm's work force	0.073	(0.092)
Fraction of whites in the firm's work force	0.021	(0.113)
Fraction of blacks in the firm's work force	0.004	(0.177)
Average earnings at the firm	**-4.926**	(2.918)
90th percentile of average earnings	0.801	(0.871)
75th percentile of average earnings	-0.499	(2.090)
50th percentile of average earnings	3.595	(3.645)
25th percentile of average earnings	2.718	(4.692)
10th percentile of average earnings	1.108	(3.278)
Average accession rate	**0.996**	(0.138)
Multi-plant dummy	-0.030	(0.040)
Firm age	0.000	(0.002)
Firm age censored dummy	0.067	(0.056)
State government firm	-0.148	(0.215)
Local government firm	-0.238	(0.193)
Private sector firm	-0.095	(0.170)
State of employment set of 31 dummies	Yes	
Metropolitan area indicator	0.056	(0.040)
Time trend	**0.003**	(0.001)
Constant	-28.122	(25.221)

Notes: Initial experience is equal to the total number of months of work experience as of the beginning of the SIPP coverage. The quarter-specific accession rate is defined as the number of workers with positive earnings in quarter t who were not employed in quarter t-1 divided by the average number of workers in quarters t-1 and t. All workforce demographic, size, earnings and turnover characteristics are further averaged over all quarters of data available for an employer. Firm age is equal to the number of quarters an employer is observed in the LEHD. Firm age is left censored (firm age indicator = 1) if an employer appears in the first quarter of the LEHD coverage period. Coefficient estimates in bold are significantly different from zero at the 10% level.

Figure 1
Average Monthly Separation Rates by Single Year of Age

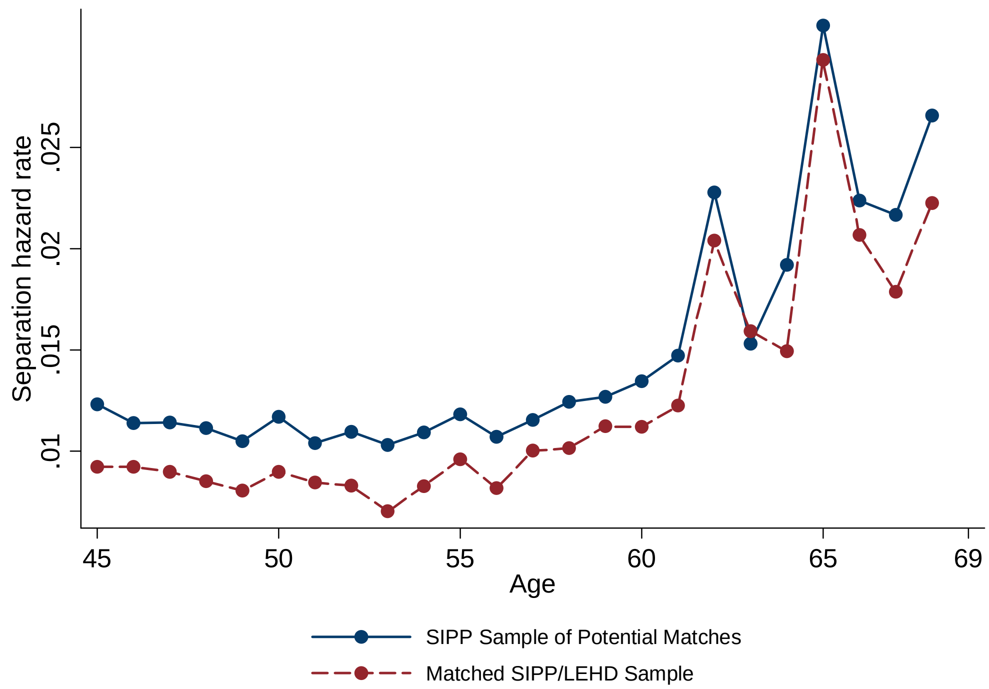

Figure 2
Predicted Monthly Separation Rate
by Single Year of Age and Employer-Specific Fraction of 65-69 Year Old Workers,
No Other Controls

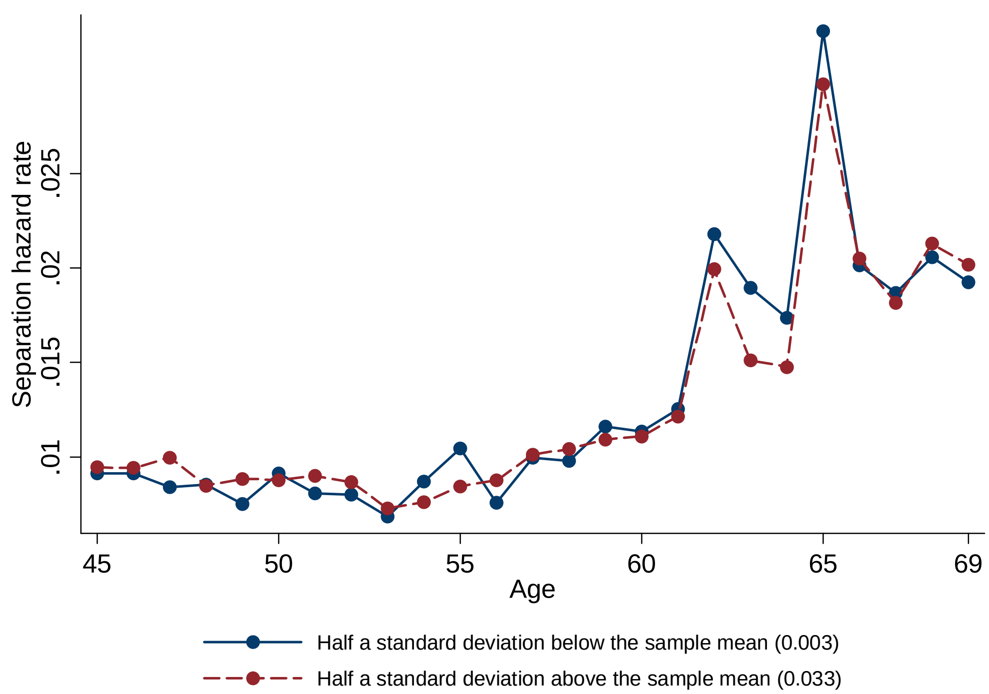

Figure 3
Predicted Monthly Separation Rate by Single Year of Age and Employer-Specific Fraction of 65-69 Year Old Workers with the Full Set of Control Variables

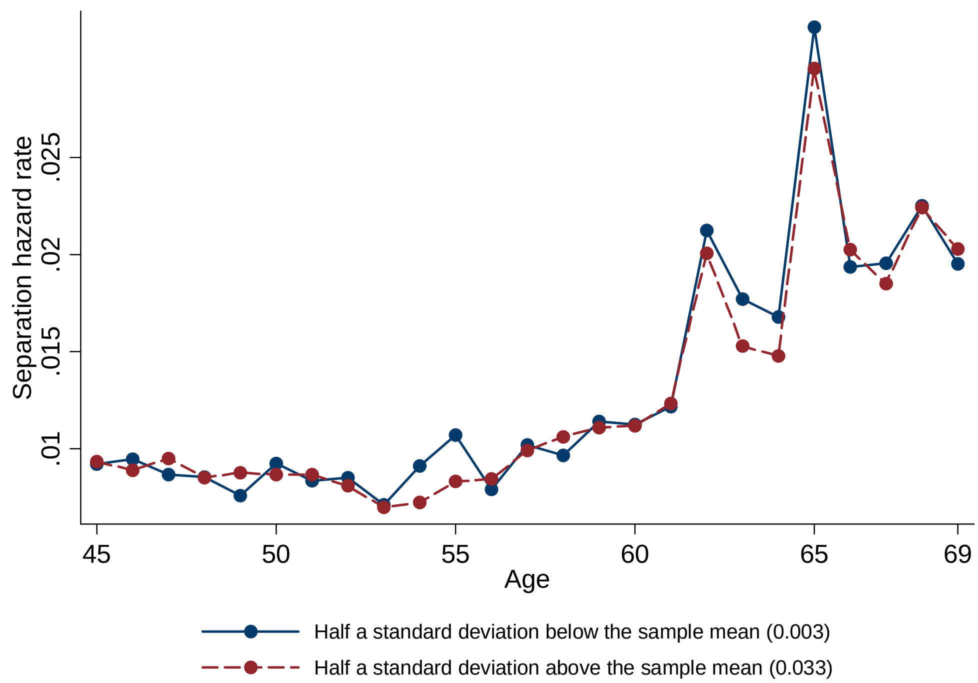

Figure 4
Predicted Monthly Separation Rate by Age and Employer-Specific Fraction of 65-69 Year Old Workers with Five Year Age Group Dummies and the Full Set of Other Control Variables (from specification 5 in Table 3)