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 trajectory from work to retirement. A potentially important source of rigidity is restrictions on hours of work imposed by firms, but such rigidities are difficult to measure directly. We explore two variables that may serve as proxies for flexibility in hours at the employer level: the share of older workers and the share of young women in the employer's workforce. We use matched worker-firm data to analyze the effects of these variables on the job separation propensity of older workers and the incidence of part-time work. The results show that older workers employed in firms with a greater share of older workers and a greater share of young female workers have a lower job separation propensity. These results provide indirect but suggestive evidence of the importance of labor market rigidities in shaping employment decisions of older workers.

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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 (Gustman and Steinmeier, 1984, Ruhm, 1990, Blau, 1994, Maestas, 2004), but less frequent than abrupt and complete exit from employment. Why do people typically retire so abruptly? If most individuals retire as a result of a health shock, then the prevalence of abrupt retirement would be understandable. Deterioration in health is an important cause of retirement, but, as shown below, most changes in employment status at older ages are not associated with a decline in self-reported health. In the absence of a health shock, it seems implausible that preferences for leisure would change abruptly at older ages. One indication that sudden 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.

Many factors could be responsible for making the labor market rigid. Retirement incentives facing older workers often change abruptly at specific ages, 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 (see Haider and Loughran, 2008, and Song and Manchester, 2007, and references cited therein). 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 do not 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 supply side of the labor market, fixed costs of being employed may make part-time employment unattractive to many individuals (Hamermesh and Donald, 2007). On the demand side of the labor market, fixed costs of hiring, training, and employing a worker, could induce firms to impose minimum hours-of-work constraints 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 work schedules. [2]

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. But the preference for flexibility in employment is stronger at older ages and among women of childbearing age. Consequently, the existence of technology-induced rigidities could be manifested in the age and gender structure of a firm's work force: the more important are technology-induced rigidities, the lower are the shares of older workers and younger women at the firm.

There is evidence that production technology differs substantially across firms, even within narrowly defined industries. 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). 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 and gender structure of the work force across firms.

In this paper, we study the effect of the employer-level age and gender composition of employment on the separation propensity and hours worked of older workers. We focus on older workers because of their rapidly growing importance, due to population aging. 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

[2] 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).

Lane, 2004). We use a difference-in-differences approach to compare the job exit behavior of older and younger workers in firms with different shares of older workers and younger women. Comparing older and younger workers makes it possible to determine whether labor market rigidities, proxied by the age and gender structure of the work force, affect older workers disproportionately, as we hypothesize. In order to ensure that the firm's age and gender composition is not merely picking up the effects of other factors, we control for many worker and firm characteristics. The empirical results show that a larger share of older workers and a larger share of younger women in a firm's work force are associated with a lower job separation propensity of older workers. We do not find any association between the age and gender structure of the firm and the incidence of part-time employment.

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

2 BACKGROUND AND LITERATURE

First, we illustrate our claim that the majority of workers retire by moving directly from full-time employment to complete retirement and that this pattern cannot be fully explained by worker characteristics and incentives.

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 six 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, 3.9% were employed on a new year-round full-time job, 5.8% were employed part-time or part-year with the same employer, and 2.8% were employed part-time or part-year with a new employer. Of the total of 30.3% of respondents who changed employment status between survey waves, the majority (58.4%) made a complete exit from employment.[3]

Deterioration in health is a major cause of retirement, but most changes in employment status at older ages are not associated with a decline in health. The second panel of 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 about 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: 69% of exits from employment were by individuals whose health remained good, compared to only 13% whose health declined from good to bad.

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, facilitating gradual retirement (Karoly and Zissimopoulos, 2004). 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 16.6% for the self-employed. This clearly suggests that wage-salary workers face a constraint on hours of work imposed by their employers.

[3] 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.

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 fourth panel of Table 1 shows that workers who are covered by DB pensions, do not have EPRHI, and are in the lower quartile of the distribution of net worth are in fact more likely to retire abruptly (72% 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 quartile of the distribution of net worth (55%). 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.[4]

Direct evidence on the demand-side 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). Using data from the HRS, 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. The majority (82%) of the establishments in a survey conducted by Hutchens and Grace-Martin (2006) reported that they have a phased retirement policy.[5] Most of these policies were informal and discretionary, and fewer than half of

[4] We verified that all of the bivariate associations shown in Table 1 are robust by estimating a multinomial logit model with the five-way classification of employment status shown in Table 1 as the dependent variable. Controlling for a full set of age fixed effects, industry, occupation, job tenure, work experience, marital status, the wage rate, physical demands of the job, race, the spouse's employment status, and gender, all of the bivariate associations described in the text are verified, and most of these conditional associations are very similar in magnitude to the bivariate associations. The only exception is the difference by health status, which falls from 13.4 percentage points to only one percentage point. This model does not directly control for Social Security incentives, but the age dummies control for the spikes in labor force exit at ages 62 and 65 that are commonly attributed to Social Security incentives.

[5] The authors surveyed 950 establishments with at least 20 employees and two white collar employees aged 55+, and posed questions about phased retirement policy.

the establishments with a phased retirement policy reported that any older employees had actually chosen to shift from full-time to part-time work in the three years prior to the survey.[6]

In order to provide additional evidence on flexibility in hours of work in relation to worker age, we examine data from the May 2001 Current Population Survey Supplement on Work Schedules and Work at Home. We analyze responses to the question "Do you have flexible hours that allow you to vary or make changes in the time you begin and end work?" Figure 1 shows the age profile of responses to this question for workers aged 30-70. The proportion reporting flexible hours is roughly constant at around 0.35 - 0.37 from age 30 to 58, and then increases sharply to over 0.50 by age 70. This age pattern persists after controlling for demographic characteristics of the worker, and detailed occupation, industry, and class of worker dummies in a regression (see Figure 1). This evidence clearly indicates that older workers are more likely to have flexible work schedules, but it does not directly address the issue of age structure of the work force as a proxy for hours flexibility. To get at this issue, we computed measures of age structure at the three-digit industry level from the 1990 U.S. Census of Population (described in more detail below) and merged them with the CPS data. Controlling for all of the variables mentioned above (including 40 worker-age dummies and 51 two-digit industry dummies), we find that the fraction of workers aged 60-64 in the three-digit industry has a positive and statistically significant effect on the probability that a worker has flexible hours. The same finding holds for the fraction of workers aged 65-69 and 70-74. A one standard deviation increase in these fractions is associated with a 1-2 percentage point increase in the probability of flexible hours. This evidence supports our contention that age structure and

[6] The survey did not inquire about any conditions that might be associated with phased retirement, such as wage cuts and pension eligibility.

flexible hours will be positively associated, although it does not establish the direction of causality.[7]

A commonly cited demand-side source of potential labor market rigidity is team production, which often requires workers to be present at fixed hours in order to work with other team members. This limits flexibility in work hours, and may reduce the attractiveness of the job for workers seeking flexibility. The only data source we could find with information on both the age and gender structure of employment and the prevalence of team production is the British Workplace Employee Relations Survey (BWERS). This 1998 survey of establishments with at least 10 employees asked managers to report the proportion of the largest occupation group that works in teams, the proportion of all employees over age 50, and the proportion of female employees. There is a statistically significant negative relationship between the prevalence of team production and the proportion of workers over the age of 50, with or without controls for other factors. A 10 percentage point increase in the prevalence of team production is associated with a 0.3 point reduction in the percentage of workers over the age of 50. For women, the unconditional association is positive, but conditional on industry and worker characteristics, the association is negative, though insignificant. These associations lend support to the notion that age and gender structure can serve as a useful proxy for demand-side labor market rigidities.[8]

For additional suggestive evidence, we computed the industry-mean proportion of employees who work in teams in the BWERS data and used it to examine the cross-industry

[7] In contrast, the association between flexible hours and the share of women aged 30 or less is negative, though very small and insignificantly different from zero. A graph like Figure 1 for women only, with the age range extended down to 20, does not show any evidence that flexible work hours are more common among women aged less than 30, compared to older women. It would be interesting to examine whether women under 30 with young children are more likely to have flexible hours, but the May CPS does not have data on family structure. This information is available in the March CPS, but merging the March and May survey is difficult and would reduce the sample size for analysis of flexibility in hours by over 80%.

[8] Hamilton, Nickerson, and Owan (2003) analyze the impact of team production on the propensity of workers to separate from a firm. However, they do not examine how the effect of teams on separation differs by worker age.

association between team production and the finer measures of age and gender structure available in the LEHD (described in Section 4).[9] The sample size in the BWERS data was too small to produce reliable statistics at a finely disaggregated industry level, so the analysis is based on only 12 industries.[10] In this sample of 12 industries, the correlation between proportion working in teams and proportion of older workers (60-64 and 65-69) is positive, rather than negative as we expected. However, there is a negative correlation between team proportion and the share of women under age 30 in the industry's workforce. Not surprisingly, these correlations are insignificant, given the small sample size as well as the fact that most of the variation in team production and age and gender structure is within-industry.[11]

3 CONCEPTUAL FRAMEWORK

Here, we describe the logic of our conceptualization of technological sources of labor market rigidity and their impact on the employment behavior of older workers, although we do not present a formal model.[12] Suppose 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

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[9] The concordance between the 1992 British SIC and the 2002 NAICS code available in the LEHD is as follows, with the British code listed first and the corresponding NAICS codes in parentheses: 1 (31-33) – manufacturing, 2 (22) – electricity, gas and water, 3 (23) - construction, 4 (42, 44, 45) – wholesale and retail, 5 (72) – hotels and restaurants, 6 (48, 49, 51) – transport and communication, 7 (52) – financial services, 8 (53-56) – other business services, 9 (92) – public administration, 10 (61) – education, 11 (62) – health, 12 (71, 81) – other community services.

[10] There is substantial variation in the BWERS team production measure even across these broadly aggregated industries, from 40 - 45% in hotels and restaurants, manufacturing, transport and communication industries, to 82% in financial services and electricity, gas and water.

[11] The R^2 in an establishment-level regression of team proportion on the 12 industry dummy variables using the BWERS establishment-level data is only 0.10. The R^2 increases to 0.39 when a much more disaggregated 5 digit industry coding is used.

[12] 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 and gender structure of employment at the firm level. Hence our approach is complementary with these alternative models.

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. As discussed above, the rigidity could arise from team production or fixed costs of hiring, training, or employment. 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 pay 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 entails 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. 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 propensity of older workers to separate from a firm will be higher for firms with a younger age structure. Thus, the age structure of a firm’s work force can serve as a proxy for the degree of technological rigidity. A similar argument applies to other groups of worker who value employment flexibility, such as younger women.

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.

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

Alternatively, suppose workers prefer to work with coworkers 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.[13]

Finally, in a steady state, the age structure of a firm's labor force is determined by age-specific inflow and outflow rates. A higher separation rate of older workers results in a lower steady state share of older workers in a firm's labor force, other things equal. Thus, reverse causality will induce a *negative* association between the share of older workers in a firm's labor force and the separation propensity of older workers, as predicted by our hypothesis, but for reasons unrelated to our explanation.

To deal with these and other possible sources of association between age structure and separation propensity that are unrelated to technology, we mainly focus on the other proxy for a firm's technological flexibility described above: 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

[13] 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.

are much more likely to occupy part-time and flexible-hours jobs than are other workers.[14] If the share of young female workers is negatively associated with the separation propensity of older workers, this would be difficult to explain by mechanisms other than technology-based inflexibility in hours of work. Finally, reverse causality is unlikely to be a problem when using the share of younger women as a proxy for flexible employment practices. A higher separation rate of older workers will cause a higher share of younger workers at the firm in the steady state, but there are no implications for the steady state share of any particular group of younger workers, such as women under age 30.[15]

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, 2006; Van der Klaauw and Wolpin, 2008).[16] We augment this list with measures of the age and gender composition of employment at the individual's firm. A simple illustration of our empirical specification is

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

[14] 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.

[15] We explored several Instrumental Variables strategies to deal with the endogeneity of the share of older workers, including using the share of younger women as an instrument, but none were successful.

[16] It is straightforward to include the wage rate in the analysis, 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 interpret the effects of variables such as the wage rate, years of work experience, and job tenure as operating in part via their impact on anticipated future pension and Social Security benefits.

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 a proxy for the degree of employment flexibility at firm j. This is a discrete time hazard model of the risk of separation, and is estimated by logit.

The coefficient of interest is δ: the difference between the effect of the flexibility proxy 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 flexibility proxy γ captures any effects of age and gender composition of employment that are independent of the worker's own age. For example, firms with relatively few older workers may tend to be newer, and firm age may affect the separation propensity of all workers at the firm. Controlling for pension and health insurance coverage, occupation, industry, and the wage rate (all included in X), we interpret differential effects of a firm's workforce age and gender composition on the separation hazard of 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 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 and family formation 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 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 in each state. 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 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 (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.[17] 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.

Table 2 presents summary statistics. 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

[17] 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.

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 a binary 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 the reasons discussed in the previous paragraph. Figure 2 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 fraction of women aged less than 30 in an employer's work force. This is our preferred proxy for the flexibility of the employer's working conditions. We also present some results using the share of workers aged 65-69 as the proxy for flexibility, but as noted above, this measure is very likely endogenous. We use the employer-specific fraction of younger women averaged across all observed quarters for a given employer. This provides a relatively stable measure that is not subject to transitory quarter-to-quarter variation. We also control for the industry-level share of young women. We compute this 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-level age and gender composition variables with the SIPP based on a worker's self-reported three-digit industry. The mean employer-specific share of women under age 30 is 0.106, compared to 0.129 at the industry level.

5 RESULTS

To illustrate the basic patterns of interest, we first estimated a logit model of the monthly hazard of separation using a set of single-year age dummies, the fraction of women aged less than 30 (abbreviated as sharewomlt30 henceforth) at the individual's employer, and interactions of these variables, with no other control variables. A similar model was estimated using the fraction of 65-69 year old workers (abbreviated as share65-69). Figure 3a depicts the pattern of the predicted monthly separation hazard rate for two different values of the sharewomlt30: half a standard deviation below the sample mean (0.06) and half a standard deviation above the mean (0.15). The separation rate is predicted for each person-month and then averaged by age. The results in Figure 3a suggest that beginning at age 62 the separation propensity of workers is lower when the sharewomlt30 is higher. This is the pattern predicted by our hypothesis. Interestingly, this bivariate association pattern is reversed at younger ages: workers aged 45-60 firms separate at a higher rate from firms with larger sharewomlt30. Figure 3b shows a similar pattern for share65-69: a higher value of this proxy for employment flexibility is associated with a lower separation propensity at older ages with few and much smaller differences at younger ages.

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, 16 industry dummies, 12 occupation dummies, 6 class of worker dummies, job tenure, work experience, pension plan type, health insurance coverage, size of the employer (number of workers), the demographic characteristics and earnings distribution of the employer's workforce, ownership type, a multi-plant indicator, the employer's age[18], region indicators, and a linear time trend. This specification also controls for the industry-level proxies for employment flexibility and their interactions with single-year age dummies. Figures 4a and 4b present the average predicted separation propensity by age based on this specification, for alternative values of sharewomlt30 and share60-65. The negative association between sharewomlt30 and the separation hazard at older ages remains visible even after controlling for many other factors that are likely to influence employment behavior. The same finding is apparent for share65-69. These results suggest an association between the share of older workers in a firm and the separation propensity of older workers. Controlling for worker and firm characteristics eliminates the differences in the separation propensity at younger ages across firms with different age and gender workforce structure.

Table 3 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 omitted age category is 45-49). Results are presented for both flexibility proxies, and for both the employer and industry level flexibility proxies. The table shows logit estimates of selected coefficients (α, γ, and δ from eq. 1), using sharewomlt30 as the proxy for employment flexibility in the first column, and share6569 in the third column. The main effects of the employer-level sharewomlt30 and share65-69 are both positive, while the age interaction effects are negative. Two of the four interaction effects are significantly different from zero in each case. Our hypothesis implies that the age interaction effects should be negative and increase in absolute value with age. This pattern is evident for the sharewomlt30, but is more irregular for

[18] Firm age is equal to the number of quarters an employer is observed in the LEHD. Firm age is left censored if an employer appears in the first quarter of the LEHD coverage period. A dummy variable indicating whether the firm's age is left censored is included in the model.

the share65-69. The estimates in column 1 imply that a one standard deviation increase in the sharewomlt30 (0.089) would cause the log odds of the monthly separation probability to fall by 0.099 at ages 60-64 ($0.089 \times [0.441 - 1.550]$) and by 0.106 at ages 65-69 ($0.089 \times [0.441 - 1.629]$). Evaluated at the mean separation probability (0.010), the implied effects on the probability of separation are -0.00099 and -0.00106, or about 10% of the mean separation probability.[19] The estimates in column 3 imply that a one standard deviation increase in the share65-69 (0.03) would cause the monthly separation probability to decline by 0.00058 and 0.00004 at ages 60-64 and 65-69 respectively, or 5.8% and 0.4% of the mean.[20]

It is interesting to note that there are also strong negative age interaction effects with the industry-level share65-69. A one standard deviation (0.009) increase in the industry-level share65-69 would reduce the probability of separation by 0.001 at ages 60-64 and by 0.0014 at ages 65-69. However, the effects of the industry-level sharewomlt30 are small and insignificantly different from zero at older ages. The strong effect of the industry-level share65-69 raises the question of whether the employer-level flexibility proxies capture the effects of unobserved characteristics of the employer's industry that affect worker turnover. The specification in columns 1and 3 controls for 16 broad industry dummies and the industry-level flexibility proxies (measured at the three digit industry level), but this may be too crude to capture industry effects. Columns 2 and 4 control for three-digit industry fixed effects instead of the broader industry fixed effects used in columns 1 and 3. The three-digit industry fixed effects control for all industry-level factors that could be associated with the separation propensity,

[19] The full set of parameter estimates for this specification is presented in the Appendix Table.
[20] We also estimated a specification that included both the share65-69 and the share of younger female workers. The estimates (available on request) of the key interaction effects of interest are very similar to those reported in columns 1 and 3.

including observed factors such as the industry-specific age structure, and other unobserved factors. The main results are robust to this specification change.

As noted above, the control variables include an indicator for whether the SIPP worker is covered by a DB pension plan. These plans often contain strong incentives to leave the employer at the plan's early retirement age, which is typically between 55 and 62 and varies across plans. DB pension plans with strong early retirement incentives could be more prevalent in firms with inflexible technology. For example, if it is difficult for older workers to keep up with the desired work pace on an assembly line (a classic example of team production), firms might use a DB pension plan to give such workers an incentive to retire early. In this case, the estimated negative effects of the employer flexibility proxies would be biased away from zero (i.e. too large in absolute value). The SIPP does not collect information about the early retirement age in DB plans, so we cannot control for it, and the LEHD has no information about pensions. Instead, we re-estimated the models in Table 3 on a sample that excludes workers who reported being covered by a DB plan (31% of the sample; see Table 1). The results were very similar to those reported in Table 3 (we omit them for brevity). Hence this does not seem to be a problem in practice.

6 ALTERNATIVE OUTCOMES AND SPECIFICATIONS

Next, we examine the destination of job separations, in order to determine whether greater employer flexibility reduces both exits from the labor force and job switching. 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. Separations resulting in unemployment, withdrawal from the labor force, and of undetermined destination are classified as leading to non-employment. Slightly more than 20% of monthly job separations are followed by a change of employers within 30 days of the separation (see Table 2). Panel 1 of Table 4 presents selected estimates from a multinomial logit model in which the outcomes are separation leading to a change in employer, separation leading to non-employment, and no separation using sharewomlt30 as the flexibility proxy. The results indicate that greater flexibility (a larger sharewomlt30) reduces separations leading to non-employment at ages 60-69. 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. The effects of larger sharewomlt30 on employer-to-employer separations are actually positive at these ages. This result is inconsistent with our story.

Another way to disaggregate separations is by the proximate cause: employer-initiated (laid off, fired, plant closed) versus worker-initiated (quit, retired). The results in panel 2 of Table 4 are from a multinomial logit model in which the outcomes are employer-initiated separation, worker-initiated separation, and no separation. The results indicate that employer flexibility reduces both employer-initiated and worker-initiated separations at older ages. Our reasoning predicts a negative association between employment flexibility and worker-initiated separations, but has no predictions about employer-initiated separations. The finding of similar effects for both types suggests that sharewomlt30 may capture other firm characteristics in addition to work schedule flexibility. Alternatively, the distinction between the two types of separations may not be meaningful, as in some theories of efficient turnover.

Table 5 presents estimates separately for men and women. The sharewomlt30 has a stronger negative effect on the separation propensity of older men compared to older women. Women may have a much stronger demand than men for flexible hours during the childbearing years, but it seems that men have a stronger preference for flexibility at older ages.[21] Finally, Table 6 shows results from a binary logit model of part-time work, conditional on employment. We expected that employment flexibility would have a positive effect on part-time work for older workers. However, the results do not show this: the age interaction effects are all of the wrong sign and insignificantly different from zero.[22] Despite the high prevalence of part-time work among young women (see note 8), a higher sharewomlt30 is not associated with a higher probability of part-time work among older workers. Using the share65-69 as the flexibility proxy yields the same result (not shown): no evidence of greater part-time employment in firms with a greater share65-69. It is possible that flexible hours can take forms other than part-time work; for example, flexible work days and schedules, long vacations, etc. Nevertheless, the absence of an association between sharewomlt30 and share65-69 and the incidence of part-time employment suggests caution in accepting our interpretation of these variables as proxies for employment flexibility.

7 CONCLUSIONS

This study analyzes the association between the age and gender 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 and a larger share of young female workers. This evidence is consistent with the hypothesis that technology-driven labor market rigidities are manifested in the age and gender structure of employment. Although we have no direct measure

[21] The result that the separation rates of older men are lower in firms with a higher fraction of young female workers also means that our proxy is not just picking up the fact that separation rates are higher at all ages in "female-oriented" firms and industries.
[22] It would be of considerable interest to use the share of part-time employment in an employer's workforce as an explanatory variable, but the LEHD data do not contain information on hours of work.

of technology-induced labor market rigidities, we argue that the share of older workers and the share of younger women are useful proxies for the flexibility of technology at the firm. There is reason to expect that the share of older workers is endogenous to the separation propensity of older workers, but there is no reason to expect the share of younger women to be endogenous, so the robustness of the findings across these two measures is reassuring. 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 and gender structure. This reduces the likelihood that our results are driven by some alternative source of correlation between age and gender structure and turnover behavior. 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 employment behavior of older workers, but clearly not as definitive evidence.

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 an increasing 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 the demand-side sources of 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 sharewomlt30 would reduce the monthly separation probability by about 10% at older ages. 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.

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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	69.7	3.9	5.8	2.8	13,462
Health in wave t, wave t+1						
good, good	15.2	71.9	4.2	5.7	3.0	10,762
good, bad	30.2	58.7	2.8	6.7	1.5	1,057
bad, good	22.0	66.1	4.4	5.3	2.3	664
bad, bad	28.9	59.9	2.4	6.5	2.2	979
Class of worker						
Wage-salary	18.6	70.6	3.5	4.6	2.7	11,616
Self-employed	12.6	64.1	6.7	13.5	3.1	1,846
Pension, Retiree Health Insurance, and Wealth Status						
Does not have DB pension; has EPRHI, wealth in upper quartile	18.7	67.2	3.9	7.3	3.0	1,045
Has DB pension, does not have EPRHI, wealth in lower quartile	16.4	77.1	2.9	2.0	1.6	890
Others	17.8	69.3	4.0	6.0	2.9	11,527

Source: Health and Retirement Study.

Notes: The sample is the HRS cohort born 1931-1941 or married to someone in that birth cohort; age 51-72 at the date of survey; first seven survey waves (1992-2004). 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)		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)
***Employer-specific f*raction of 65-69 year old workers**		0.018 (0.030)
***Employer-specific f*raction 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
Age50-54	0.049	0.073	-0.075	0.094
	(0.101)	(0.091)	(0.106)	(0.080)
Age55-59	**0.299**	0.268	**0.268**	0.199
	(0.150)	(0.142)	(0.154)	(0.136)
Age60-64	**0.555**	**0.599**	**0.659**	**0.497**
	(0.203)	(0.198)	(0.208)	(0.192)
Age65-69	**0.999**	**1.026**	**1.073**	**0.822**
	(0.271)	(0.262)	(0.279)	(0.254)
***Industry*-specific fraction females < 30**	0.699			
	(0.455)			
***Industry*-specific fraction aged 65-69**			-0.455	
			(2.904)	
Age50-54 × *industry*-specific fraction	0.307		**9.367**	
	(0.601)		(4.087)	
Age55-59 × *industry*-specific fraction	-0.588		-5.240	
	(0.660)		(4.308)	
Age60-64 × *industry*-specific fraction	0.369		**-10.768**	
	(0.716)		(4.849)	
Age65-69 × *industry*-specific fraction	-0.184		**-15.451**	
	(0.968)		(5.816)	
***Employer*-specific fraction females < 30**	0.441	0.516		
	(0.330)	(0.307)		
***Employer*-specific fraction aged 65-69**			1.362	2.052
			(1.363)	(1.336)
Age50-54 × *employer*-specific fraction	-0.277	-0.145	**-3.369**	-2.778
	(0.543)	(0.382)	(1.884)	(1.787)
Age55-59 × *employer*-specific fraction	-0.474	**-0.854**	-1.690	-2.606
	(0.543)	(0.435)	(1.774)	(1.723)
Age60-64 × *employer*-specific fraction	**-1.550**	**-1.465**	**-3.311**	**-4.474**
	(0.588)	(0.468)	(1.796)	(1.818)
Age65-69 × *employer*-specific fraction	**-1.629**	**-1.868**	-1.496	**-2.476**
	(0.802)	(0.657)	(1.466)	(1.448)
N(person-months)	473,034	471,104	473,034	471,104
N(individuals)	22,372	22,296	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 specification in columns 2 and 4 uses three digit industry dummies instead of two digit dummies, and omits the industry-specific share variables, since they are calculated at the three-digit industry level. Coefficient estimates in bold are significantly different from zero at the 5% level.

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

(standard errors in parentheses)

	1. Destination of separation		2. Reason for separation	
	New employer	**Non-employment**	**Employer-initiated**	**Worker-initiated**
Age50-54	-0.079	0.111	0.297	-0.135
	(0.198)	(0.121)	(0.163)	(0.132)
Age55-59	0.305	**0.339**	0.355	0.303
	(0.328)	(0.172)	(0.242)	(0.194)
Age60-64	-0.212	**0.669**	0.393	**0.645**
	(0.502)	(0.228)	(0.343)	(0.253)
Age65-69	-1.606	**1.232**	0.432	**1.226**
	(0.827)	(0.293)	(0.469)	(0.329)
***Industry*-specific fraction females < 30**	0.874	0.589	**1.530**	0.322
	(0.812)	(0.551)	(0.762)	(0.584)
Age50-54 × *industry*-specific fraction	1.768	-0.092	-0.098	0.711
	(1.021)	(0.749)	(1.008)	(0.777)
Age55-59 × *industry*-specific fraction	-0.820	-0.460	1.083	**-1.732**
	(1.425)	(0.765)	(1.076)	(0.865)
Age60-64 × *industry*-specific fraction	0.087	0.580	2.009	-0.647
	(1.844)	(0.792)	(1.203)	(0.883)
Age65-69 × *industry*-specific fraction	1.453	-0.143	-0.962	-0.133
	(2.665)	(1.018)	(1.977)	(1.131)
***Employer*-specific fraction females < 30**	-0.087	0.625	-0.811	**1.317**
	(0.623)	(0.385)	(0.569)	(0.421)
Age50-54 × *Employer*-specific fraction	-0.737	-0.206	-0.256	-0.324
	(0.851)	(0.600)	(0.929)	(0.588)
Age55-59 × *Employer*-specific fraction	-1.364	-0.340	-0.880	-0.322
	(1.330)	(0.610)	(0.928)	(0.691)
Age60-64 × *Employer*-specific fraction	0.228	**-1.807**	**-2.466**	-1.382
	(1.416)	(0.645)	(1.100)	(0.702)
Age65-69 × *Employer*-specific fraction	**3.820**	**-2.107**	0.028	**-2.518**
	(1.798)	(0.849)	(1.794)	(0.863)
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 1 of Table 3. Panel 1 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. Panel 2 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 5% level.

Table 5: Selected Coefficient Estimates from Logit Models of Monthly Job Separation by Gender

(standard errors in parentheses)

	Job Separation	
	Women	**Men**
Age50-54	0.123 (0.150)	0.028 (0.143)
Age55-59	0.429 (0.223)	0.160 (0.216)
Age60-64	**0.903** (0.294)	0.266 (0.292)
Age65-69	0.785 (0.402)	**1.014** (0.380)
***Industry*-specific fraction females < 30**	0.995 (0.597)	0.709 (0.737)
Age50-54 × *industry*-specific fraction	0.109 (0.786)	0.288 (0.999)
Age55-59 × *industry*-specific fraction	-0.605 (0.920)	-0.807 (1.054)
Age60-64 × *industry*-specific fraction	-0.612 (0.993)	0.979 (1.152)
Age65-69 × *industry*-specific fraction	0.444 (1.382)	0.361 (1.384)
***Employer*-specific fraction females < 30**	**0.852** (0.382)	-0.425 (0.670)
Age50-54 × *employer*-specific fraction	-0.195 (0.586)	-0.757 (0.905)
Age55-59 × *employer* -specific fraction	-0.617 (0.690)	-0.083 (0.944)
Age60-64 × *employer* -specific fraction	-1.440 (0.746)	-1.862 (1.012)
Age65-69 × *employer* -specific fraction	-0.673 (1.050)	**-3.895** (1.247)
N(person-months)	236,815	236,219
N(individuals)	11,212	11,160

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. Coefficient estimates in bold are significantly different from zero at the 5% level.

Table 6: Selected Coefficient Estimates from a Logit Model of Part-time Employment

(standard errors in parentheses)

	Part-time Employment
Age50-54	-0.003
	(0.134)
Age55-59	-0.027
	(0.172)
Age60-64	0.178
	(0.204)
Age65-69	0.146
	(0.266)
***Industry*-specific fraction females < 30**	**1.669**
	(0.643)
Age50-54 × *industry*-specific fraction	-0.235
	(0.829)
Age55-59 × *industry*-specific fraction	-0.472
	(0.907)
Age60-64 × *industry*-specific fraction	-0.484
	(1.021)
Age65-69 × *industry*-specific fraction	-0.084
	(1.339)
***Employer*-specific fraction females < 30**	-0.384
	(0.492)
Age50-54 × *employer*-specific fraction	-0.527
	(0.634)
Age55-59 × *employer*-specific fraction	-0.862
	(0.705)
Age60-64 × *employer*-specific fraction	-0.763
	(0.795)
Age65-69 × *employer*-specific fraction	-0.512
	(1.511)
N(person-months)	473,034
N(individuals)	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. Coefficient estimates in bold are significantly different from zero at the 5% level.

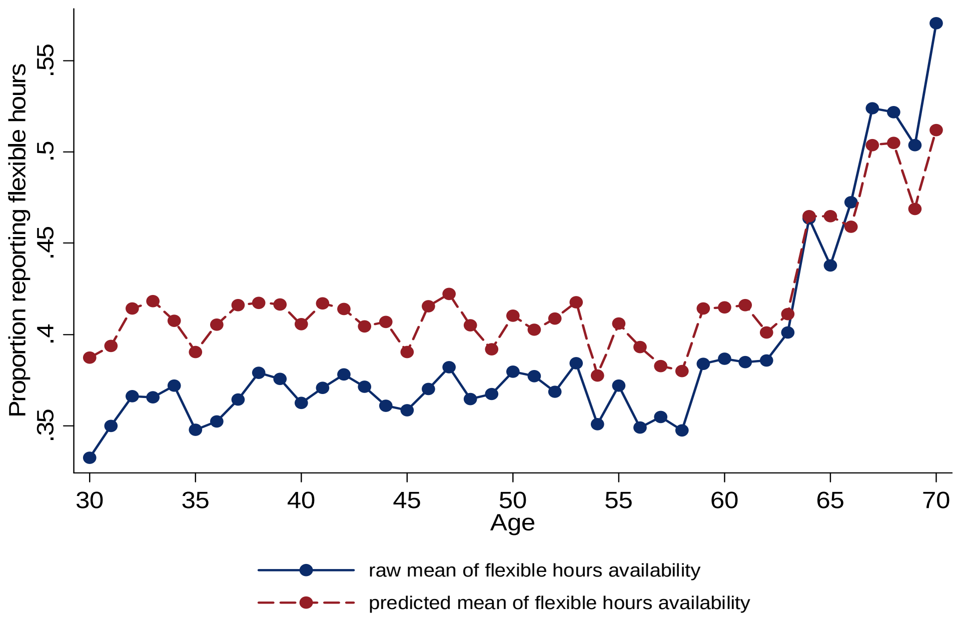
Appendix Table: Logit Parameter Estimates of the Monthly Job Separation Hazard

(standard errors in parentheses)

	Coefficient	Robust St. Err.	Continued	Coefficient	Robust St. Err.
Age 50-54	0.048	(0.101)	Repair services	0.204	(0.168)
Age 55-59	**0.299**	(0.150)	Personal services	0.027	(0.181)
Age 60-64	**0.555**	(0.203)	Recreation services	0.008	(0.190)
Age 65-69	**0.999**	(0.271)	Health services	-0.079	(0.173)
Industry-specific fraction females < 30	0.699	(0.455)	Educational services	0.058	(0.178)
Age 50-54 × industry fraction females < 30	0.307	(0.601)	Other services	0.090	(0.172)
Age 55-59 × industry fraction females < 30	-0.588	(0.660)	Public administration	0.084	(0.193)
Age 60-64 × industry fraction females < 30	0.369	(0.716)	***Occupation:***		
Age 65-69 × industry fraction females < 30	-0.184	(0.968)	Executives	-0.061	(0.060)
Employer-specific fraction females < 30	0.441	(0.330)	Professionals	**-0.233**	(0.100)
Age 50-54 × employer fraction females < 30	-0.277	(0.494)	Technicians	-0.027	(0.066)
Age 55-59 × employer fraction females < 30	-0.474	(0.543)	Sales	0.030	(0.055)
Age 60-64 × employer fraction females < 30	**-1.550**	(0.588)	Administrative support	**-0.824**	(0.310)
Age 65-69 × employer fraction females < 30	**-1.629**	(0.802)	Private household	0.017	(0.126)
Age	**-2.269**	(0.873)	Protective service	**-0.211**	(0.068)
Age squared	**0.041**	(0.016)	Farming, forestry and fishing	0.142	(0.152)
Age cubed	**-0.000**	(0.000)	Craft and repair	0.033	(0.065)
Male	-0.009	(0.041)	Machine operators	0.008	(0.074)
Black	**-0.156**	(0.064)	Transportation and material moving	0.037	(0.083)
American Indian	0.085	(0.133)	Handlers, helpers, and laborers	0.066	(0.094)
Asian	-0.154	(0.098)	***Class of worker:***		
Married, Spouse Absent	0.063	(0.153)	Private non-profit	**-0.145**	(0.070)
Widowed	0.095	(0.073)	Federal government	**-0.203**	(0.107)
Divorced	**0.153**	(0.044)	State government	-0.009	(1.116)
Separated	0.064	(0.098)	Local government	-0.129	(1.114)
Never married	**0.138**	(0.070)	Armed forces	**-1.747**	(0.441)
Education	**0.012**	(0.007)	Family business	**-2.292**	(0.671)
Real income of other household members	1.397	(1.049)	***Other employer characteristics:***		
Total household wealth	-0.001	(0.002)	Firm size <= 5 workers	**-0.452**	(0.089)
Indicator: Wealth imputed	**-0.427**	(0.098)	Firm size 6-10 workers	**-0.249**	(0.084)
Real wage	**0.006**	(0.003)	Firm size 11-25 workers	**-0.148**	(0.070)
Indicator: Wage imputed	**1.448**	(0.074)	Firm size 26-50 workers	0.010	(0.068)
Tenure	**-0.005**	(0.001)	Firm size 51-75 workers	0.014	(0.076)
Tenure squared	**0.000**	(0.000)	Firm size 76-100 workers	-0.073	(0.083)
First quarter of tenure	**0.151**	(0.051)	Firm size 101-200 workers	0.003	(0.059)
First year of tenure	**0.171**	(0.059)	Firm size 201-500 workers	0.011	(0.053)
Year 2-5 of tenure	0.061	(0.052)	Firm size 500-1000 workers	**-0.126**	(0.058)
Initial experience	**-0.006**	(0.002)	Average number of workers	**-0.040**	(0.019)
Indicator: Experience imputed	-0.043	(0.066)	Fraction of females in the firm's work force	0.013	(0.110)
Pension plan indicator	**-0.270**	(0.102)	Fraction of whites in the firm's work force	0.028	(0.113)
DB pension plan indicator	**0.186**	(0.082)	Fraction of blacks in the firm's work force	-0.005	(0.177)
Employer contributions indicator	-0.024	(0.083)	Average earnings at the firm	-4.512	(2.881)
Indicator: Pension information imputed	**1.974**	(0.047)	90th percentile of average earnings	0.626	(0.876)
Disabled	**0.397**	(0.045)	75th percentile of average earnings	-0.555	(2.115)
Bad health	-0.010	(0.047)	50th percentile of average earnings	3.868	(3.689)
Indicator: Self-reported health imputed	**-0.617**	(0.072)	25th percentile of average earnings	2.223	(4.635)
Health insurance, own name	**-0.285**	(0.051)	10th percentile of average earnings	1.454	(3.227)
Health insurance, others name	**0.087**	(0.047)	Average accession rate	**0.992**	(0.143)
Employer provided health insurance	**-0.368**	(0.048)	Multi-plant dummy	-0.029	(0.040)
Industry:			Firm age	-0.000	(0.002)
Mining	**0.449**	(0.245)	Firm age censored dummy	0.071	(0.056)
Construction	**0.296**	(0.170)	State government firm	-0.176	(0.214)
Non-durables	0.216	(0.167)	Local government firm	-0.234	(0.192)
Durables	0.199	(0.165)	Private sector firm	-0.104	(0.169)
Transportation	-0.048	(0.179)			
Public utilities	**0.365**	(0.181)	State of employment set of 31 dummies	Yes	
Wholesale trade	0.172	(0.170)	Metropolitan area indicator	0.057	(0.040)
Retail trade	-0.042	(0.169)	Time trend	**0.003**	(0.001)
Finance	-0.018	(0.171)	Constant	-31.589	(25.303)

Notes: The estimates correspond to Table 3, column 1.Initial experience is equal to the total number of months of individual's job experience as of the beginning of the SIPP coverage. 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
Raw and Predicted Means of Flexible Hours Availability by Age



Note: The predicted mean is based on a regression equation with controls for demographic characteristics, single year-of-age dummies, and detailed industry, occupation, and class of worker controls. The predicted mean holds constant all of the regressors other than the age dummies.

Source: Calculations from the May 2001 Current Population Survey.

Figure 2
Average Monthly Separation Rates by Single Year of Age

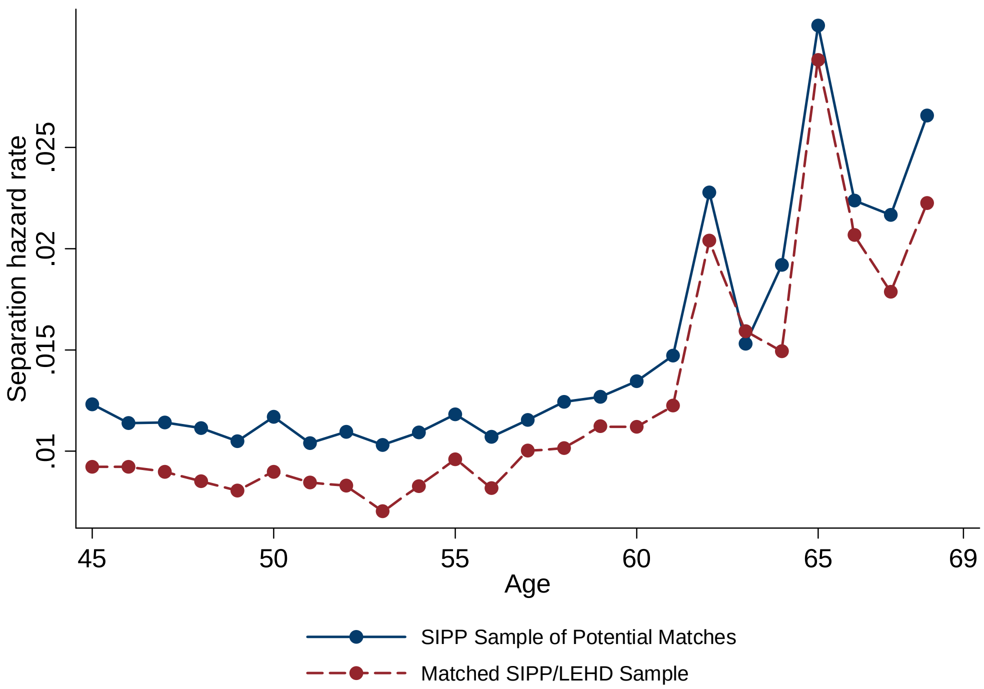


Figure 3a
Predicted Monthly Separation Rate
by Single Year of Age and Employer-Specific Fraction of Female Workers aged less than 30 years, No Other Controls

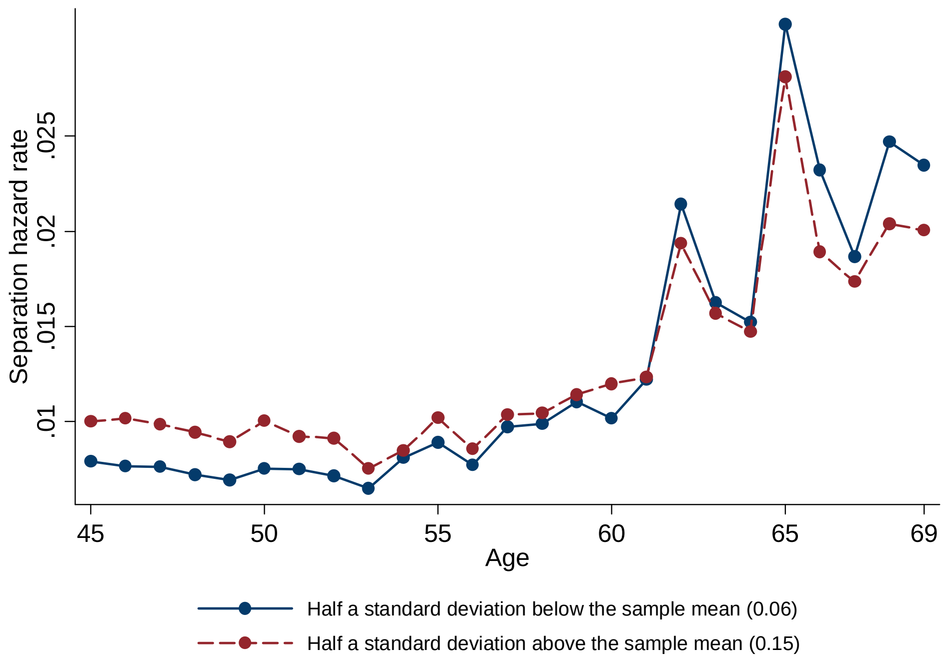


Figure 3b
Predicted Monthly Separation Rate
by Single Year of Age and Employer-Specific Fraction of 65-69 Year Old Workers, No Other Controls

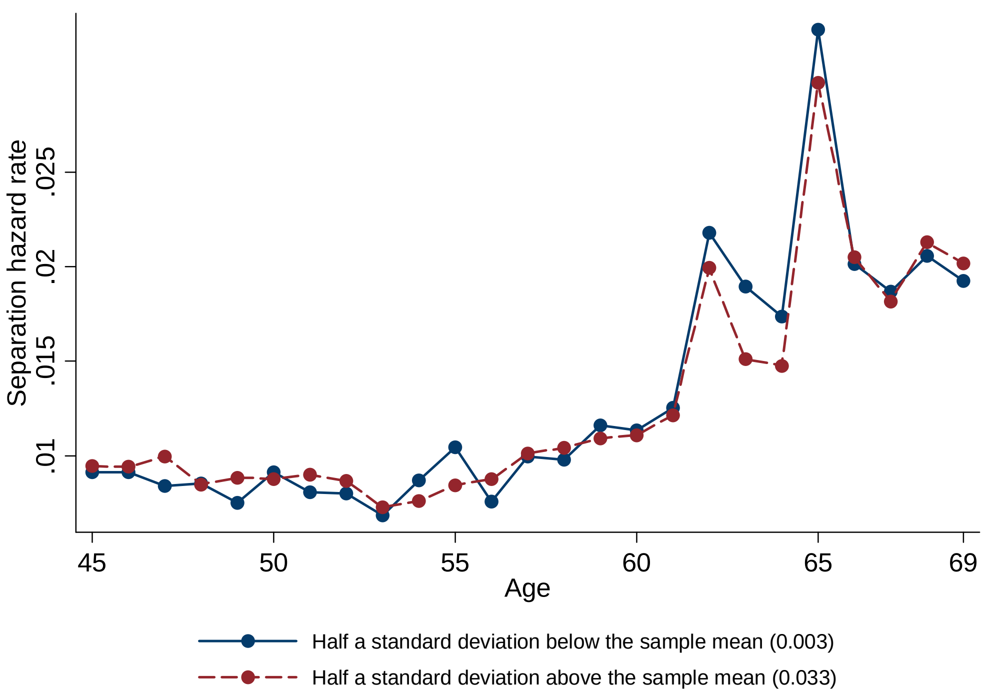


Figure 4a
Predicted Monthly Separation Rate by Single Year of Age and Employer-Specific Fraction of Female Workers aged less than 30 years with the Full Set of Control Variables

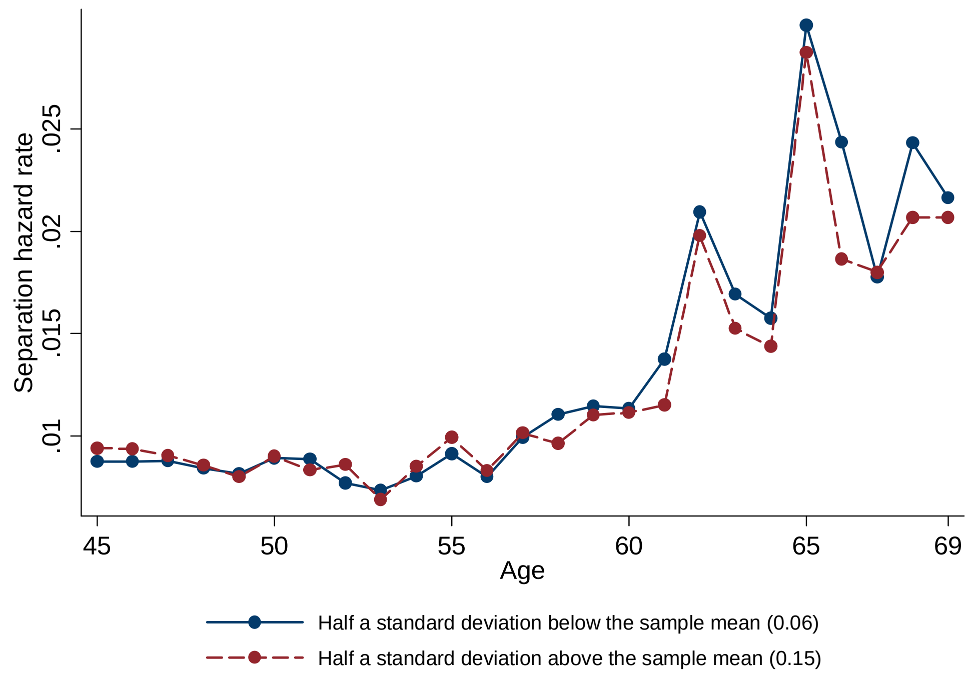


Figure 4b
Predicted Monthly Separation Rate by Single Year of Age and Employer-Specific Fraction of 65-69 Year Old Workers with the Full Set of Control Variables