

Preliminary Draft

Job Loss and the Decline in Job Security in the United States *

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Abstract

Job tenure and the incidence of long-term employment have declined sharply in the United States. However, rates of job loss as measured by the Displaced Workers Survey (DWS), while cyclical, have not increased. This presents a puzzle that has several potential solutions. One is that, while overall rates of job loss have not increased, rates of job loss for high-tenure workers have increased relative to those for lower-tenure workers. Another is that there has been an increase in rates of job change that is not captured in the limited questions asked in the DWS. Some of this seemingly voluntary job change (e.g., the taking of an offered buy-out) may reflect the kind of worker displacement that the DWS was meant to capture but is not reported as such by workers.

In this study, I address these issues by 1) documenting the decline in job tenure and long-term employment using data from various supplements to the Current Population Survey (CPS) from 1973-2006, 2) documenting the lack of secular change in rates of job loss using data from the DWS from 1984-2006, and 3) exploring the extent to which the observed patterns result from a relative increase in rates of job loss among high-tenure workers. I find that there has been no such relative change and that reconciliation of the trends in the tenure and displacement data must lie with a failure to identify all relevant displacement in the DWS.

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1 Introduction

There is ample evidence that long-term employment is on the decline in the United States. The common understanding that, after some turnover early in careers, most workers find a job (relationship with an employer) that lasts for a long period of time (a “life-time” job), has been challenged in the last fifteen to twenty years, both in academic research and in the media, as large corporations have engaged in highly publicized layoffs and the industrial structure of the U.S. economy has shifted in the face of global competitive pressures. However, there is little evidence that rates of job loss have increased. This leaves a puzzle regarding the mechanism through which long-term employment relationships are becoming less common.

One possible explanation is that there has been an increase in the rate of job change by workers that is not captured by the DWS. This could be in the form of voluntary job change. accounts for the decline in job durations. The interpretation what is a “job loss” and what is a “voluntary” job change is left to the respondent in the DWS. The result may be an underestimate of employer-initiated separations. For example, a firm may offer workers “buy-outs.” Workers who take these buy-outs in lieu of a layoff may not report a job loss in the DWS. Unfortunately, there are no large-scale surveys that measure job change and reasons for job change generally.

Another possible explanation is that, while the overall rate of job loss has not increased, higher tenure workers have become more susceptible to job loss. In this study, I address the validity of this explanation by 1) documenting the decline in job tenure and long-term employment using data from various supplements to the Current Population Survey (CPS) from 1973-2006, 2) documenting the lack of secular change in rates of job loss using data from the DWS from 1984-2006, and 3) exploring the extent to which the observed patterns result from a relative increase in rates of job loss among high-tenure workers. I find that there has been no such relative change and that reconciliation of the trends in the tenure and displacement data must lie with a failure to identify all relevant displacement in the DWS.

2 Background and Earlier Literature

The evolution of the job durations in the U.S. has played out in the context of dramatic growth in employment over the last 40 years. Civilian employment was 85.1 million in 1973 and rose to 144.4 million in 2006.¹ Thus, almost sixty million jobs have been created on net in the past 33 years, for an average rate of employment growth of 1.6 percent per year over this period. Despite this record of sustained growth in employment in the United States, there is longstanding concern that the quality of the stock of jobs in the economy more generally is deteriorating. The concern about job quality is based in part on the fact that the share of employment that is in manufacturing has been declining over a long period of time.² This has led to the view that, as high-quality manufacturing jobs are lost, perhaps to import competition, they are being replaced by low-quality service sector jobs (so-called hamburger-flipping jobs). The high-quality jobs are characterized by relatively high wages, full-time employment, substantial fringe benefits, and, perhaps most importantly, substantial job security (low rates of turnover). The low-quality jobs are characterized disproportionately by relatively low wages, part-time employment, an absence of fringe benefits, and low job security (high rates of turnover).

The perceived low quality of many newly-created jobs fuels the concern that the nature of the employment relationship in the United States is changing from one based on long-term full-time employment to one based on more short-term and casual employment is. There has been concern that employers are moving toward greater reliance on temporary workers, on subcontractors, and on part-time workers. Potential motivation for employers to implement such changes range from a need for added flexibility in the face of greater uncertainty regarding product demand to avoidance of increasingly expensive fringe benefits and long-term obligations to workers. The public's concern arises from of the belief that these changes result in lower quality (lower paying and less secure) jobs for the average

¹ These statistics are taken from U.S. Bureau of Labor Statistics Series ID LNU02000000. This is the civilian employment level derived from the Current Population Survey for workers aged 16 and older.

² The manufacturing share of non-farm employment has been falling for over fifty years. Manufacturing's share was 30.9 percent in 1950 and fell to 10.4 percent in 2006. These statistics are taken from U.S. Bureau of Labor Statistics Series ID CEU00000001 and CEU30000001 derived from the Current Employment Statistics payroll data.

worker.

2.1 Literature on Job Stability

There have been a series of analyses of job stability that have relied on mobility supplements to various January Current Population Surveys. An influential early analysis was carried out by Hall(1982). He used published tabulations from some of the early January mobility supplements to compute contemporaneous job retention rates. Hall found that, while any particular new job is unlikely to last a long time, a job that has already lasted five years has a substantial probability of lasting twenty years. He also finds that a substantial fraction of workers will be on a “lifetime” job (defined as lasting at least twenty years) at some point in their life. Ureta (1992) used the January 1978, 1981, and 1983 mobility supplements to recompute retention rates using artificial cohorts rather than contemporaneous retention rates.

Several more recent papers have used CPS data on job tenure to examine changes in employment stability. Swinnerton and Wial (1995), using data from 1979 through 1991, analyze job retention rates computed from artificial cohorts and conclude that there has been a secular decline in job stability in the 1980’s. In contrast, Diebold, Neumark, and Polsky (1994), using CPS data on tenure from 1973 through 1991 to compute retention rates for artificial cohorts, find that aggregate retention rates were fairly stable over the 1980’s but that retention rates declined for high school dropouts and for high school graduates relative to college graduates over this period. I interpret a direct exchange between Diebold, Polsky, and Neumark (1996) and Swinnerton and Wial (1996) as supporting the view that the period from 1979-91 is not a period of generally decreasing job stability. Farber (1998), using CPS data on job tenure from 1973 through 1993, finds that the prevalence of long-term employment has not declined over time but that the distribution of long jobs has shifted. He finds that less-educated men are less likely to hold long jobs than they were previously but that this is offset by a substantial increase in the rate at which women hold long jobs. Farber (2000) examines CPS data on job tenure from 1973 through 1996, and he finds that the prevalence of long- term employment relationships among men declined by 1996 to its lowest level since 1979. In contrast, long-term employment relationships became somewhat

more common among women.

Rose (1995) uses data from the Panel Study of Income Dynamics (PSID) to measure job stability by examining the fraction of male workers who do not report any job changes in a given time period, typically ten years. Rose finds that the fraction of workers who reported no job changes in given length of time was higher in the 1970's than in the 1980's. He argues that this is evidence of increasing instability of employment.

The Russell Sage Foundation sponsored a conference organized by David Neumark on "Changes in Job Stability and Job Security" in 1998.³ The evidence presented here is mixed regarding whether job tenure was declining. Jaeger and Stevens (1999) use data from the PSID and the CPS mobility and benefit supplements on (roughly) annual rates of job change to try to reconcile evidence from the CPS and PSID on job stability. They find no change in the share of males in short jobs and some decline between the late 1980s and mid-1990s in the share of males with at least ten years of tenure.⁴ Neumark, Polsky, and Hansen (1999) find a similar decline in long-term employment but conclude that this does not reflect a secular trend. Gottschalk and Moffitt (1999) use monthly data from the Survey of Income and Program Participation (SIPP) along with annual data from the SIPP and the PSID, and they find no evidence of an upward trend in job insecurity in the 1980s and 1990s. Valletta (1999) uses data from the PSID from 1976-1993 and finds some decline in long-term employment relationships.

In more recent work, Stewart (2002) uses data from the March CPS to investigate two aspects of job security. The first, the likelihood of leaving a job, shows no particular trend from 1975 through 2000 based on these data. The second, the likelihood of making an employment-to-employment transition, increased over this period while the likelihood of making an employment-to-unemployment transition decreased. Stewart concludes that the cost of changing jobs has decreased.

Stevens (2005) examines data from several longitudinal histories of older male workers (late 50s and early 60s) with regard to changes over time in the length of longest job held

³ The Proceedings of this conference are published in Neumark (2000), and a number of these papers are published in *The Journal of Labor Economics* Volume 17, Number 4, Part 2, October 1999

⁴ Unfortunately, due to the design of the PSID, neither of these studies examine the mobility experience of women.

during careers. She finds that there has been no change between the late 1960s and late early 2000s and concludes that there has not been a decline in the incidence of “lifetime jobs”. A careful reading of her results show an increase in average longest tenure from about 22 years among older workers in 1969 to 24 years in 1980 followed by a decline to 21.4 years in 2002. A reasonable interpretation of this pattern is that the earliest cohorts had jobs interrupted by service in World War II, resulting in lower average longest tenure than subsequent cohorts. The decline since 1980 may then reflect a real decline in job durations. Additionally, the most recent cohort examined by Stevens was born in the 1940s so that her analysis cannot shed light on the experience of more recent birth cohorts.

Farber (2007) uses data from 21 supplements to the Current Population Survey (CPS) over the 1973-2006 period that contain information on how long workers have been employed by their current firm. He finds that, by virtually any measure, more recent cohorts of male workers have been with their current employers for less time at specific ages. He does not find a corresponding decline in age-specific tenure for women. This contrast reflects the increased commitment of women to the labor force tempered by the fact that many working women, when have young children, either exit the labor force for a period of time or change jobs to one with different or more flexible hours.

Taken as a whole, I conclude from this earlier literature that there has been a decline in job tenure and in the incidence of long-term employment relationships.

2.2 Literature on Job Loss

In an earlier paper (Farber 1993), I used the five DWSs from 1984 to 1992 to examine changes in the incidence and costs of job loss over the period from 1982-1991. I found that there were slightly elevated rates of job loss for older and more educated workers in the slack labor market in the latter part of the period compared with the slack labor market of the earlier part of the period. But I found that job loss rates for younger and less educated workers were substantially higher than those for older and more educated workers throughout the period. These findings are consistent with the long-standing view that younger and less educated workers bear the brunt of recessions. I also confirmed the conventional view that the probability of job loss declines substantially with tenure.

Gardner (1995) carried out the first analysis of which I am aware that incorporated the 1994 DWS. She examined the incidence of job loss from 1981-92. While she found roughly comparable overall rates of job loss in the 1981-82 and 1991-92 periods, she found that the industrial and occupational mix of job loss changed over this period. There was an decreased incidence of job loss among blue-collar workers and workers in manufacturing industries and an increase in job loss among white-collar workers and workers in non-manufacturing industries.

In another paper (Farber 1997), I used the seven DWSs from 1984 to 1996 to revisit the issue of changes in the incidence and costs of job loss. I found that the overall rate of job loss increased in the first half of the 1990s despite the sustained economic expansion. Hipple (1999) carried out the first analysis of the 1998 DWS, and he finds that the displacement rate among workers who had held their jobs for at least three years fell only slightly between the 1993-1994 period and the 1995-1996 period despite the sustained economic expansion.

There is a substantial literature using the DWS to study the post-displacement employment and earnings experience of displaced workers.⁵ This work demonstrates that displaced workers suffer substantial periods of unemployment and that earnings on jobs held after displacement are substantially lower than pre-displacement earnings. In my earlier work (Farber 1993), I found that there was no difference on average in the consequences of job loss between the 1982-83 recession and the the 1990-91 recession.

The earnings loss suffered by displaced workers is positively related to tenure on the pre-displacement job. On the other hand, Kletzer (1989) found further that the post-displacement earnings *level* is positively related to pre-displacement tenure, suggesting that workers displaced from long jobs are more able on average than those displaced from shorter jobs. In more recent work, Neal (1995) using the DWS and Parent (1995) using the National Longitudinal Survey of Youth (NLSY) found that workers who find new employment in the same industry from which they were displaced earn more than do industry switchers. This work suggests that Kletzer's finding that post-displacement earnings are positively related to pre-displacement tenure may be a result of the transferability of industry-specific capital.

⁵ See, for example, Podgursky and Swaim (1987), Kletzer (1989), Topel (1990), Farber (1993), Farber (1997).

Workers who are re-employed in the same industry “earn a return” on their previous tenure while those re-employed in a different industry do not.

In Farber (2004) I examined changes in the incidence and consequences of job loss between 1981 and 2001 using data from the Displaced Workers Surveys (DWS) from 1984-2002. I found that the overall rate of job loss has a strong counter-cyclical component but that the job loss rate was higher than might have been expected during the mid-1990’s given the strong labor market during that period. I found substantial earnings declines for displaced workers relative to what they earned before displacement. Additionally, foregone earnings growth (the growth in earnings that would have occurred had the workers not been displaced), is an important part of the cost of job loss for re-employed full-time job losers. There is no evidence of a decline during the tight labor market of the 1990s in the earnings loss of displaced workers who were reemployed full-time. In fact, earnings losses of displaced workers have been increasing since the mid 1990s. In Farber (2005), I update my earlier work to include data on job loss through 2003. Not surprisingly, there were higher job loss rates and lower post-displacement reemployment probabilities during the recession of the early 2000s.

With regard to overall rates of job loss, this literature suggests that job loss rates have a strong cyclical component. However, aside from several years with unusually high job-loss rates in the mid-1990s, there has been no secular increase in rates of job loss.

3 The Decline In Long-Term Employment

In this section, I present evidence on job durations from a sample consisting of not-self-employed workers aged 20-64 from the 21 CPS supplements covering the period from 1973 to 2006. The sample contains 876,063 workers. These data are described in more detail in Appendix I and in Farber (2007).

3.1 Measuring the Change in Tenure Over Time

I organize my analysis of changes over time in the distribution of job durations by examining age-specific values of various distributional measures of job tenure for each sampled year.

No one statistic can completely characterize a distribution, and I focus on several measures here:

- Mean job tenure (years with the current employer). Note that this is not mean completed job duration since since the jobs sampled are still in progress.
- The age-specific probability that a worker reports being on their job at least ten years. Because younger workers cannot have accumulated substantial job tenure, I restrict this analysis to workers at least 35 years of age, and I examine how these probabilities have changed over time.
- The age-specific probability that a worker reports being in their job at least twenty years. Because younger workers cannot have accumulated substantial job tenure, I restrict this analysis to workers 45 years of age and older
- The age-specific probability that a worker reports being their job for less than one year. This provides information on changes over time in the transition from the early job-shopping phase of a career to more stable longer-term employment relationships.

An important measurement issue is related to cyclical changes in the composition of the sample. It is clear that workers with little seniority are more likely than high-tenure workers to lose their jobs in downturns (Abraham and Medoff, 1984). Thus, we would expect that the incidence of long-term employment, as measured by the fraction of workers with tenure exceeding some threshold, to be counter-cyclical. Tight labor markets will lead the distribution of job durations to lie to the left of the distribution in slack labor markets, and these cyclical influences need to be kept in mind when interpreting the results.

3.2 Mean Tenure

Figure 1 contains separate plots by sex of mean tenure by age for the four decades from 1973 through 2006.⁶ These figures show clearly that 1) mean tenure is rising with age and 2) women have lower mean tenure than men after about age 30. With regard to shifts over

⁶ Means are calculated weighted by CPS final sample weights.

time in the tenure distribution, age-specific mean tenure for males has declined substantially, particularly for older workers. For example mean tenure for males at age 50 declined from 13.6 years in the 1970s to 11.8 years in the early 2000s. In contrast, age-specific mean tenure for older women has increased. Mean tenure for females at age 50 increased from 8.9 years in the 1970s to 9.7 years in the early 2000s. This reflects the increased attachment to the labor force of more recent cohorts of women.

Another approach to summarizing the data is to estimate a linear model of the natural logarithm of tenure of the form

$$\ln(T_{ijt}) = Y_t + A_j + \epsilon_{ijt}, \quad (3.1)$$

where T_{ijt} is tenure in years for individual i at age j in year t , Y_t is a calendar year indicator, and A_j is a years-of-age indicator. This logarithmic specification embodies the plausible implicit assumption that proportional year effects on mean tenure are constant across ages and, equivalently, that the proportional age effects on mean tenure are constant across years.⁷ A more detailed investigation would allow for year effects that vary by age since changes in job security could express themselves differentially at various ages. However, the model in equation 3.1 fits the data quite well, and it serves as a good summary of the data.⁸

I estimate the model in equation 3.1 separately for men and women using ordinary least squares (OLS), weighted by the CPS final sample weights. The estimated year effects on mean tenure, normalized at zero in 1973, are converted to proportional differences in mean tenure relative to 1973 as $\exp(\hat{Y}_t - \hat{Y}_{1973}) - 1$. These proportional differences are plotted in figure 2, and they show a sharp decline of about 20 percent in age-specific mean tenure for male workers between the 1973 and 2006.

The time-series pattern is quite different for female workers. Age-specific mean tenure for female workers rose from 1980 through the mid-1990s and subsequently fell. The increase in

⁷ I do not estimate this model using absolute tenure because the implicit assumption in that case would be that absolute year effects on mean tenure are constant across ages and, equivalently, that absolute age effects on mean tenure are constant across years. This is clearly not plausible on inspection of figure 1, given the fact that younger workers have very low levels of tenure.

⁸ I computed (separately for men and women) weighted mean tenure for each age/year combination and regressed these measures on a complete set of age and year fixed effects. This is essentially the main-effects model in equation 3.1 aggregated to the cell level. The R-squared from the mean regression is 0.99 for men and 0.98 for women.

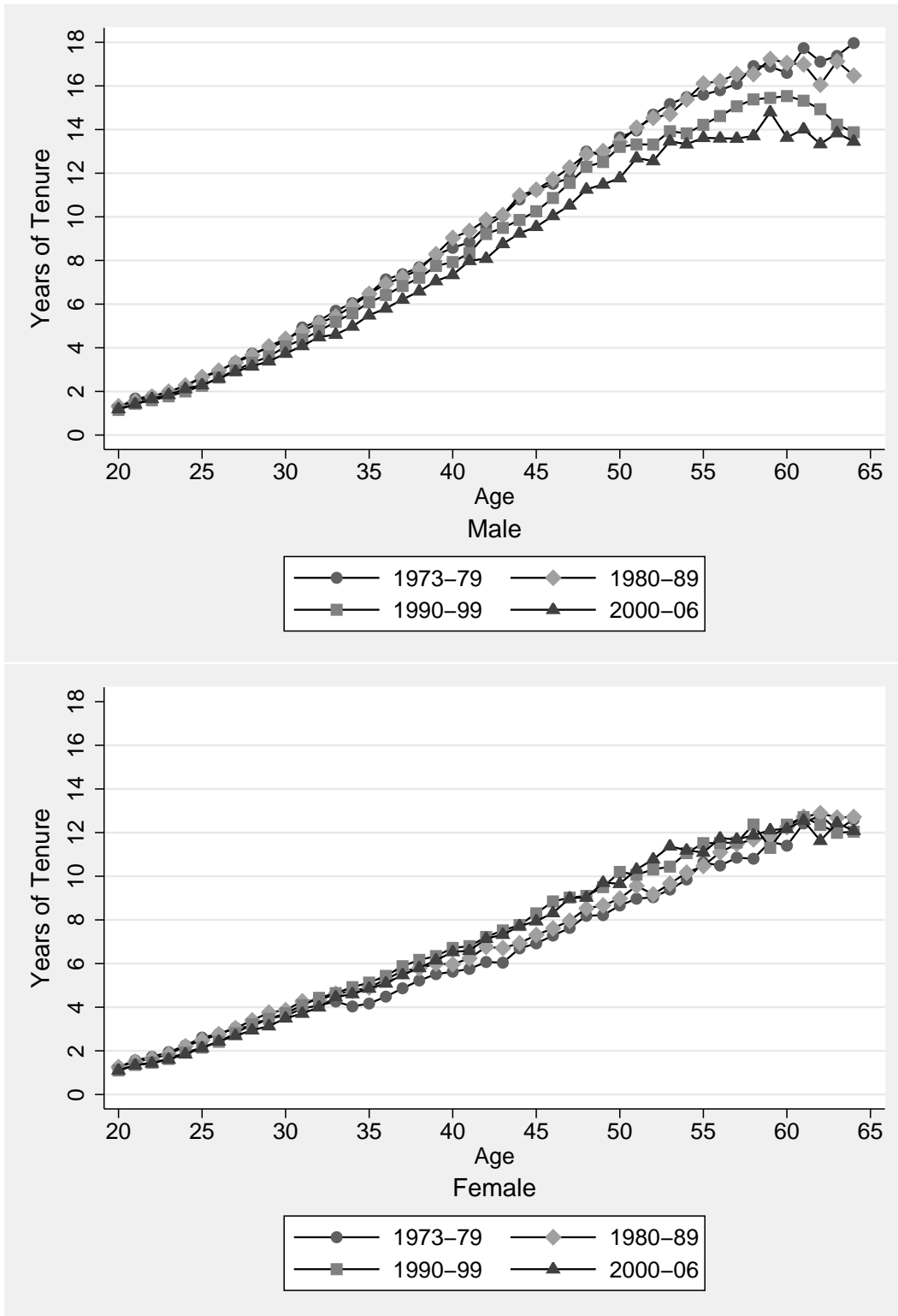


Figure 1: Mean Tenure, by Sex, Age, and Year

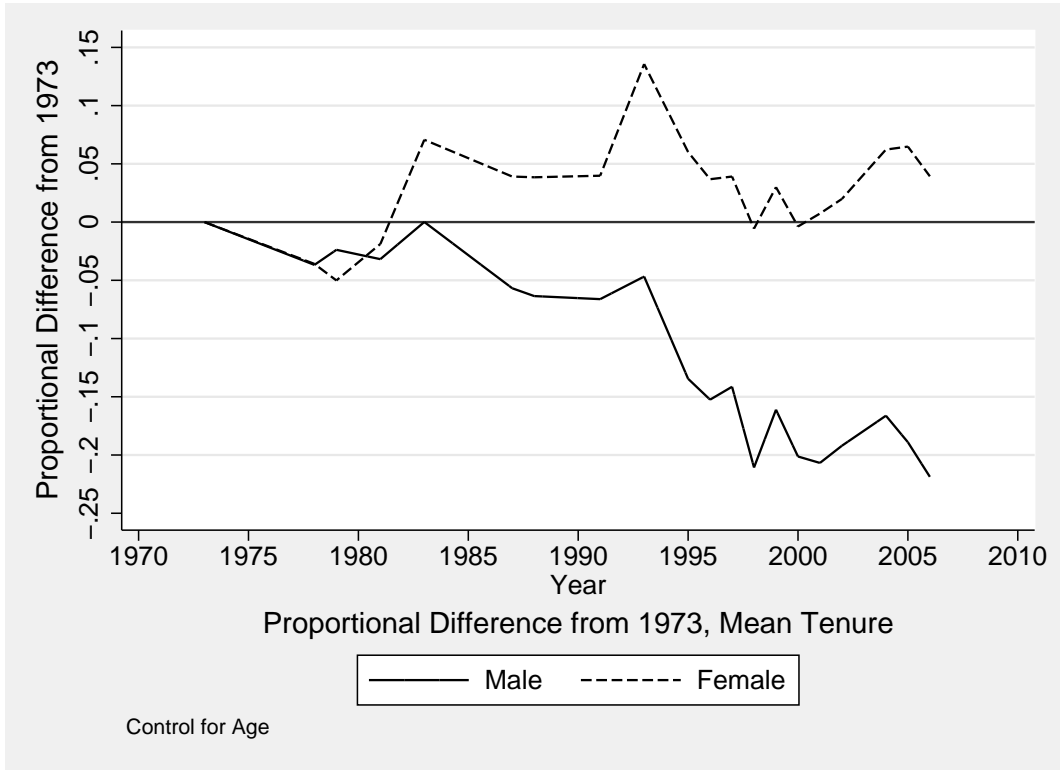


Figure 2: Proportional Difference from 1973, Mean Tenure, Controlling for Age.

mean tenure for women between reflects the increased commitment of women to the labor force for women born in this period tempered by 1) high rates of withdrawal from the labor force, even if only for a short time, in the child-bearing years and 2) the general decline in long-term employment opportunities apparent in the data for males. The subsequent decline in tenure for females may reflect a continued general decline in long-term employment opportunities that is not offset by a further increase in female commitment to the labor force.

In addition to the increased presence of women in the labor force, there are other important changes that could be related to the decline in tenure. First is the well-known large increase in average educational attainment during the 20th century. While there is not a clear relationship between educational attainment and tenure, I investigate how the decline in mean job tenure is related to the general increase in educational attainment.⁹

⁹ Mean tenure in my analysis sample for each of the four educational categories are ED<12: 7.3 years, ED=12: 7.3 years, ED 13-15: 6.4 years, and ED ≥ 16: 7.3 years.

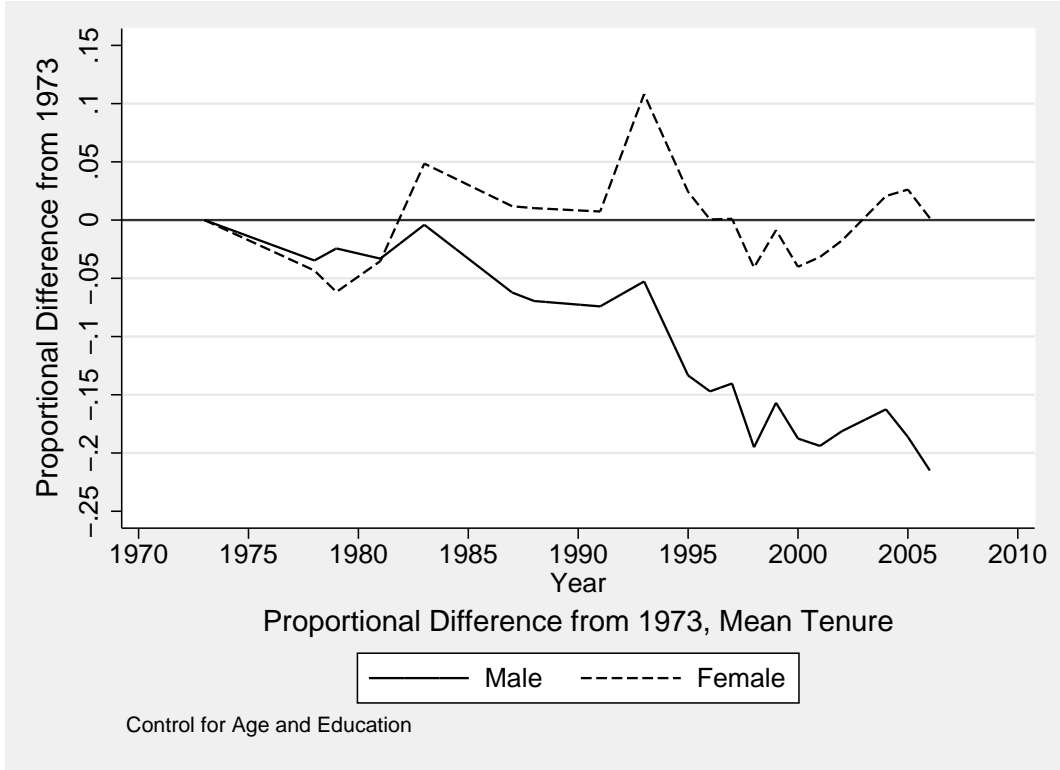


Figure 3: Proportional Difference from 1973, Mean Tenure. Controlling for Age and Education.

In order to provide a summary across educational categories of the proportional change in mean tenure over time accounting for changes in the educational distribution over time, I estimate an augmented version of the regression model for mean tenure in equation 3.1 as

$$\ln(T_{ijt}) = ED_i\gamma + Y_t + A_j + \epsilon_{ijt}, \quad (3.2)$$

where ED_i is a vector of dummy variables indicating educational attainment and γ is a vector of associated coefficients. The estimated proportional change in mean tenure relative to 1973 ($\exp(\hat{Y}_t - \hat{Y}_{1973}) - 1$) are plotted in figure 3, and, while they are very similar in shape to those derived without controlling for education (figure 2), there are some differences for females. Accounting for changes in the distribution of education, the estimated increase in mean tenure for females between 1973 and the mid 1990s was about 10 percent compared with an increase of about 14 percent without accounting for differences in education.

Table 1: Proportion Immigrants by Race and Hispanic Ethnicity, 1995-2006

Year	All	White NonHisp	Nonwhite NonHisp	All Hisp	White Hisp	Nonwhite Hisp
1995	0.095	0.030	0.187	0.506	0.509	0.492
1996	0.100	0.032	0.226	0.494	0.493	0.510
1997	0.109	0.032	0.232	0.516	0.518	0.484
1998	0.117	0.035	0.240	0.517	0.516	0.526
1999	0.111	0.033	0.222	0.495	0.498	0.448
2000	0.121	0.038	0.239	0.517	0.514	0.585
2001	0.129	0.039	0.261	0.522	0.520	0.557
2002	0.130	0.040	0.270	0.528	0.527	0.543
2004	0.142	0.042	0.280	0.531	0.538	0.439
2005	0.141	0.037	0.275	0.538	0.545	0.439
2006	0.147	0.039	0.282	0.550	0.556	0.469
All	0.122	0.036	0.249	0.521	0.524	0.491

Note: Based on data for not self employed workers 20-64 years of age from 11 CPSs covering the period from 1995 to 2006. Weighted by CPS final sample weights. N=467,399.

A second and potentially more important factor that could account for the decline in tenure is the increased presence of immigrants in the U.S. labor force. By definition, newly arrived immigrants cannot have substantial tenure. Data on immigration are not available in any CPS with tenure data prior to 1995. Analysis of the data since 1995 illustrates both the sharp increase in immigrant share in the labor force and the fact that immigrants have lower job tenure than natives. The weighted immigrant fraction of the labor force in my sample increased steadily from 9.45 percent in 1995 to 14.7 percent in 2006. On average between 1995 and 2006, immigrants had 2.14 years lower tenure than natives (s.e. = 0.034). Immigrants were only slightly younger than natives over the same period (average difference = 0.98 years (s.e. = 0.050)).¹⁰

An important question is how much of the decline in observed tenure since 1973 is due to the increased immigrant presence in the labor force. While not directly observable prior to 1995, immigrant status is strongly correlated with race and Hispanic ethnicity, which is observed in all years. Table 1 contains the immigrant proportion by race and Hispanic ethnicity for the 1995-2005 CPS data. The overall immigrant proportion of workers rose from 9.5 percent in 1995 to 14.7 percent in 2006. These immigrants are highly concentrated

¹⁰ See Farber (2007) for a detailed analysis of the change in job tenure since 1995 that controls directly for immigrant status.

among nonwhites and Hispanics. Only 3.6 percent of white non-Hispanics are immigrants, while over fifty percent of Hispanics (white and nonwhite) are immigrants.¹¹ Additionally, a growing fraction of nonwhite non-Hispanics are immigrants, rising from 18.7 percent in 1995 to 28.2 percent in 2006. The rising overall immigrant share over this period is reflected in the growing share of Hispanics and nonwhites in the labor force. The Hispanic share of employment in my sample increased from 9.0 percent in 1995 to 13.4 percent in 2006 and the nonwhite share of employment increased from 15.2 percent to 17.2 percent over the same period.

In order to account, at least partly, for the role of increased immigration in the decline in tenure, I estimate age-specific proportional differences in mean tenure relative to 1973 controlling for race and ethnicity as well as age and education. I derive the year effects by estimating

$$\ln(T_{ijt}) = \alpha_1 NW_i + \alpha_2 H_i + ED_i \gamma + Y_t + A_j + \epsilon_{ijt}, \quad (3.3)$$

where NW_i is an indicator for nonwhite and H_i is an indicator for Hispanic ethnicity.

Figure 4 contains separate plots for males and females of the proportional differences from 1973 in mean tenure based on equation 3.3. The estimated differences for males show a decline in age-specific tenure of about 18 percent between the 1973 and 2006. This contrasts with an estimated decline over the same period of about 21.5 percent when there are no controls for race and Hispanic ethnicity (figure 3). Thus, about 16 percent of the decline in age-specific average tenure is related to changes in racial and ethnic composition. Age-specific average tenure for females, estimated when race and ethnicity are accounted for in figure 4, peaks 1993 and returns to the 1973 level by 2006. This pattern is very close to the pattern for females shown in figure 3 without controls for race and ethnicity.

It is clear from the analysis in this sub-section that age-specific mean tenure has declined dramatically over time and that only one sixth of this decline can be accounted for by the sharp growth in immigrants in the labor market. This decline is concentrated among men and is less apparent among women, where labor supply considerations are more important.

¹¹ The rather sharp drop in the immigrant proportion among nonwhite Hispanics is due to the change in the race identification coding in the CPS in 2004.

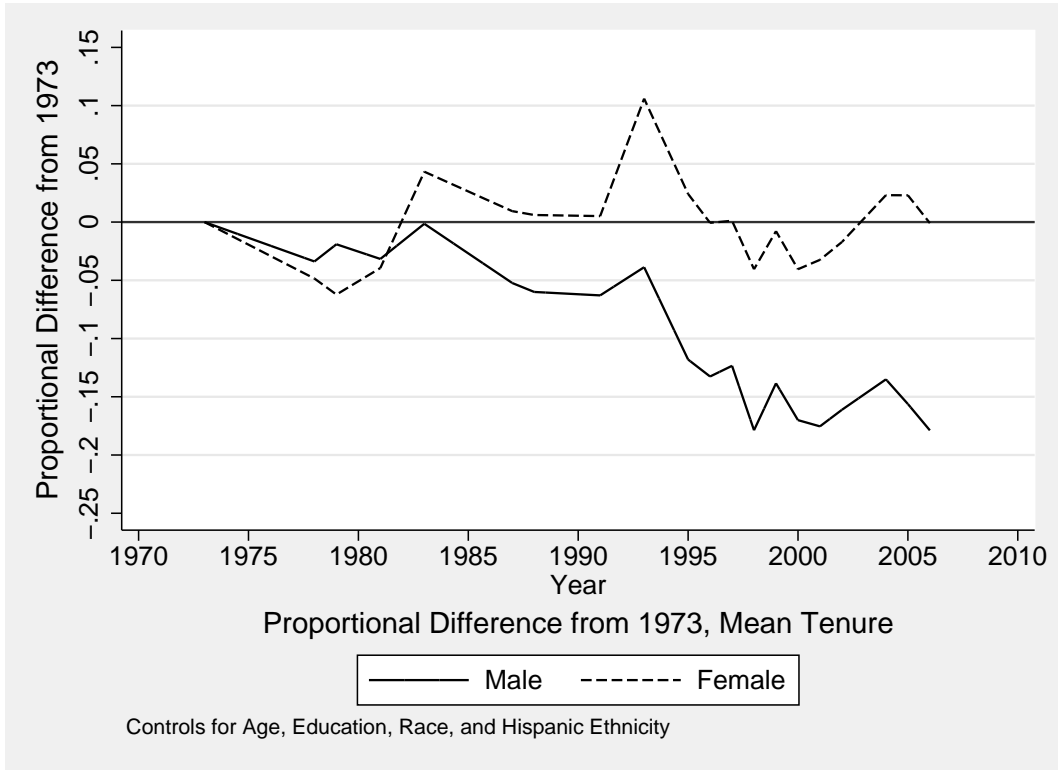


Figure 4: Proportional Difference from 1973, Mean Tenure. Controlling for Age, Education, Race, and Hispanic Ethnicity.

3.3 Long-Term Employment

Long-term employment is common in the U.S. Labor Market. In this analysis I consider two measures of long-term employment:

- the fraction of workers aged 35-64 who have been with their employer at least ten years (the “10-year rate”), and
- the fraction of workers aged 45-64 who have been with their employer at least twenty years (the “20-year rate”).

Figure 5 contains plots of these two measures over the 1973-2006 period for men and women. It is clear that the incidence of long term employment has declined dramatically for men, with the 10-year rate falling from about 50 percent to less than 40 percent and the 20-year rate falling from about 35 percent to less than 25 percent between 1973 and 2006. In

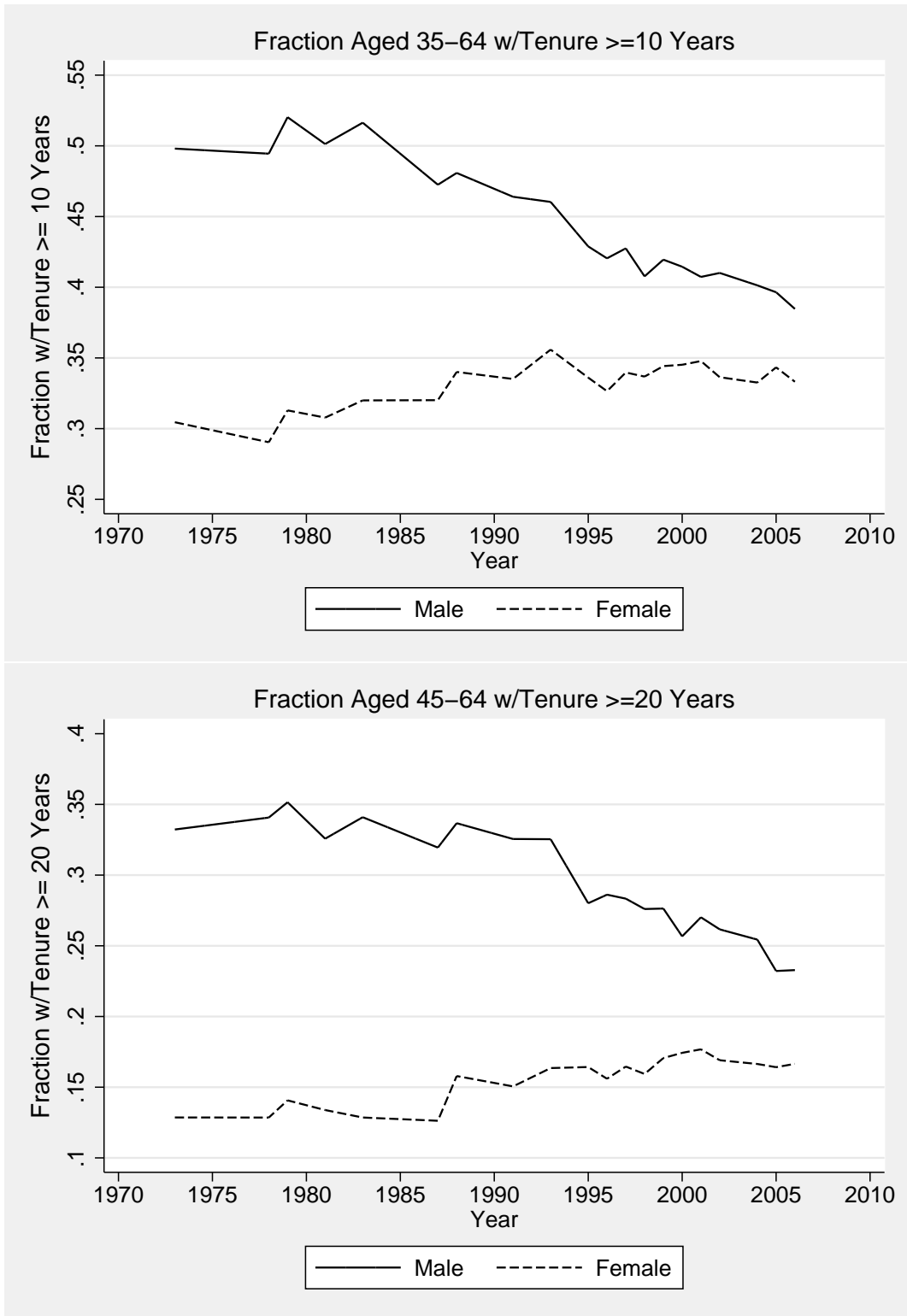


Figure 5: Fraction of Workers in Long Term Jobs, by Year.

contrast the incidence of long-term employment increased among women, with the 10-year rate increasing from about 30 percent to about 34 percent and the 20-year rate increasing from about 13 percent to about 17 percent between 1973 and 2006.

Because these measures are sensitive to the age distribution and other observable characteristics, I estimate age-specific year effects using the same approach I used for mean tenure. I estimate linear probability models using the same specification of explanatory variables (year, age, education, race, Hispanic ethnicity) in equation 3.3, and I report the estimated year effects from this analysis in figure 6.

The top panel of figure 6 contains separate plots for males and females of year effects (1973=0) for the 10-year rate. The age-specific probability that a male worker has been with his employer for at least ten years increased by about 4 percentage points between 1973 and the early 1980s before falling by about 14 percentage points subsequently. The age-specific probability that a female worker has been with her employer for at least ten years increased by 7 percentage points between 1973 and 1993 before falling by 5.5 percentage points by 2006.

The bottom panel of figure 6 contains separate plots for males and females of the year effects (1973=0) for the 20-year rate. The age-specific probability that a male worker has been with his employer for at least twenty years was roughly constant between 1973 and 1993 but fell by about 8.5 percentage points between 1993 and 2006. The pattern for females shows no change in the 20-year rate between 1973 and 1987 before increasing by about 4 percentage points between 1987 and 2001.

Taken together, the analysis of the changes in average tenure (figure 4) and in the likelihood of long-term employment (figure 6) across cohorts shows clearly that average tenure has declined and long-term employment has become much less common for males. Among females, average tenure increased slightly from the 1973 through 1993 before declining. The likelihood of long-term employment for females grew slowly over the sample period.

This difference in patterns between males and females reflects the common factors reducing tenure for all workers offset for females by their dramatically increased commitment to the labor force over the past half century. As that increase in commitment among women has slowed down average tenure and the incidence long-term employment among women

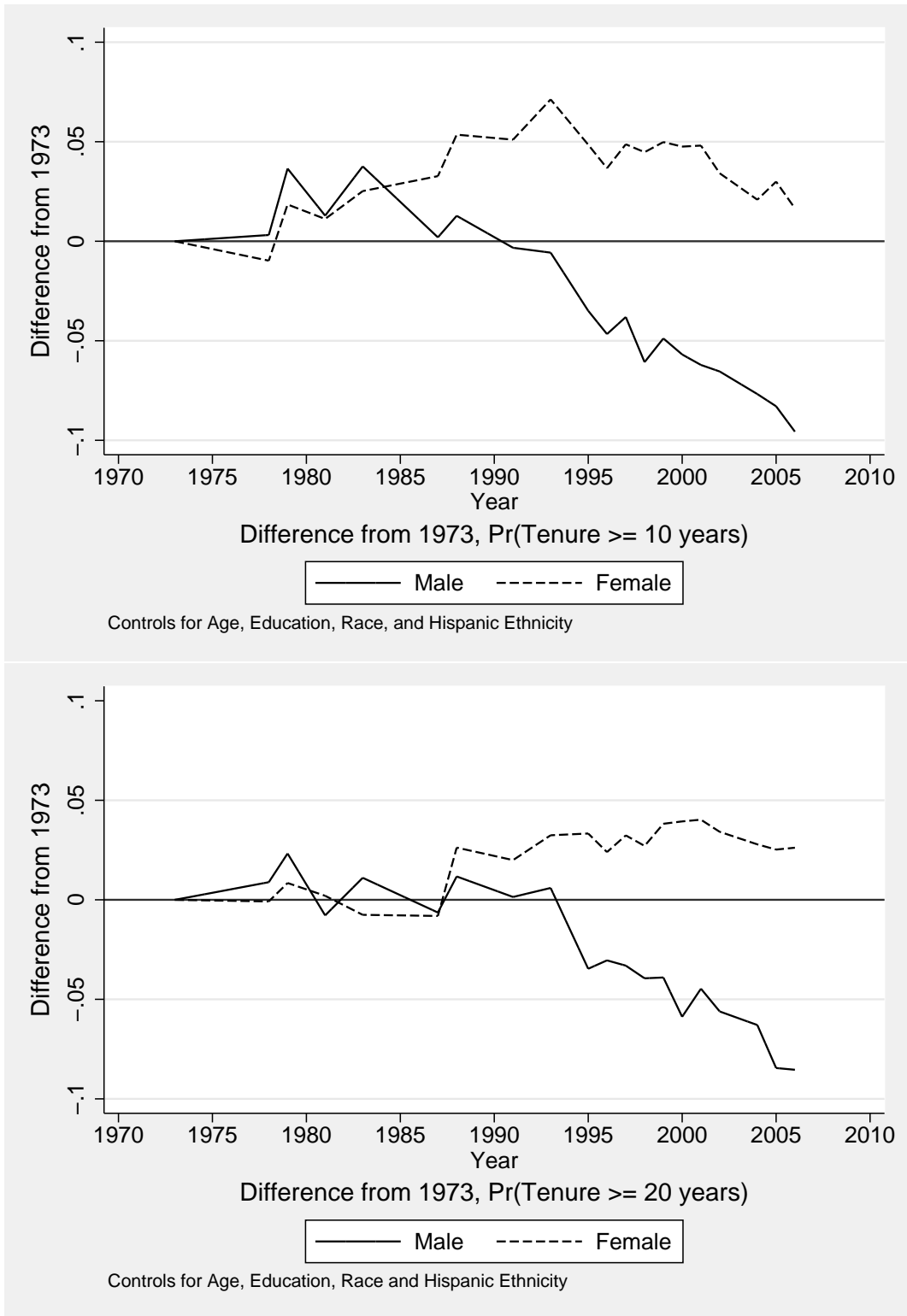


Figure 6: Proportional Difference from 1973, $Pr(T \geq 10)$ and $Pr(T \geq 20)$. Controlling for Age, Education, Race, and Hispanic Ethnicity.

have begun to decline.

A key conclusion is that the structure of employment in the United States appears to have become less oriented toward long-term jobs. It appears that young workers today will be less likely than their parents to have a “life-time” job.

3.4 Churning: Are There More Very Short Jobs?

The opposite but related pole of the job tenure distribution is short-term jobs. Farber (1994, 1999) presents evidence that half of all new jobs (worker-employer matches) end within the first year. As I show below, a substantial fraction (around 20 percent) of all jobs have current tenure less than one year (“new jobs”). Not surprisingly, young workers are more likely than older workers to be in new jobs. High rates of job change among young workers are a natural result of search for a good job or a good match.¹²

Table 2 contains the new-job rate by ten-year age group for males and females. This illustrates the sharp decline in the new-job rate as workers age through their twenties especially and into their thirties. The new-job rate is slightly higher for females in all age groups, but the general pattern is the same as that for males.

Table 2: New Job Rate, by Sex, 1973-2005

Age	All	Male	Female
Age 20-29	0.349	0.335	0.365
Age 30-39	0.181	0.162	0.205
Age 40-49	0.124	0.111	0.139
Age 50-59	0.090	0.084	0.097
Age 60-64	0.077	0.075	0.079
All	0.191	0.176	0.206

Note: Based on data for not self employed workers 20-64 years of age from 19 CPSs covering the period from 1973 to 2006. Weighted by CPS final sample weights. N=876,063.

This contrast by age raises two interesting questions regarding the decline in mean tenure and long term employment and how this decline is related to the rate of “churning” in the

¹² Burdett (1978) presents a model of job search with this implication. Jovanovic (1979) presents model of matching in the labor market with the same implication.

labor market:

1. Are young workers taking longer to find good (long-lasting) matches or jobs? This would imply an increase in the new-job rate among younger workers.
2. Are older workers having more difficulty finding good matches when they lose jobs that may formerly have been “lifetime” jobs? This would imply an increase in the new-job rate among older workers.

In order to investigate how the new-job rate has changed over time, I estimate age-specific year effects using the same approach I used for means and for the probability of long-term employment. I estimate linear probability models of the probability of being in a new job using the same specification of explanatory variables (year, age, education, race, Hispanic ethnicity) in equation 3.3.

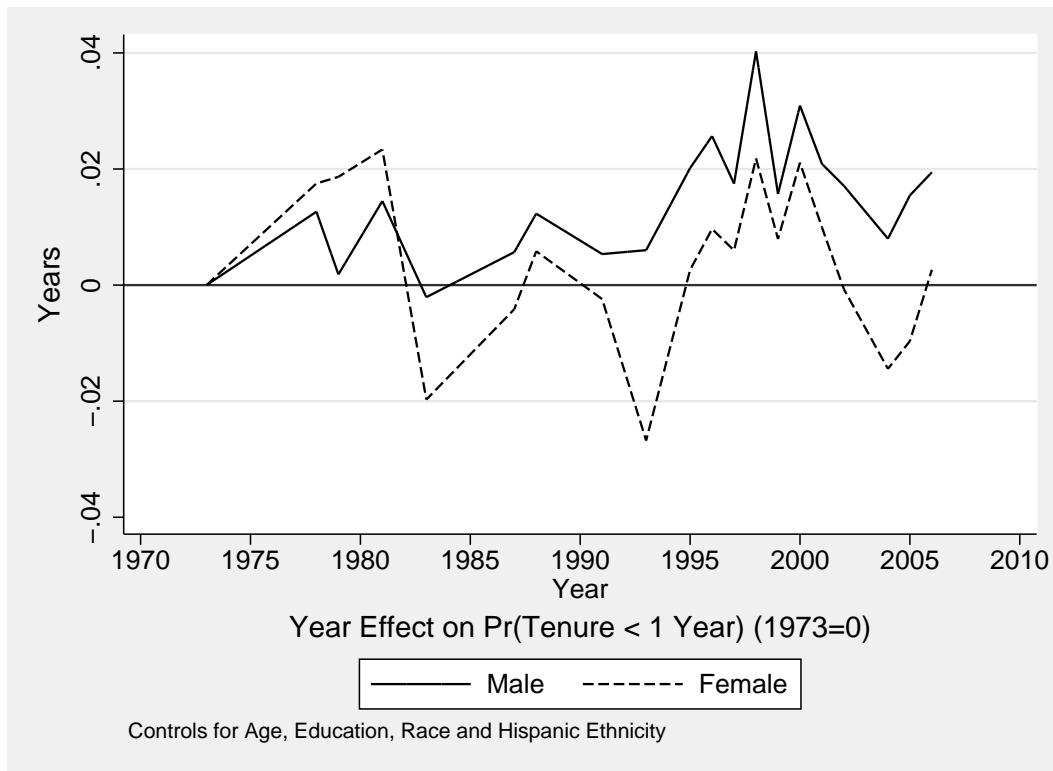


Figure 7: Year Effects on $Pr(T < 1)$

Figure 7 contains separate plots for males and females of the difference by year in the new-job rate relative to the 1973. The age-specific probability that a male worker has been with his employer for less than one year increased by about 4 percentage points between the 1973 and 1998 before falling by 2 percentage points by 2006. The age-specific probability that a female worker has been with her employer for less than one year fluctuates, returning to the 1973 level in 2006. These patterns appear to have a strong cyclical component, with higher new-job rates in periods with stronger labor markets and lower new-job rates in periods with weaker labor markets.

An implicit constraint in my model is that year effects are constant across age groups. Given the role that job change plays in matching and job search early in careers, I estimate separate year effects for different age groups. The top panel of figure 8 contains differences by year in the new-job rate relative to the 1973 estimated using a sample of workers aged 20-29. These estimates, which are very similar for males and females, show the pro-cyclical pattern noted above.

The bottom panel of figure 8 contains differences by year in the new-job rate relative to 1973 estimated using a sample of workers aged 30-39. These estimated year effects differ substantially from those for workers in their twenties. There is an increase of about 4 percent in the new job rate for males, and decrease of about 4 percent for females between the 1973 and 2006. The pattern for males is consistent with the hypothesis that men are job shopping in their twenties but have become less likely to settle into longer-term jobs in their thirties. The pattern for females likely reflects an increase in commitment to the labor force by women as they enter their thirties.

Given that older workers are less likely to be in long-term jobs, I next investigate how the new-job rate has changed for workers aged 40 and older. The top panel of figure 9 contains differences by year in the new-job rate relative to 1973 estimated using a sample of workers aged 40-49. The bottom panel of this figure contains differences by year in the new-job rate relative to 1973 using a sample of workers aged estimated using a sample of workers aged 50-64. Both plots show an increase in the probability of being on a new job for males. The magnitude of the increase (about 2 percentage points) is substantial when compared to the overall mean new job rates for older men in table 2. The new job rate for women in their

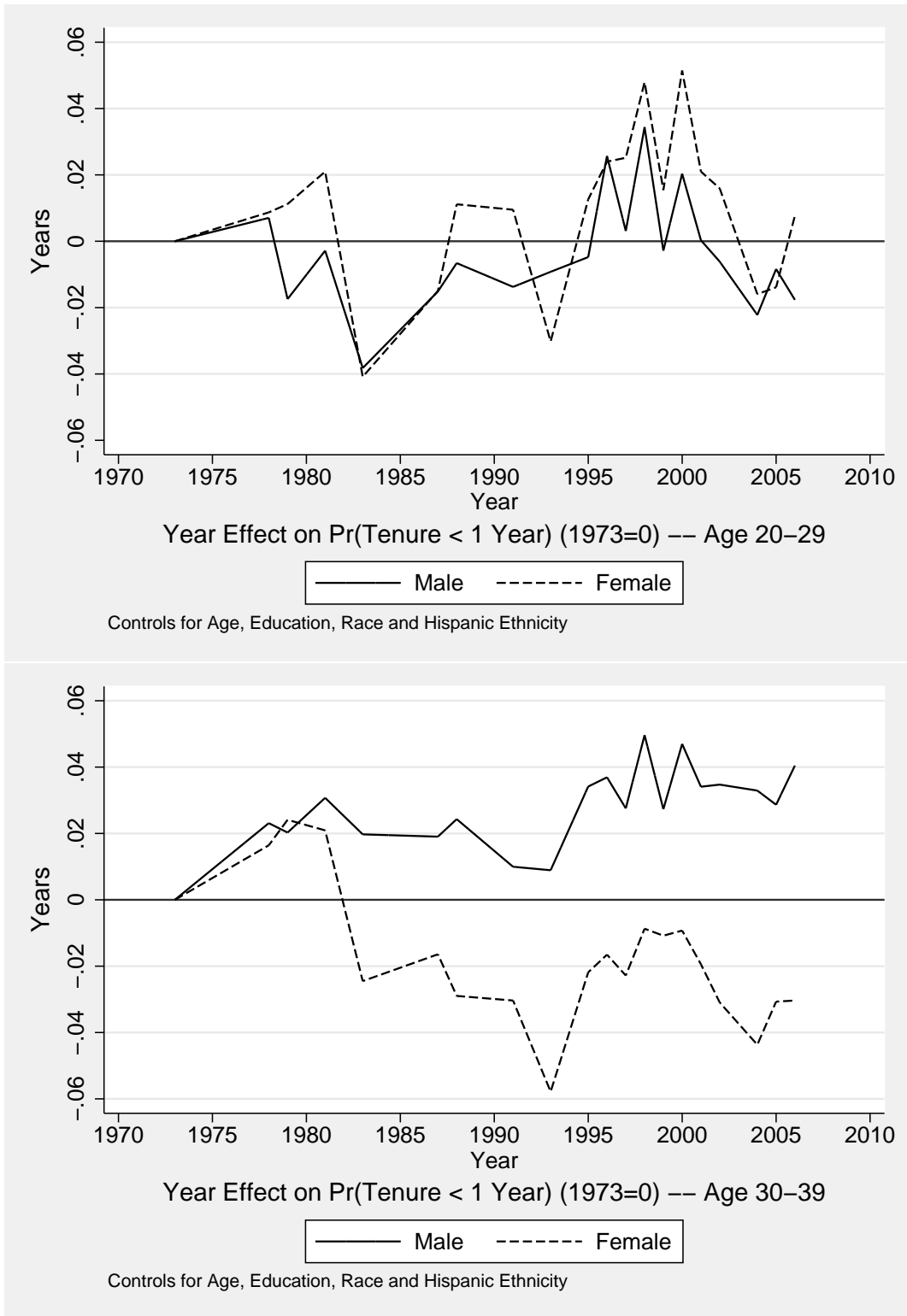


Figure 8: Year Effects on $Pr(T < 1)$ (Age 20-29 and 30-39)

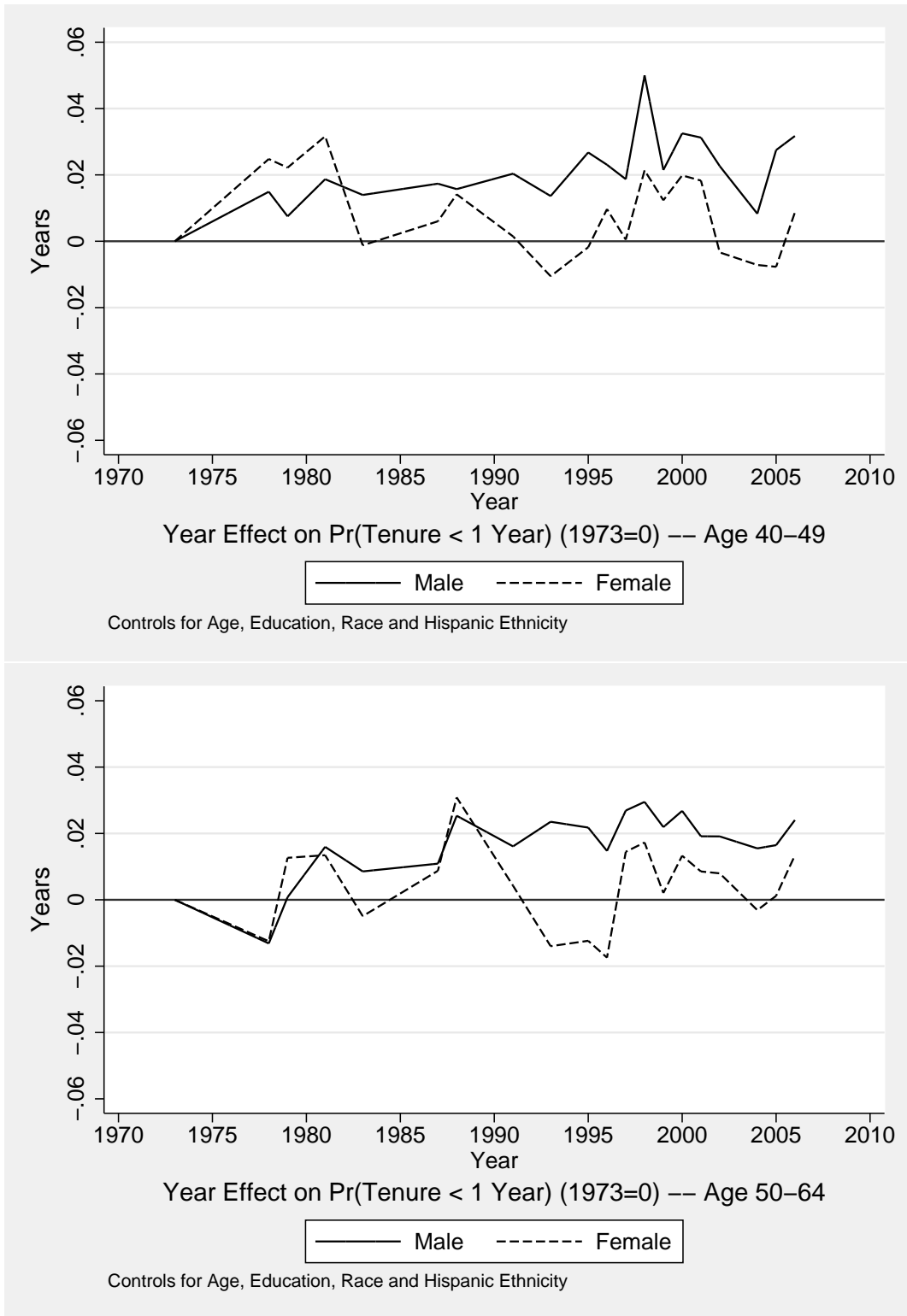


Figure 9: Year Effects on $Pr(T < 1)$ (Age 40-49 and 50-64)

forties or older shows no consistent pattern.

The overall time-series pattern of the age-specific new-job rate is a general increase over time for men aged 30 and older. Part of this reflects an extension of the period of “job-shopping” early in careers and part reflects increased probabilities of jobs ending later in careers. There is not much change over time in the age-specific new-job rate for women aside from a substantial decline for women in their thirties, likely reflecting a reduced likelihood of withdrawing from and subsequently re-entering the labor force at that age.

4 The Rate of Job Loss

Perhaps the most comprehensive source of information on the incidence and costs of job loss in the United States is the Displaced Workers Survey (DWS), administered every two years since 1984 as a supplement to the Current Population Survey (CPS). In this section, I analyze data on 913,083 individuals between the ages of twenty and sixty-four from the DWSs conducted as part of the January CPSs in 1984, 1986, 1988, 1990, 1992, 2002, 2004, and 2006 and the February CPSs in 1994, 1996, 1998, and 2000.

There are three important issues of measurement and interpretation that arise when comparing job loss rates calculated using the DWS over time.

1. The DWS asks only about a single involuntary job loss. The survey does not capture multiple job losses by the same worker. Neither does it capture worker terminations “for cause.” The survey is meant to capture worker terminations as the result of business decisions of the employer unrelated to the performance of the particular employee (e.g., a plant closing, a layoff, the abolition of a job). Thus, the measure of the job loss rate that I calculate is the fraction of workers who lost at least one job not “for cause” in the relevant period rather than the rate of destruction of worker-employer matches.
2. The DWS from 1984-1992 asked about job separations in the previous five years while the later DWS asked about job separation in the previous three years. The measure of job loss that I use is adjusted to account for this change in the recall period so that all rates are reported on a three-year basis. This adjustment is detailed in Farber (1997).

3. The basic wording of key questions changed since the inception of the DWS in 1984. This may have affected whether survey respondents would report a job separation in a particular circumstance as an involuntary separation in one survey but would not report a separation in the same circumstance as involuntary in another year. In Farber (1998) and Farber (2004), I use additional data from debriefing questions asked of a fraction of DWS respondents in 1996, 1998, and 2000 to investigate how changes in the wording of the key question may have affected the likelihood that a worker reported a particular separation as an involuntary job change. I use the results of that analysis to calculate re-weighted job loss rates that I present in this study.¹³

In these surveys, I count as job losers workers who reported a job loss in the three calendar years prior to the survey. Based on these data, I calculate the rate of job loss as the ratio of the number of reported job losers divided by the number of workers who were either employed at the survey date or reported a job loss but were not employed at the survey date. I then adjust these job loss rates as described in Farber (2004) to account for the change in the recall period from five years to three years in 1994 and changes in the wording of the key job loss question.

Figure 10 contains plots of adjusted three-year job loss rates computed from each of the ten DWSs from 1984-2004 along with the average civilian unemployment rate for each three-year period.¹⁴ The cyclical behavior of job loss is apparent, with job-loss rates clearly positively correlated with the unemployment rate ($\rho = 0.44$). Both unemployment and job-loss rates were high in the 1981-83 period, and they both fell sharply during the expansion of the mid-1980's. However, the job-loss rate rose much more sharply from the 1987-89 to the 1989-91 period than did the unemployment rate. The job-loss rate rose by fully 3.1 percentage points (from 7.1 percent to 10.2 percent) while the average unemployment rate rose by only 0.2 percent (from 5.7 percent to 5.9 percent) over this period. Between 1993 and 1999, both the job-loss and unemployment rates fell sharply, but the gap between them

¹³ Job losers are asked to report the reason for their job loss. One allowable response is "other." The adjustment for changes in the wording of the key job loss question discounts job loss rates for "other" reasons by 37.4% for the 1984-1992 DWS and by 74.8% for the 1994 and later DWSs. See Farber (1998) for details.

¹⁴ Information on rates of job loss is presented most accessibly in graphical form, and the discussion here is organized around a series of figures. All counts are weighted using the CPS sampling weights.

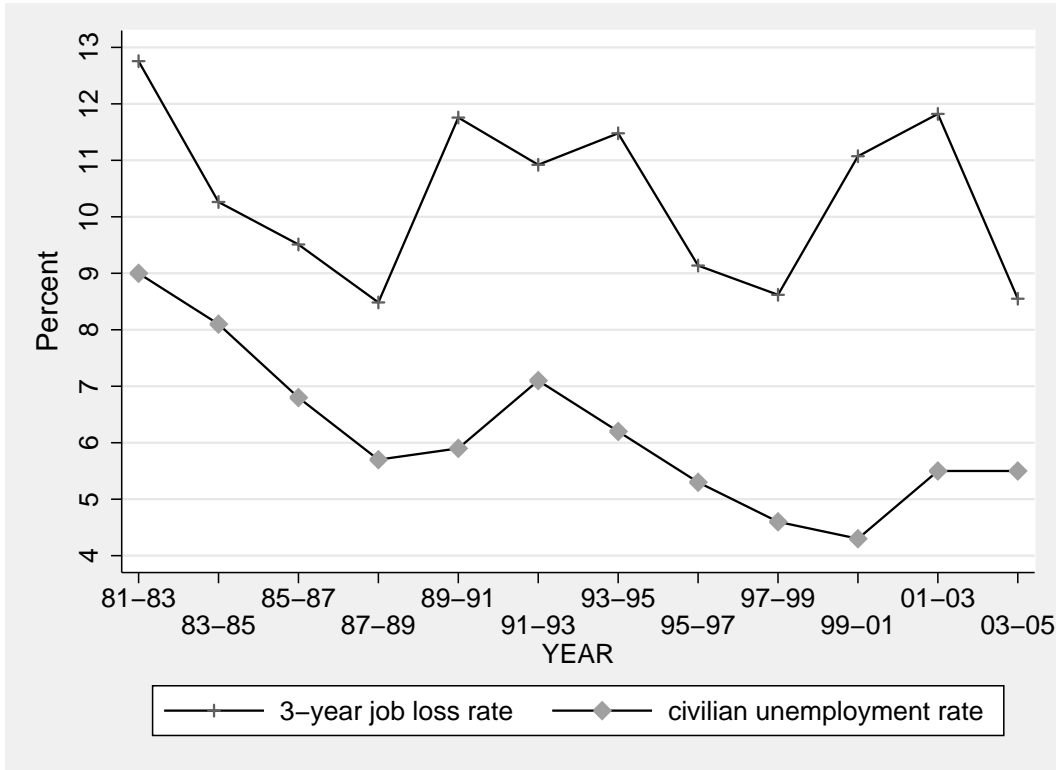


Figure 10: Unemployment and Job Loss Rates, by Year

remained larger than in the strong labor market of the late 1980s.

The unemployment rate continued to fall in the 1999-2001 period before rising somewhat in the 2001-03 period. The job loss rate rose sharply after the 1997-99 trough through the 2001-03 period before falling off in the 2003-05 period.¹⁵

Figure 11 contains three-year rates of job loss by year for each of four education categories. Not surprisingly, job loss rates are dramatically higher for less educated workers than for more educated workers, although this gap has narrowed over time. There is a strong cyclical pattern in job loss rates for all groups of worker. It does appear that there was an upsurge in job loss rates for more educated workers in the early and mid-1990s and again early in the new century. The job loss rate for workers with at least 16 years of education attained

¹⁵ The use of three-year averages here hides the facts that the job loss rate was steady in 1999 and 2000 before increasing sharply in 2001 while the unemployment rate declined slightly in 1999 and 2000 before increasing slightly in 2001.

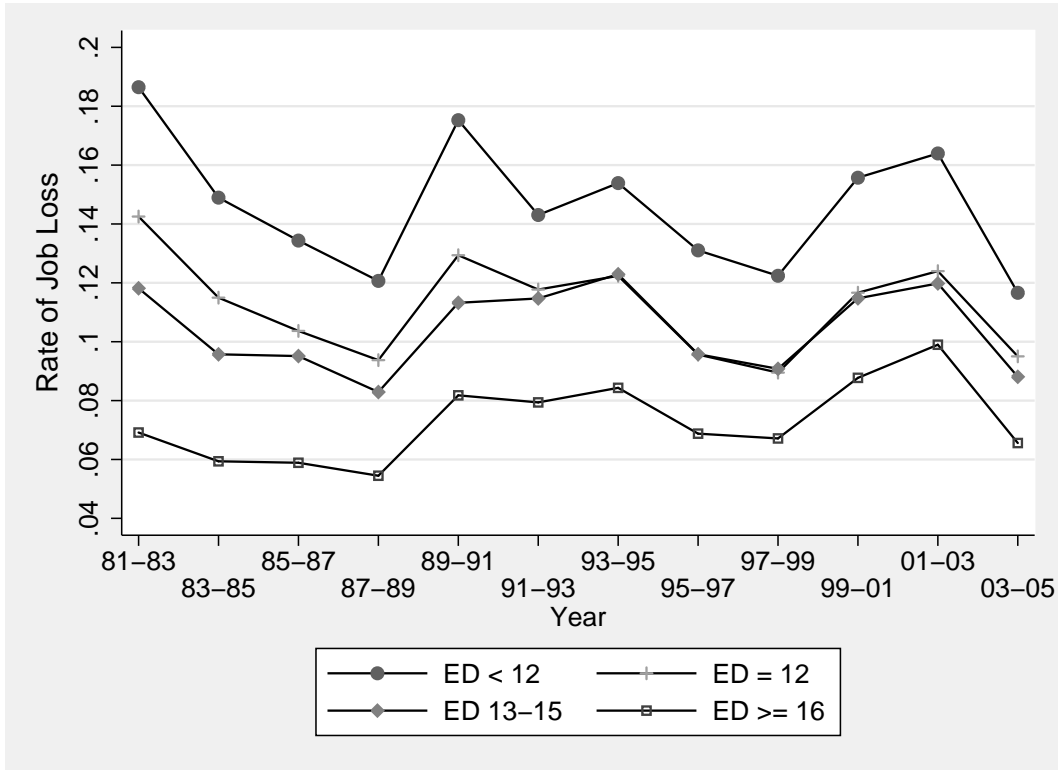


Figure 11: Three-Year Job Loss Rate by Education, 1981-2003.

its highest level in the 2001-2003 period, reaching 9.9 percent. Job loss rates for the other educational groups show a cyclical pattern but no upward trend.

Figure 12 contains three-year job loss rates by year for four age groups covering the range from 20-64. Job loss rates are highest for the youngest workers (20-29) and generally show the standard cyclical pattern. The job loss rates of the oldest two group, ages 40-49 and 50-64, are very similar. There is little evidence here of a secular increase in job loss rates among older workers.

5 Reconciling the Trends in Job Tenure and Job Loss

The sharp decline in job tenure combined with the lack of any trend in rates of job loss presents a puzzle. In this section, I explore the possibility that this apparent contradiction is due to a relative increase in job-loss rates among high-tenure workers.

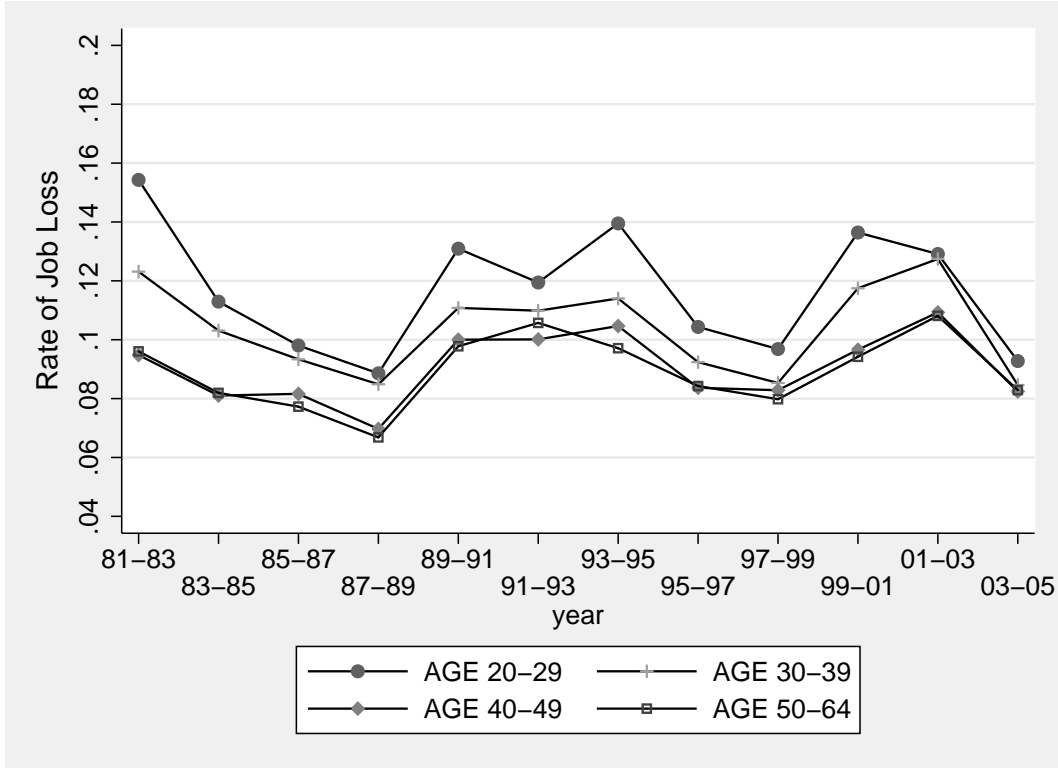


Figure 12: Three-Year Job Loss Rate by Age, 1981-2003.

Unfortunately, the available data do not allow direct calculation job-loss rates by tenure. Tenure is not observed in the DWS for workers who do not report that they have been displaced. I use Bayes' rule to recover the probability of job loss by tenure category using the tenure data from the CPS as well as the DWS. The appropriate relationship defining the probability that worker i in tenure category j is displaced in period t is

$$Pr(D_{it} = 1|T_{it} = j) = \frac{Pr(T_{it} = j|D_{it} = 1)Pr(D_{it} = 1)}{Pr(T_{it} = j)} \quad (5.1)$$

where $D_{it} = 1$ if worker i in period t suffers a job loss and $T_{it} = j$ if worker i is in tenure category j in period t . The only piece of equation 5.1 that is not available directly from the DWS is $Pr(T_{it} = j)$, the unconditional tenure distribution. However, data from the CPS on job tenure are available for all workers in at least one of the two years immediately prior to each DWS other than the 1986 DWS. I use this information to compute unconditional

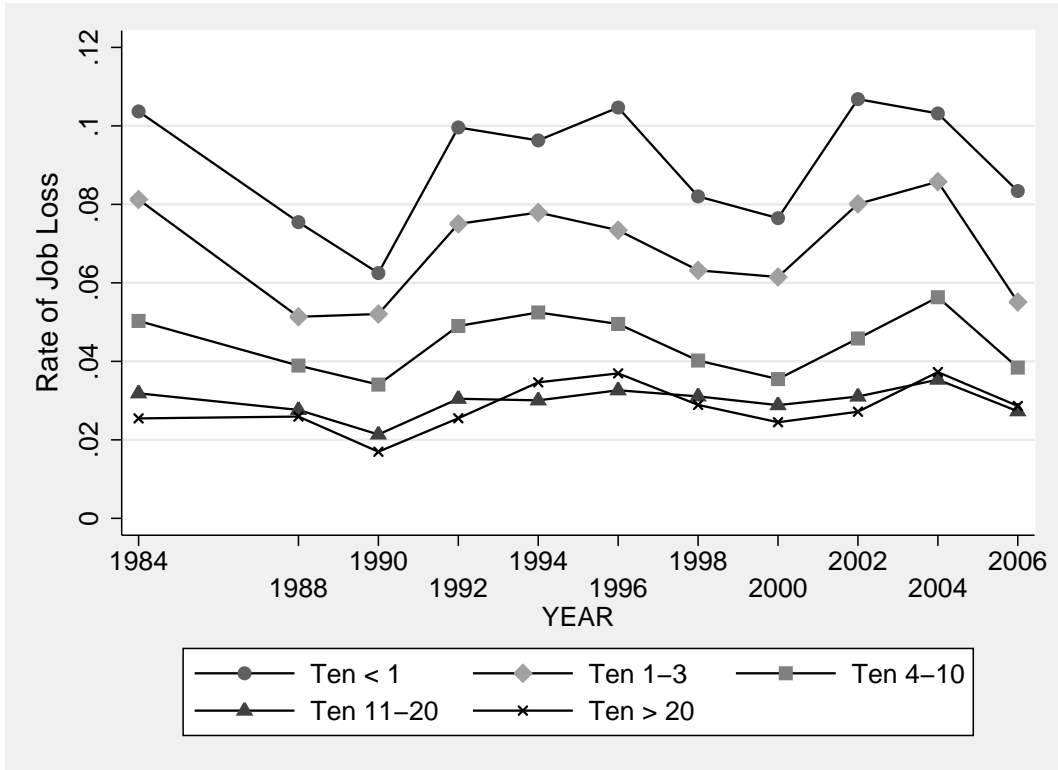


Figure 13: Three-Year Job Loss Rate by Tenure, 1984-2006.

tenure distributions appropriate to each DWS (other than 1986).¹⁶

Tenure-specific job-loss rates calculated based on equation 5.1 are plotted in figure 13. These show the unsurprising result that the probability of job loss is strongly monotonically declining in tenure. The likelihood of job loss (averaged across years) falls from 9.0 percent for workers in their first year on the job to 2.8 percent for workers with more than twenty years on the job. The results are more remarkable for what is not found. There seems to be no secular trend in the tenure-specific probabilities of job loss.

Covariation between sex, age, education, and tenure in determining the rate of job loss suggests that a multivariate analysis of job loss would be useful. Without longitudinal data

¹⁶ Since 1994, the DWS collects information on tenure on the lost job only for the subset of displaced workers who report a reason for job loss of 1) slack work, 2) position or shift abolished, or 3) plant closing. This information is not collected for displaced workers who report a reason for job loss of 1) temporary job ended, 2) seasonal job ended, or 3) other. For this reason, the succeeding analysis measures rates of job loss only for the “big three” reasons. See Farber (2005) for more details.

on all variables, a micro-level analysis is not feasible, but I can compute cell-mean level rates of job loss using the Bayes' Rule approach. I compute job-loss rates for groups defined by education category, tenure category, age category, sex, and time period. There are 1936 cells, defined by these five variables, for which I have computed job loss rates using information on individual characteristics and job displacement from the CPS mobility data and the DWS. Unfortunately, cell sizes are too small to expand the breakdown to include other variables.

The job loss rates for each category are computed using Bayes' rule applied to the conditional probabilities associated with job loss, tenure, sex, and education. It is straightforward to show that

$$Pr(D_{it} = 1|T_{it} = j, X_{ikrmt} = 1) = \frac{Pr(T_{it} = j|D_{it} = 1, X_{ikrmt} = 1)Pr(D_{it} = 1, X_{ikrmt} = 1)}{Pr(T_{it} = j, X_{ikrmt} = 1)} \quad (5.2)$$

where $X_{ikrmt} = 1$ if individual i in year t is of sex k in age category r with education level m . This particular representation of the conditional probability of job loss is used because it allows computation of the probability of displacement conditional on tenure and the other controls without direct information on the joint distribution of displacement and tenure.

Table 3 contains estimates of weighted least squares (WLS) regressions of the log of the probability of job loss in a cell on the characteristics defining the cell. The weights used for each cell are estimates of the variance of the log probability of job loss for that cell.¹⁷

Specification 1 in table 3 contains estimates of a model pooling all years of the log probability of job loss with main effects for sex, age category, education level, tenure category, and DWS survey year. These results demonstrate several strong patterns:

- There is a strong negative relationship between the probability of job loss and tenure.

Workers with more than twenty years of tenure have a probability of job loss that is

¹⁷ The log of the probability of job loss in equation 5.2 is a linear combination of the logs of its component probabilities. Specifically,

$$\begin{aligned} \ln(Pr(D_{it} = 1|T_{it} = j, X_{ikrmt} = 1)) &= \ln(Pr(T_{it} = j|D_{it} = 1, X_{ikrmt} = 1)) \\ &+ \ln(Pr(D_{it} = 1, X_{ikrmt} = 1)) - \ln(Pr(T_{it} = j, X_{ikrmt} = 1)). \end{aligned}$$

The variance of each component probability is $p(1-p)/n$, where p is the relevant probability and n is the sample size that the estimate of the probability is based on. The variance of the log of each probability is computed using the delta method as $(1-p)/pn$, and the variance of $\ln(Pr(D_{it} = 1|T_{it} = j, X_{ikrmt} = 1))$ is computed as the sum of the variances of the log of the three component probabilities.

Table 3: WLS Estimates of Log-Probability of Job Loss by Sex, Education, Tenure, Age and Year Cells, 1984-2006

Variable	(1) 1984-06	(2) 1984-90	(3) 1992-00	(4) 2002-06
Ten 1-3	-0.270 (0.018)	-0.255 (0.033)	-0.273 (0.027)	-0.294 (0.032)
Ten 4-10	-0.744 (0.019)	-0.700 (0.035)	-0.741 (0.028)	-0.811 (0.033)
Ten 11-20	-1.176 (0.028)	-1.123 (0.052)	-1.180 (0.041)	-1.244 (0.049)
Ten > 20	-1.274 (0.039)	-1.298 (0.078)	-1.237 (0.056)	-1.325 (0.064)
Female	-0.446 (0.014)	-0.590 (0.026)	-0.414 (0.020)	-0.361 (0.024)
Ed < 12	-0.029 (0.022)	-0.006 (0.035)	-0.042 (0.033)	-0.042 (0.043)
Ed 13-15	-0.071 (0.017)	-0.182 (0.033)	-0.037 (0.025)	-0.020 (0.030)
Ed \geq 16	-0.293 (0.019)	-0.556 (0.038)	-0.253 (0.028)	-0.172 (0.032)
Age 25-34	0.236 (0.022)	0.195 (0.037)	0.226 (0.033)	0.320 (0.041)
Age 35-44	0.264 (0.023)	0.173 (0.042)	0.291 (0.034)	0.345 (0.043)
Age 45-54	0.277 (0.026)	0.075 (0.049)	0.300 (0.038)	0.416 (0.045)
Age 55-64	0.045 (0.031)	-0.156 (0.058)	0.065 (0.046)	0.210 (0.052)
Year FEs	Yes	Yes	Yes	Yes
N	1936	528	880	528
R-squared	0.732	0.794	0.715	0.761

Note: Based on calculated conditional cell probabilities from equation 5.2 using data from the 1984-2006 DWS and tenure data from various CPSs from 1983-2005. The estimates are weighted by the inverse sampling variance of the log probability of job loss in each cell. See Note 17 for details.

28 percent of the rate of job loss of workers in their first year on the job.

- Females have lower rates of job loss than men (about 15 percent lower).
- Rates of job loss are inversely related to education. The rate of job loss for college graduates is 75 percent of the rate of job loss of otherwise-equivalent high school graduates (the base group).
- Controlling for tenure, there is a non-monotonic relationship between age and the rate of job loss. Workers ages 25-54 have a rate of job loss that is about 25 percent higher than either workers aged 20-24 or workers aged 50-64.

Of course, the key question is whether these relationships have changed over time. In order to investigate this, I reestimated the model separately for three sub-periods: 1984-1990, 1992-2000, and 2000-2006.¹⁸ These results are interesting both for the changes revealed in some dimensions and the lack of change in others.

- Conditional on age, there is no systematic change over time in the relationship of the displacement rate with job tenure. This suggests that we cannot reconcile the decline in age-specific job tenure in the U.S. with the lack of increase in job loss with changes in the tenure distribution of job loss.
- The female-male differential in job loss has narrowed, falling from 45 percent lower for females in the 1980s to 30 percent lower in the 2000s.
- The strong advantage of education in reducing job loss rates has diminished substantially. The college/high-school differential in job loss rates fell from 43 percent lower in the 1980s to 16 percent lower in the 2000s.
- Conditional on tenure, older workers have become substantially more susceptible to job loss. Workers aged 45-54 had roughly equivalent job-loss rates to workers in 20-24 in the 1980s. By the 2000s, these workers job-loss rates were 52 percent higher than those of workers 20-24.

¹⁸ These years refer to the DWS survey years. The 1984-1990 sample covers job loss from 1981-1989, the 1992-2000 sample covers job loss from 1989-1999, and the 2000-2006 sample covers job loss from 1999-2005.

Table 4: WLS Estimates of Log-Probability of Job Loss by Sex, Education, Tenure, and Year Cells, 1984-2006

Variable	(1) 1984-06	(2) 1984-90	(3) 1992-00	(4) 2002-06
Ten 1-3	-0.257 (0.037)	-0.244 (0.071)	-0.254 (0.051)	-0.273 (0.061)
Ten 4-10	-0.690 (0.037)	-0.654 (0.071)	-0.689 (0.051)	-0.727 (0.061)
Ten 11-20	-1.121 (0.037)	-1.135 (0.072)	-1.100 (0.051)	-1.141 (0.062)
Ten > 20	-1.213 (0.038)	-1.373 (0.074)	-1.161 (0.052)	-1.171 (0.062)
Female	-0.411 (0.024)	-0.588 (0.048)	-0.379 (0.033)	-0.321 (0.040)
ED < 12	-0.053 (0.038)	0.005 (0.062)	-0.100 (0.053)	-0.116 (0.071)
ED 13-15	-0.052 (0.030)	-0.176 (0.061)	-0.023 (0.040)	0.007 (0.048)
ED \geq 16	-0.383 (0.033)	-0.767 (0.073)	-0.321 (0.045)	-0.244 (0.052)
Age	No	No	No	No
Year FEs	Yes	Yes	Yes	Yes
N	438	118	200	120
R-squared	0.835	0.883	0.842	0.864

Note: Based on calculated conditional cell probabilities from equation 5.2 using data from the 1984-2006 DWS and tenure data from various CPSs from 1983-2005. The estimates are weighted by the inverse sampling variance of the log probability of job loss in each cell. See Note 17 for details.

The strong positive relationship between age and tenure may be masking a change in relationship between the probability of job loss and tenure. To this end, I re-estimate the model presented in table 3 without controlling for age.¹⁹ The results of this analysis are presented in table 4, and they are quite similar to those controlling for age. There is no evidence that the strong negative relationship between tenure and rates of job loss has weakened.

Thus, the puzzle of declining job tenure over time with no increase in measured rates

¹⁹ In other words, I estimate the model over cells that do not condition on age.

of job loss in the DWS remains. One real possibility is that the DWS does not measure rates of job loss very well, perhaps because of ambiguity among workers who leave jobs in why they, in fact, left. For example, workers who accept buy-outs to leave their jobs may not understand and report this as a job loss but as a voluntary job change. To the extent that this is an important phenomenon, rates of job change, whether classified as voluntary or involuntary, have increased. A more comprehensive survey of job changes and the underlying circumstances is needed in order to understand the decline in long-term employment more fully.

Appendix I: The CPS Data on Employer Tenure

At irregular intervals, the Census Bureau has appended mobility supplements to the January or February Current Population Surveys. The years in which they did so include 1951, 1963, 1966, 1968, 1973, 1978, 1981, 1983, 1987, 1991, and in even years from 1996-2006. These supplements contain information on how long workers have been continuously employed by their current employer, and they are asked of all eight CPS rotation groups. However, only the supplements since 1973 are available in machine-readable form. Information on job durations is also available in pension and benefit supplements to the CPS in May of 1979, 1981, 1983, and 1988, and in April 1993. These supplements contain information on how long workers have been working for their current employer, and they are asked of four of the eight CPS rotation groups. Finally, information on job durations is available in the continuous and alternative employment arrangement supplements (CAEAS) to the CPS in February of 1995, 1997, 1999, 2001, and 2005. In total there are 21 CPS supplements with information on employer tenure available in machine readable form over the period from 1973 to 2006, and my analysis relies on these data.

A question of comparability of the data over time that must be kept in mind when interpreting the results arises because of a significant change in the wording of the central question about job duration. The early mobility supplements (1951-1981) asked workers what year they started working for their current employer. In later mobility supplements (1983-2006), in all of the pension and benefit supplements (1979-1993), and in all of the CAEAS supplements (1995-2005) workers were asked how many years they worked for their current employer. If the respondents were perfectly literal and accurate in their responses (a strong and unreasonable assumption), then these two questions would yield identical information (up to the error due to the fact that calendar years may not be perfectly aligned with the count of years since the worker started with his/her current employer). But responses are not completely accurate, and this is best illustrated by the heaping of responses at round numbers. The empirical distribution function has spikes at five-year intervals, and there are even larger spikes at ten-year intervals. In the early question, the spikes occur at round calendar years (1960, 1965, etc.). Later, the spikes occur at round counts of years (5, 10, 15, etc.).

There are also subtle but potentially important changes in wording of the key questions even within these surveys. All of the mobility supplements since 1983 ask individuals how long they have worked *continuously* (italics added) for their current employer. However, neither the pension and benefit supplements nor the CAEAS include the word “continuously”. The May 1979 and 1983 pension and benefit supplements ask individuals how long they have worked for their current employer and specify that if there was an interruption greater than one year to count only the time since the interruption. The May 1988 and April 1993 supplements and the CAEAS ask individuals how long they have worked for their current employer without any reference to interruptions or continuity. Thus, it might be the case that the mobility supplements would yield shorter tenures than the pension and benefit supplements and the CAEAS due to the requirement of continuity in the former. And it

might be the case that the early two pension and benefit supplements would yield shorter durations than the later two pension and benefit supplements due to the consideration of long interruptions given in the early supplements. I make no explicit allowance for these differences in my analysis, but they should be kept in mind when interpreting the results.

With the exception of jobs of less than one year, all of the supplements before the February 1996 mobility supplement collect data on job duration in integer form reporting the number of years employed. For jobs of less than one year, the mobility supplements report the number of months employed while the pension and benefit supplements report only the fact that the job was less than one year old. The February 1996 and later mobility supplement ask workers how long they have worked continuously for their current employer and accepts a numerical response where the worker specifies the time units. The 1995-2005 CAEAS ask workers how long they have worked for their current employer and accepts a numerical response where the worker specifies the time units. Virtually all workers in jobs even five years old and all workers in jobs 10 years old or longer, report job durations in years.

One reasonable interpretation of the integer report of the number of years is that workers round to the nearest integer when they report jobs of duration of at least one year.²⁰ For example, a response of 10 years would imply tenure greater than or equal to 9.5 years and less than 10.5 years. In order to create a smooth tenure variable, I assume that the distribution of job tenure is uniform in these one-year intervals. Given a reported tenure of T years, I replace T by $T - 0.5 + u$ where u is a random variable distributed uniformly on the unit interval.²¹

My sample consists of 876,063 not self employed workers aged 20-64 from the 21 CPS supplements covering the period from 1973 to 2006. The self-employed are not included because the concept of employer tenure is less clear for the self-employed, and, in any case, the CPS supplements do not contain consistent information on tenure for the self-employed.

²⁰ This ignores the heaping of the tenure distribution at multiples of five and ten years.

²¹ Where reported tenure is zero years, I assume that tenure is uniformly distributed between zero and one and define tenure as u . Given that jobs are more likely to end earlier in the first year than later in the first year, this is not completely accurate (Farber, 1994). However, the measures used in my analysis will not be affected by this representation. Where reported tenure is exactly one year, I assume that true tenure is uniformly distributed between 1 and 1.5 and define tenure as $1 + u/2$.

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