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5 Are Lifetime Jobs Disappearing? Job Duration in the United States, 1973–1993

Henry S. Farber

5.1 Introduction

The public perception is that there has been a fundamental deterioration of job security in the United States. It is not unusual to see reports in the media to this effect. Headlines such as "Jobs in an Age of Insecurity" are not uncommon. Neither are statements like "Thirty months into recovery, Americans are realizing that the Great American Job is gone" (*Time*, 22 November 1993, p. 32). The same article in *Time* reports survey results finding that "two-thirds believed that job security has deteriorated over the past two years, although those years have seen continuous economic growth." These stories may not only reflect but also help shape the generally reported view that job security is declining.

Job security is not a precisely defined concept and has several dimensions. One dimension is the subjective perception of how secure one's job is. This depends both on how likely it is that the worker will be terminated involuntarily from his or her job and on how valuable that job is to the worker. If the job can be replaced easily (at low pecuniary and nonpecuniary cost) with an equivalent job, then the worker may not feel terribly insecure regardless of the likelihood of losing the current job. On the other hand, if replacing the job is difficult, then even low probabilities of losing the job may engender feelings of insecurity. On this basis, one way to investigate changes in job security is to measure changes in the likelihood and costs of job loss. Some of my earlier work through 1993 from the Displaced Workers Supplements to the Current

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Population Survey provides evidence that the costs of job loss in terms of post-displacement employment probabilities and earnings are substantial but have not increased since the early 1980s, when the Displaced Workers Supplement was initiated (Farber 1993, 1997). My analysis of the same data shows a small increase in the likelihood of job loss, particularly for more educated workers. It is difficult to find strong evidence in these data of more job insecurity.

An alternative, complementary, and perhaps longer run view of job security is based on the idea that stable long-run employment relationships are an important component of job security for workers, and it is this concept of job security that shapes my analysis. I examine evidence on job durations in order to determine if, in fact, a systematic change in the likelihood of long-term employment occurred between 1973 and 1993.

There is relatively long-standing concern that the basic 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. There has been concern that employers are moving toward greater reliance on temporary workers, on subcontractors, and on part-time workers. Potential reasons 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 the belief that these changes result in lower quality (lower paying and less secure) jobs for the average worker.

The analysis in this paper is based on evidence regarding the duration of jobs in progress from supplements to the Current Population Survey with relevant information for selected years from 1973 to 1993. In order to measure changes in the distribution of job durations, I examine changes in selected quantiles (the median and the .9 quantile) of the distribution of duration of jobs in progress. I also examine selected points in the cumulative distribution function including the fraction of workers who have been with their employers (1) no more than 1 year, (2) more than 10 years, and (3) more than 20 years. These data and the distributional measures used are described in more detail in section 5.2.

The central findings, presented in sections 5.3 and 5.4, are clear. No systematic change has occurred in various measures of the overall distribution of job duration over the past two decades. However, the overall figures mask two important, though perhaps unsurprising, changes in the job durations of particular groups of workers. First, individuals, particularly men, with little education (less than 12 years) are less likely to be in jobs of long duration today than they were 20 years ago. This is consistent with the declining real earnings (both relative and absolute) of the least educated workers in the U.S. economy, and it may be part of the mechanism of this decline. Second, women with at least a high school education are substantially more likely to be in long-term jobs today than they were 20 years ago. This is likely a natural result of the declining frequency with which women withdraw from the labor market for periods of time. The increased job durations for women may also help explain the decline in the male-female wage gap in the 1980s (Wellington 1992).

5.2 Data and Measurement Issues

5.2.1 Current Population Survey Data on Job Duration

At irregular intervals, the Census Bureau has appended mobility supplements to the January Current Population Survey (CPS). The years in which it did so include 1951, 1963, 1966, 1968, 1973, 1978, 1981, 1983, 1987, and 1991.[1] These supplements contain information on how long workers have been continuously employed by their current employers. However, only the supplements since 1973 are available in machine-readable form.[2] Information on job duration is also available in pension and benefit supplements to the CPS in May 1979, 1983, and 1988 and in April 1993.

Others have used these data to analyze job duration. An important early paper is by Hall (1982), who used published tabulations from some of the January mobility supplements to compute contemporaneous retention rates. Hall found that, while any particular new job is unlikely to last a long time, a job that has already lasted 5 years has a substantial probability of lasting 20 years. He also finds that a substantial fraction of workers will be on a "lifetime" job (defined as lasting at least 20 years) at some point in their lives. Ureta (1992) used the January 1978, 1981, and 1983 mobility supplements to recompute retention rates using artificial cohorts rather than contemporaneous retention rates.

Two recent papers have examined changes in employment stability using data from the mobility and pension supplements to the CPS. Swinnerton and Wial (1995), using data from 1979–91, analyze job retention rates computed from artificial cohorts and conclude that there was a secular decline in job stability in the 1980s. In contrast, Diebold, Neumark, and Polsky (1994), using data from 1973–91 to compute retention rates for artificial cohorts, find that aggregate retention rates were fairly stable over the 1980s but retention rates declined for high school dropouts and for high school graduates relative to college graduates over this period.

In my analysis, I use data from the mobility supplements to the January 1973, 1978, 1981, 1983, 1987, and 1991 CPS and from the pension and benefit supplements to the May 1979 and April 1993 CPS.[3] These surveys cover 8 years over the 20-year period from 1973 to 1993. One feature that will distin-

1. There was also a mobility supplement to the February 1996 CPS, but it was not available at the time this analysis was performed.

2. Only summary tables are available for the 1951, 1963, 1966, and 1968 surveys.

3. There are two pension and benefit supplements that I did not use for different reasons. I did not use the May 1983 supplement because I already have data for 1983 in the January mobility supplement. I did not use the May 1988 supplement because it did not have data on duration for self-employed workers.

guish my analysis is that it uses more recent data (April 1993) than even the newest of the earlier work.

A question of comparability of the data over time arises because of substantial changes in the wording of the central question about job duration. The early January supplements (1951–81) asked workers what year they started working for their current employers (the early question). In later January supplements (1983–91) and in all of the pension and benefit supplements (1979–93), workers were asked how many years they had worked for their current employers (the later question). If the respondents were perfectly literal and accurate in their responses (a strong and unreasonable assumption), 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 or 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 5-year intervals, and there are even larger spikes at 10-year intervals.[4] 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.). The two questions may also evoke systematically different responses. Although I do not deal with the comparability problem directly, a preliminary comparison of quantiles of the 1979 distribution of job durations (based on the new question) with quantiles of the 1978 and 1981 distributions of job durations (based on the old question) does not show any systematic difference.

With the exception of jobs of less than one year, the data on job duration are collected in integer form (what year started or how many years employed). This raises questions of interpretation that are particularly serious in examining movements in quantiles. Interpreting the integer responses requires some arbitrary decisions. First consider the early question, which asked what year the worker started working for the current employer. For a survey conducted in January of year T_S, a response of year T_0 to the question of when the job was started was interpreted as a job duration of $D = \max(T_S - T_0, 1)$. Thus a duration of D years computed this way represents a "true" duration (D_T) that is (approximately) in the interval $D - 1 < D_T \leq D$. If there were a uniform distribution of job durations within intervals, D would overstate D_T by one-half year on average. Now consider the later question, which asked how many years the worker has been with the current employer. Call this response Y. If a worker has been with the employer less than one year, he or she is asked the number of months with the employer. I ignore the information on months for these workers and interpret the job duration as $D = \min(Y, 1)$. Thus all workers with durations less than or equal to one year are coded as having durations of one year. The interpretation of workers with reported durations of one year or

4. Ureta (1992) accounts for these spikes explicitly in her estimation procedure. Swinnerton and Wial (1995) work around these spikes in selecting intervals over which to compute retention rates.

longer depends on the rounding rules used by the respondents. One reasonable rule would be rounding to the nearest integer so that a response of Y would represent durations in the range from $Y - .5$ to $Y + .5$. Another reasonable rule would be for the respondent to perform the calculation of current year minus starting year and report the difference. This rule seems more reasonable for longer term jobs, and it yields a result equivalent to the procedure I use for the early question. The result is again to overstate job duration by one-half year on average.

There is no way to get direct evidence about how respondents interpret the later-style duration question. However, as noted above, a comparison of the distribution of responses to the 1979 question (later style) with the distributions of responses to the 1978 and 1981 questions (early style) does not show any systematic bias.[5] I proceed assuming that respondents answer the later question as if they report the difference in calendar years between the current date and the job start date. Thus a measured duration of D is interpreted throughout as representing a true duration in the interval $D - 1 < D_T \leq D$.

5.2.2 Interpolated Quantiles

Because job duration data are available in integer form with substantial fractions of the data at particular values, it is difficult to examine movements in quantiles. For example, the median job duration for a specific group of workers might be five years, and it might be the case that 10 percent of the sample reports job durations of five years. Ten years later, the distribution of job durations might have shifted to the right fairly substantially, but the median job duration might still be five years. The problem is that the cumulative distribution function for the integer data is a step function, and the movement "along" a step will not change the quantile unless the next step is reached.

As a result, I use *interpolated quantiles,* defined as

$$\theta_\tau = (1 - \lambda)D_k + \lambda D_{k+1}, \tag{1}$$

where θ_τ is the τth interpolated quantile of the distribution of job durations, D_k is the largest job duration such that $\Pr(D \leq D_k) < \tau$, and D_{k+1} is the smallest job duration such that $\Pr(D \leq D_{k+1}) > \tau$. In this case, the true τth quantile is D_{k+1}, and the τth interpolated quantile is simply a weighted average of the τth quantile and the next smaller observed value of job duration. The weight, λ, is,

$$\lambda = (\tau - P_k)/(P_{k+1} - P_k), \tag{2}$$

where $P_k = \Pr(D < D_k)$ and $P_{k+1} = \Pr(D < D_{k+1})$. In effect, this calculation assumes that job durations are uniformly distributed within each interval. It is straightforward to use the delta method to compute sampling variances for

5. The lack of systematic bias can be examined in the tables and figures presented below. Of course, this evidence is indirect, and it is possible that there is bias but a temporary increase in the 1979 job durations is masking the bias.

these interpolated quantiles under the assumption that the value of the interpolated quantile does not move to a different interval. All quantile results shown below are interpolated quantiles as I define them here. I refer to them simply as quantiles.

5.2.3 Fractions of Workers in Short-Term and Long-Term Jobs

I also examine the fractions of workers who fall into different intervals in the job duration distribution. These are effectively selected points on the cumulative distribution function of job duration and the inverse function of the quantiles. I examine variation in the fractions of workers who report having been with their employers (1) no more than 1 year, (2) more than 10 years, and (3) more than 20 years. These points on the distribution give a clear picture of what has happened to the incidence of very short term jobs and long-term or near lifetime jobs. It is straightforward (indeed more straightforward than computation of the interpolated quantiles) to compute these fractions using the same interpretations of the job duration information that I discussed above.

5.2.4 Employment-Based and Population-Based Distributions of Job Durations

Cyclical changes in the composition of the sample raise another important measurement issue. It is clear that workers with little seniority are more likely to lose their jobs in downturns (Abraham and Medoff 1984). Thus we would expect quantiles of the distribution of job durations to be countercyclical; tight labor markets will lead the distribution of job durations to lie to the left of the distribution in slack labor markets. Since secular rather than cyclical changes are of interest here, an alternative measure of the distribution that is relatively free of cyclical movements would be useful.

In the standard analysis, we use employed individuals in a given category (e.g., workers in a particular age range) as the base group when computing distributional measures. I call quantiles computed this way *employment-based quantiles,* and I call probabilities of having job duration in a particular category (up to 1 year, more than 10 years, and more than 20 years) *employment-based probabilities.* Cyclical fluctuations in employment add or subtract individuals from the base group for the employment-based measures. A reasonable alternative would be to use the entire population in a given category (e.g., individuals in a given age range) regardless of employment status to compute the measures assuming that those not employed have zero job duration. I call these *population-based measures.*

The employment-based and population-based measures clearly measure different distributions, but both have straightforward interpretations. For example, the median computed on an employment basis is the median duration of jobs in existence at a point in time. In contrast, the median computed on a population basis is the median length of time an individual has been employed (counting as zero the duration of those not employed). As such, the population-based median will be zero if less than half of the relevant group is not working. The contrast between the employment-based and the population-based probabilities is interpreted similarly. For example, the employment-based probability of being on a job more than 10 years is the fraction of workers who have been on their jobs more than 10 years. In contrast, the population-based probability of being on a job more than 10 years is the fraction of all individuals (employed or not) who have been on their jobs more than 10 years.[6]

The population-based measures yield information about the structure of jobs that a given group of individuals hold; the employment-based measures supply information about the structure of jobs that a given group of workers hold.

The population-based measures are not without problems of interpretation. While holding the base group of individuals fixed avoids cyclical problems of movement in and out of employment, secular changes in labor supply directly affect the population-based measures. If a group has increased its labor supply over time (e.g., as women have done), the population-based measures for that group are likely to show an increase. Similarly, if a group has decreased its labor supply over time (e.g., as older men have done), the population-based measures for that group are likely to show a decrease. Changes in population-based measures due to shifts in labor supply do not reflect changes in the underlying structure of jobs. In what follows, I present statistical results on both an employment and a population basis.

5.3 Changes in Interpolated Quantiles, 1973–93

Because the age distribution of the population has changed over time and because job durations are strongly related to age, it is important to control for age when examining the distribution of job durations over time. A visual representation of changes in the distribution of job durations over time is given in figure 5.1. This figure contains plots of four weighted (by CPS sampling weights) interpolated quantiles (.25, .5, .75, .9) of the employment-based tenure distribution by year overall and broken down by sex and four 10-year age categories. This and succeeding figures do not show sampling errors. Sampling errors for these interpolated quantiles, calculated using the delta method, are generally about 0.15 years. Thus statistical significance requires differences across calendar years of about 0.4 years.

Not surprisingly, all four employment-based quantiles in figure 5.1 rise systematically with age. The plots for males look quite flat, with perhaps a slight decline for the upper quantiles of the oldest age category. The plots for females show some upward movement over time. The combined plots (no distinction by sex) look very flat. Analogous plots of population-based quantiles are con-

6. Note that the population-based fraction of individuals on a job less than or equal to one year includes those not employed in both the numerator and the denominator. This is clear from the coding of job durations of those not employed as zero. The resulting probability has a natural interpretation.

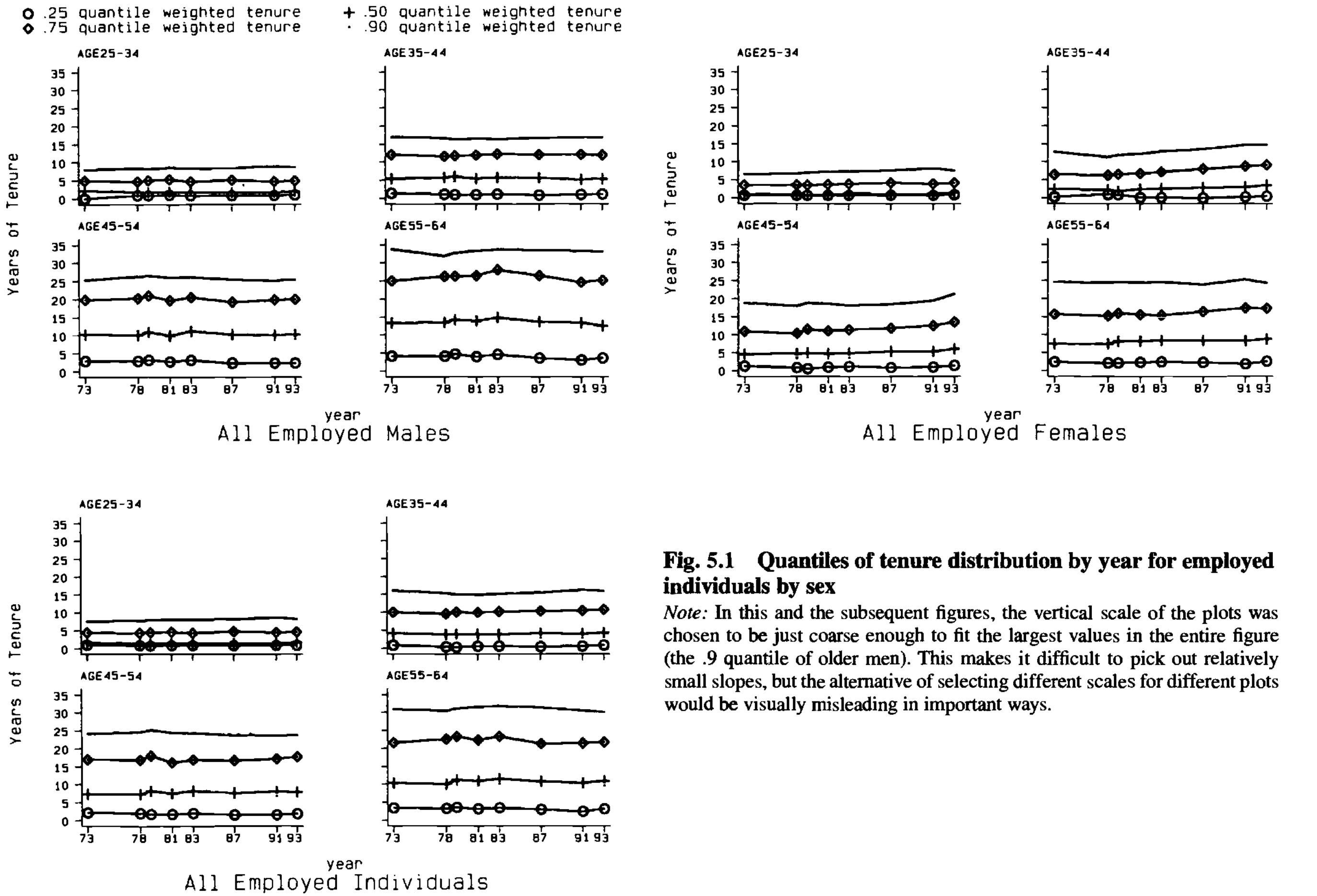

Fig. 5.1 Quantiles of tenure distribution by year for employed individuals by sex

Note: In this and the subsequent figures, the vertical scale of the plots was chosen to be just coarse enough to fit the largest values in the entire figure (the .9 quantile of older men). This makes it difficult to pick out relatively small slopes, but the alternative of selecting different scales for different plots would be visually misleading in important ways.

tained in figure 5.2. These look much like the employment-based quantiles in figure 5.1 with these exceptions: (1) there is fairly substantial upward movement in the population-based quantiles for women, and (2) there is somewhat more decline in the quantiles for older males. These changes largely represent systematic changes in labor force participation. The decrease in the frequency with which women withdraw from the labor force is doubtless an important factor in their increased job duration. The move toward earlier retirement underlies an important part of the decline in population-based measures of job duration among men aged 55–64.

Appendix tables 5A.1 through 5A.4 contain the raw data underlying the median and .9 quantiles for figures 5.1 and 5.2. Table 5A.1, which contains employment-based medians, also includes tabulations of medians by sex and age category based on the January mobility supplements for 1951, 1963, 1966, and 1968.[7] Aside from the fact that age-adjusted medians in 1951 were much lower than later, probably because many workers had to "restart" after returning from World War II, long-term trends using this longer time series are difficult to discern.

Figures 5.3, 5.4, and 5.5 contain plots of the four employment-based quantiles broken down by age and education. Figure 5.3 makes no distinction between sexes. It shows a substantial decline in job duration for workers in the lowest educational category (less than 12 years). Not much change is evident in the overall quantiles in the higher educational categories. Figure 5.4 replicates these plots for males. The substantial changes here are a decline in job duration for the least educated men and some decline for the oldest highly educated men (16 years or more). Figure 5.5 replicates these plots for females. It is interesting that there does not seem to be much decline in job duration for the least educated women. The plots also suggest that there is a fairly systematic increase in job duration for women in the three higher educational categories. This is a consequence of the decreased frequency with which women withdraw from the labor force, and it suggests that there is an increased incidence of long-term stable employment for women.

Figures 5.6, 5.7, and 5.8 replicate these plots using population-based quantiles. Here the results are more striking. There is a sharp drop in the population-based quantiles for the least educated individuals. This is attributable to a decline in job duration among men (fig. 5.7). Thus the well-known deterioration in labor market conditions for poorly educated men resulted not only in shorter jobs but also in a scarcity of jobs themselves. The quantiles of the employment-based job duration distributions for more highly educated men look fairly stable. There is also a sharp increase in job duration for women in the top three educational categories (fig. 5.8). Once again, this largely reflects the decreased frequency with which women withdraw from the labor force.

In order to provide a clearer statistical summary of changes over time in the

7. The sources for these published tabulations are Department of Labor (1963, 1967, 1969).

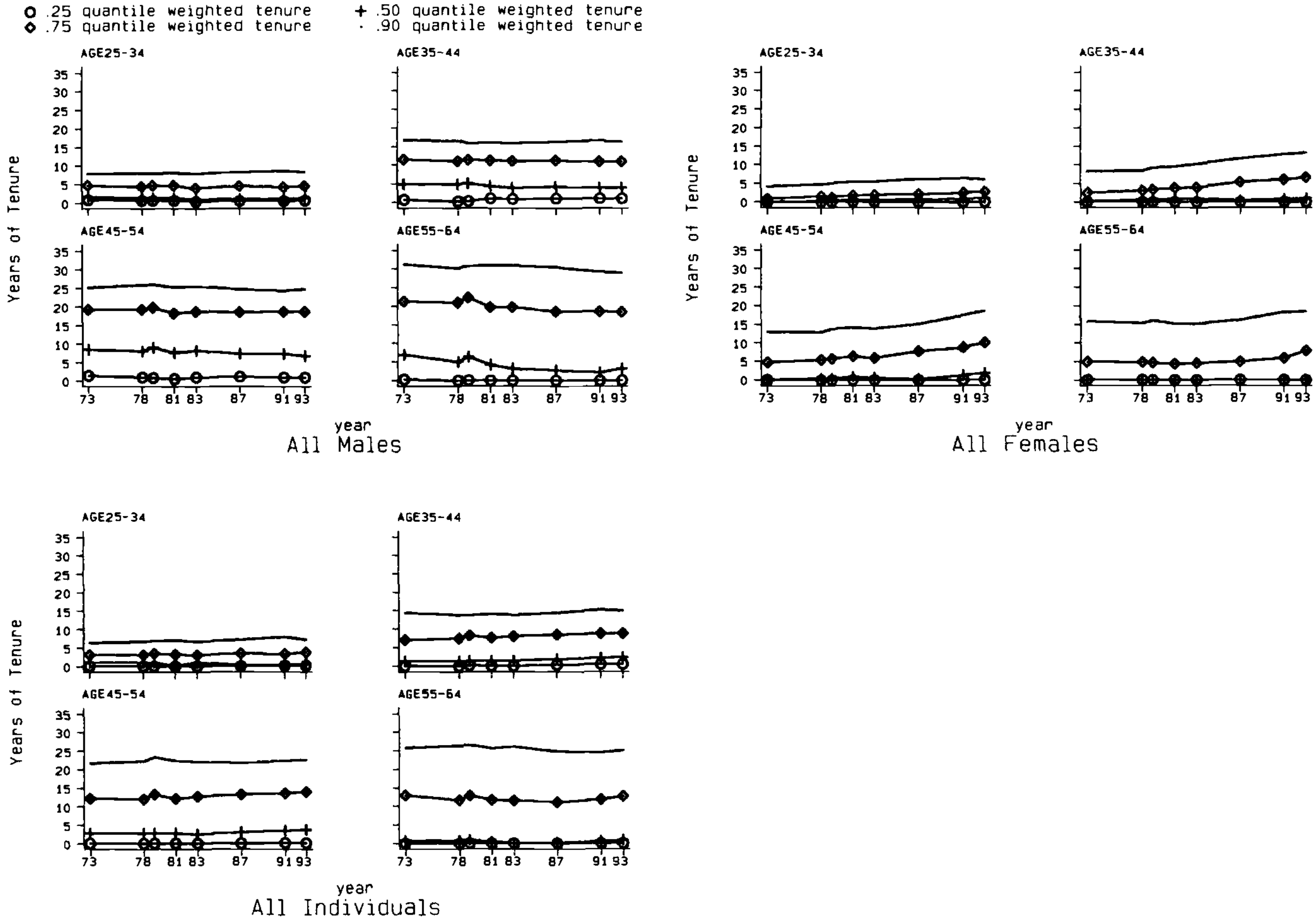

Fig. 5.2 Quantiles of tenure distribution by year for all individuals by sex

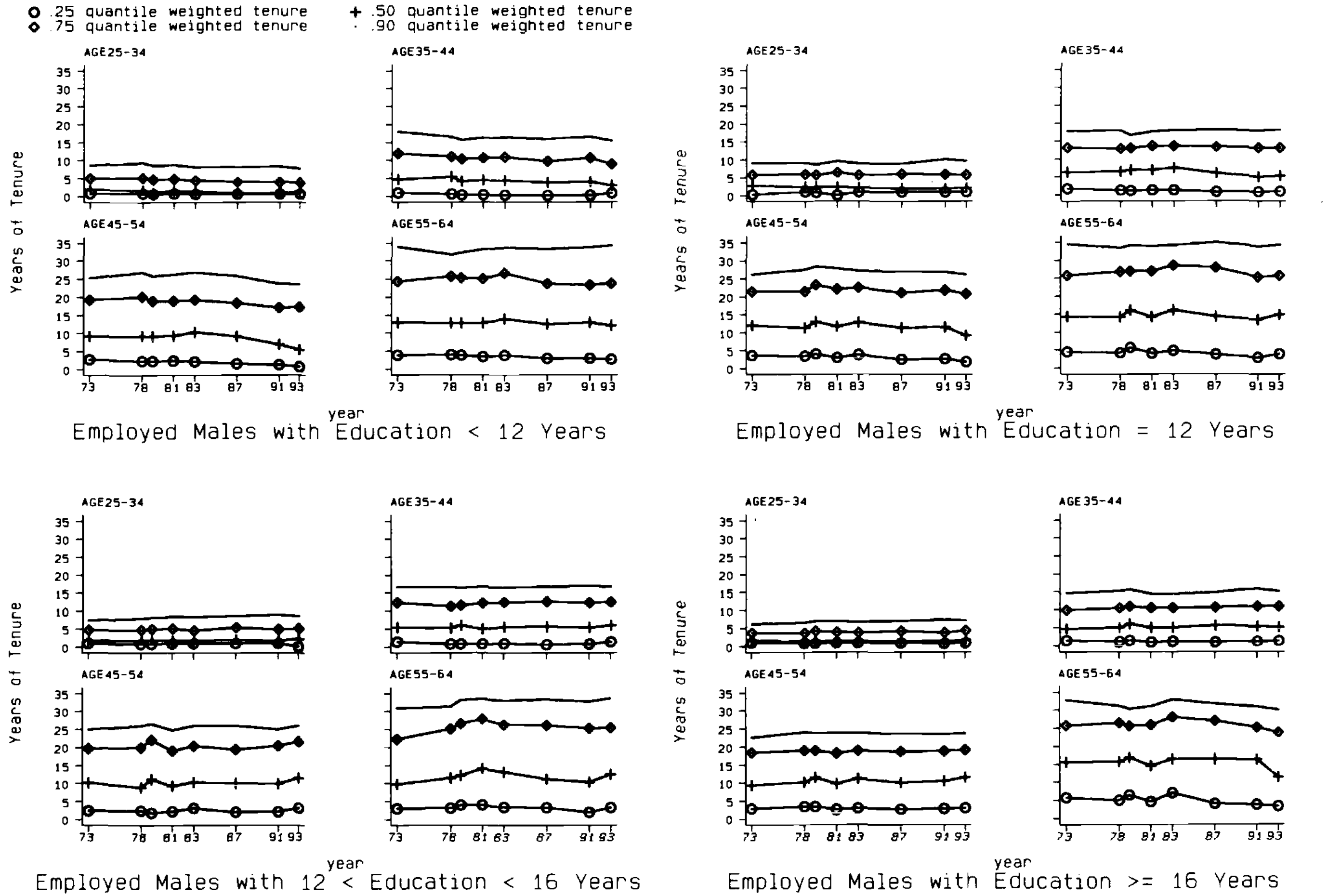

Fig. 5.3 Quantiles of tenure distribution by year for all employed individuals by education

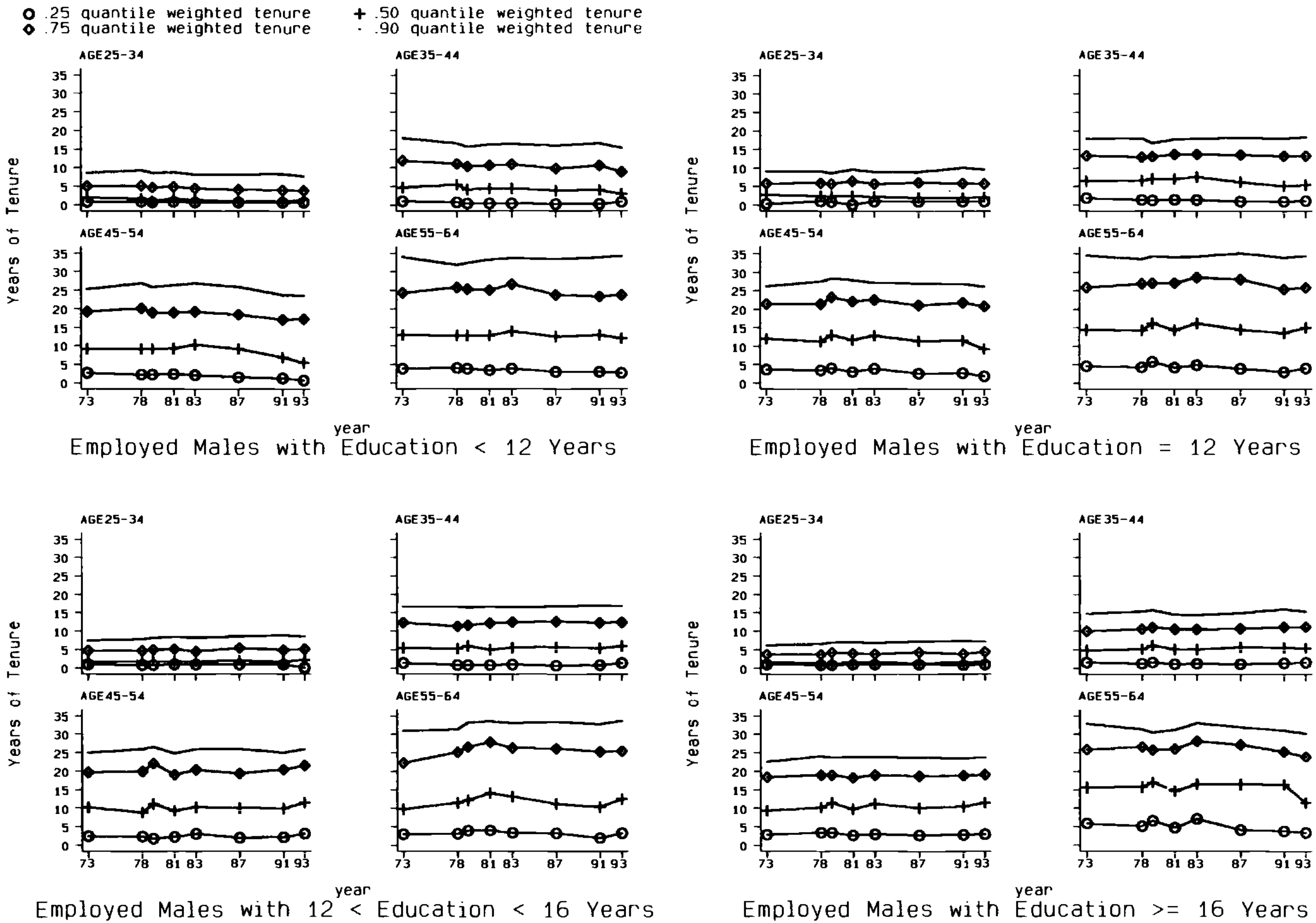

Fig. 5.4 Quantiles of tenure distribution by year for employed males by education

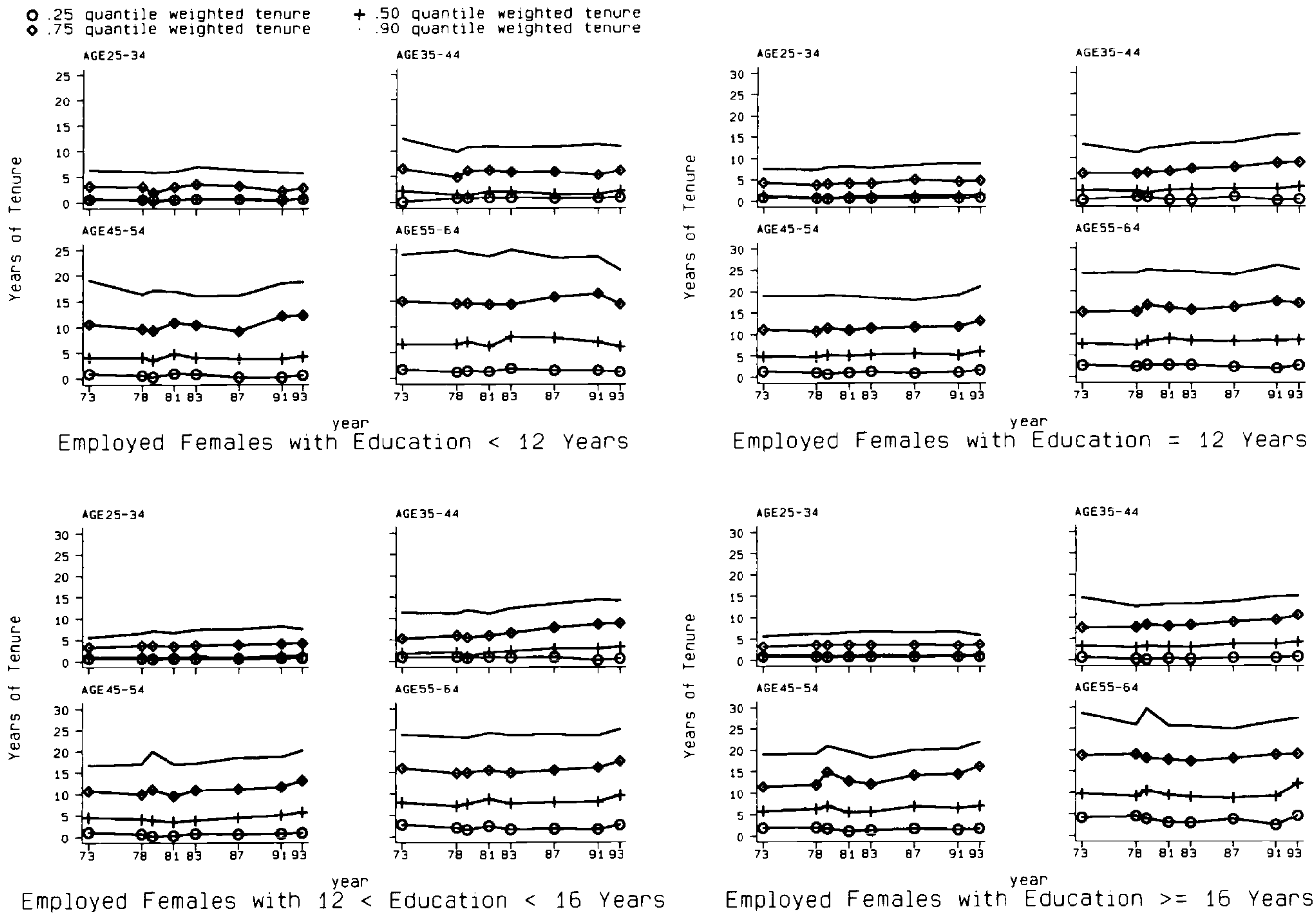

Fig. 5.5 Quantiles of tenure distribution by year for employed females by education

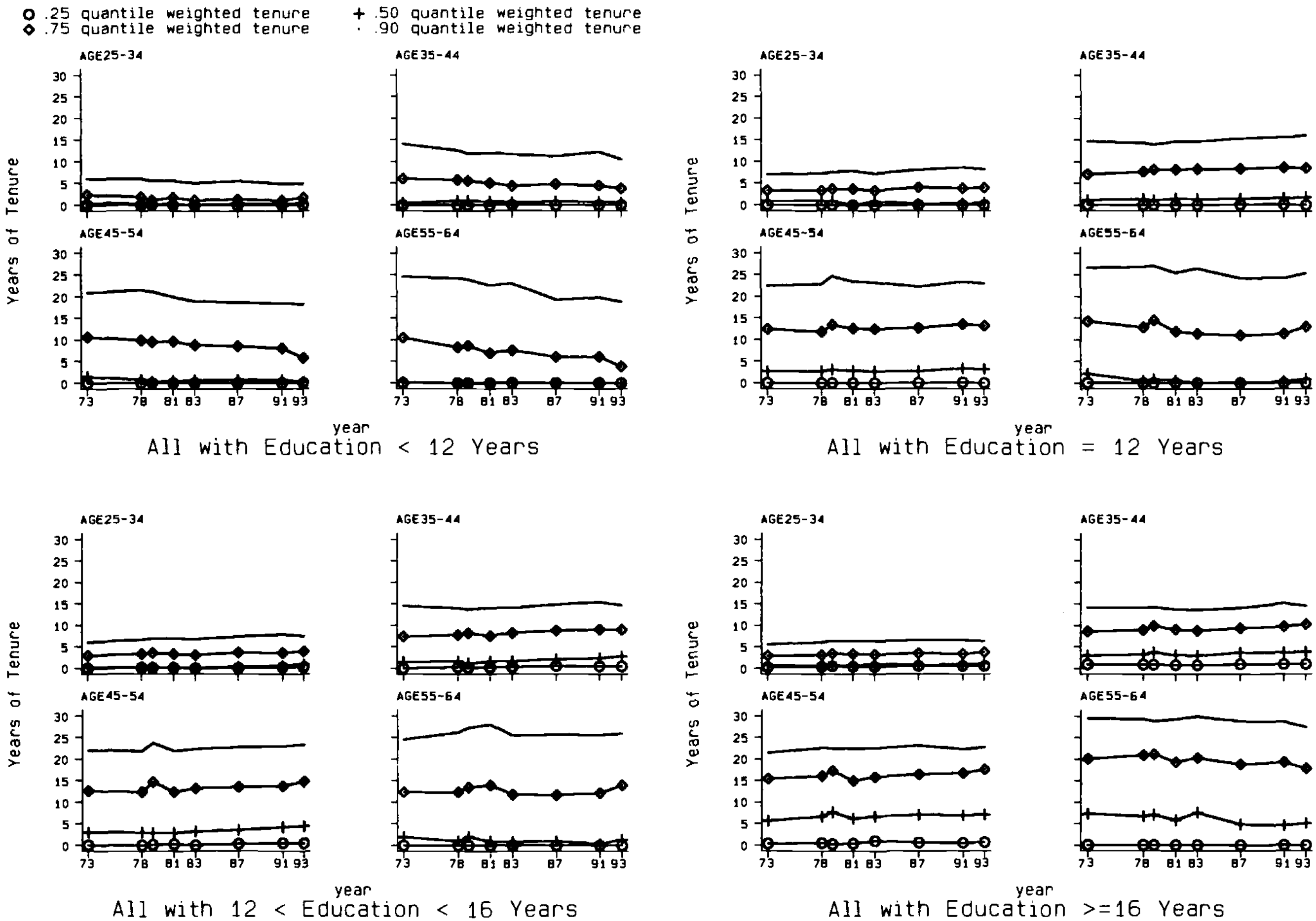

Fig. 5.6 Quantiles of tenure distribution by year for all individuals by education

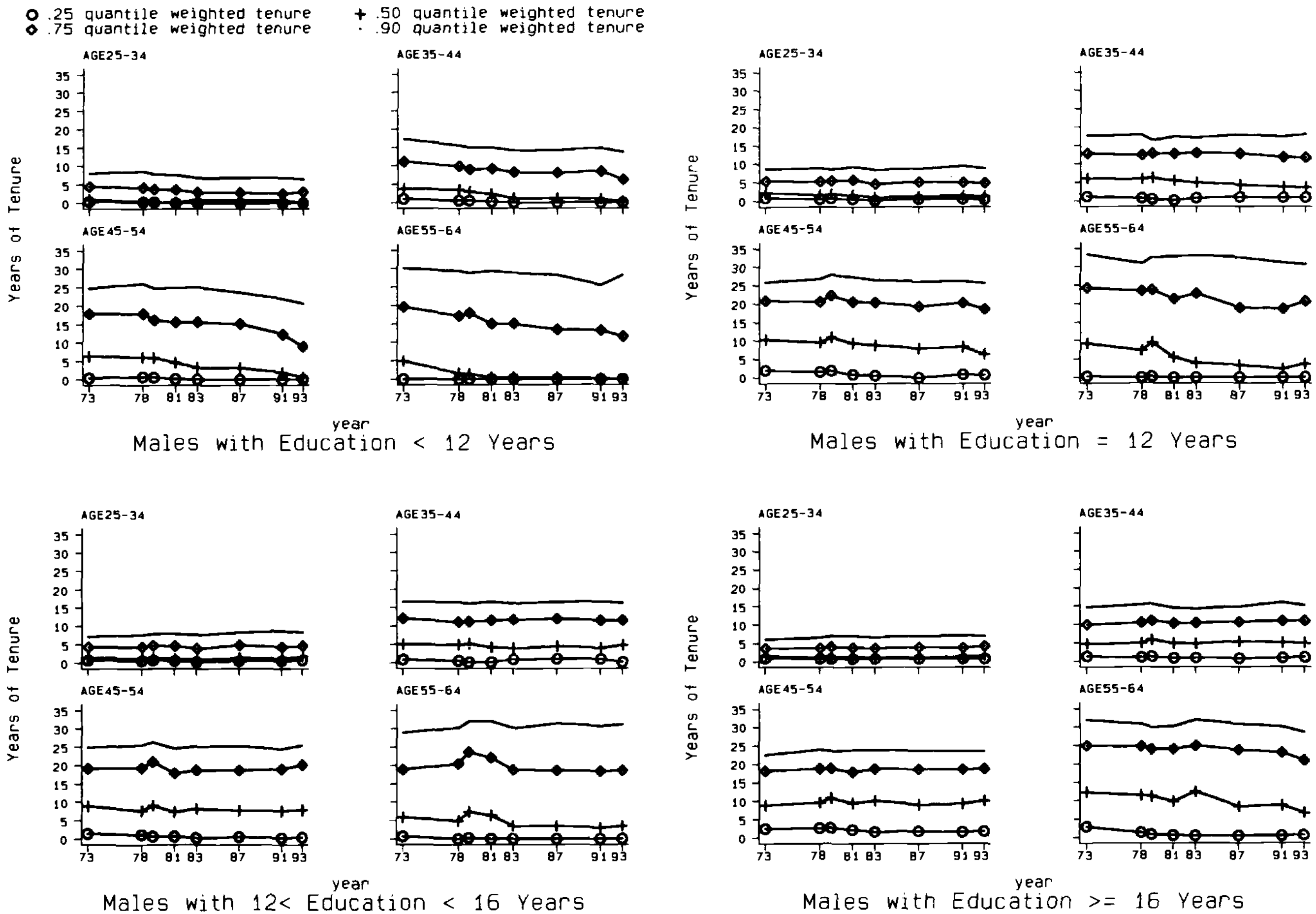

Fig. 5.7 Quantiles of tenure distribution by year for all males by education

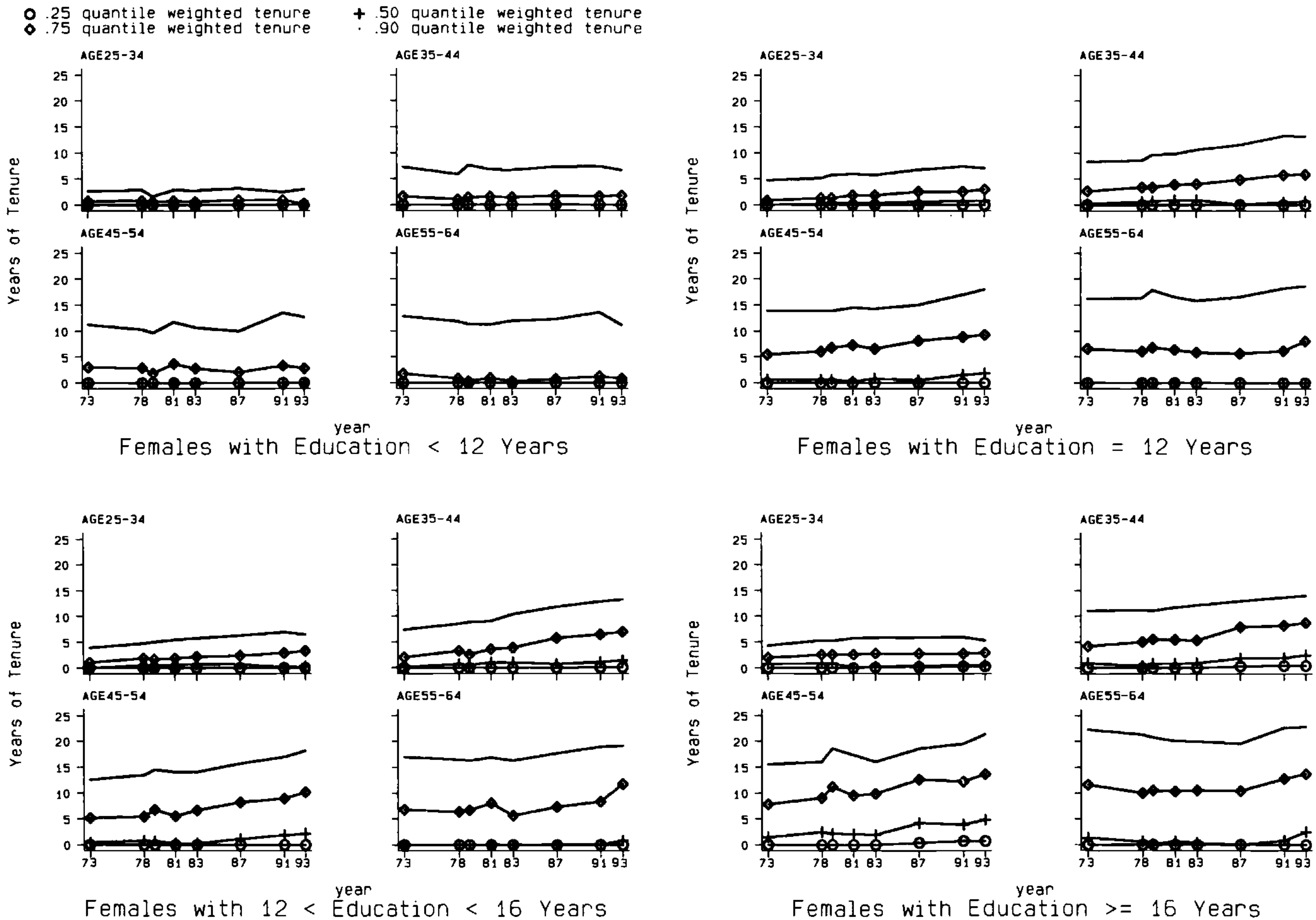

Fig. 5.8 Quantiles of tenure distribution by year for all females by education

quantiles of the distribution of job tenures, tables 5.1, 5.2, and 5.3 contain cell-based regressions of the employment-based quantiles. I compute weighted employment-based medians for cells defined by nine five-year age categories (from age 21 through age 65), four educational categories (less than 12 years, 12 years, between 12 and 16 years, and 16 years or more), and eight calendar years. I do this separately for three samples (employed individuals, employed males, and employed females). The procedure is to specify a linear model that determines the cell quantiles as a function of a set of observable characteristics of the cells.[8] Such a model for the τth quantile of observations in cell j would be

$$\theta_{\tau j} = X_j\beta + \varepsilon_{\tau j}, \qquad (3)$$

where $\theta_{\tau j}$ is the τth quantile of observations in cell j, X_j is a vector of observable characteristics for cell j, β is a vector of parameters, and $\varepsilon_{\tau j}$ is an unobserved component. This parameters of this model can be estimated using weighted least squares. One choice of weights is to use the estimated variances of cell quantiles as weights. Another choice is simply to use the number of observations in each cell as weights. Chamberlain (1994) suggests that it may be better to use the cell sizes as weights if it is possible that the model is misspecified. Since I am maintaining the specification for the cell quantiles in equation (3), I weight by cell size.

The X_j vector in tables 5.1 and 5.2 contains eight dummy variables for the age categories, three dummy variables for the educational categories, and one of two specifications of calendar year. One specification (in the odd-numbered columns) contains a complete set of eight calendar year dummy variables (and hence no constant). The other (in the even-numbered columns) contains a linear time trend (calendar year itself) and a constant. I do not present the estimates of the age effects. Not surprisingly, they have a great deal of explanatory power, with older workers having longer job durations. I focus here on the year effects.

In most cases, it is not possible to reject the single variable representation of year effects in the form of a time trend against the unconstrained dummy variable model. As such, most of the subsequent discussion will focus on models with time trends. It is also worth noting that variation in the quantiles across cells is fairly well explained by the main effects specifications used in that the R^2 of these regressions are quite large (over .95).

The estimates in columns (1) and (2) of table 5.1 show no significant relationship between employment-based median job duration and calendar year, either in the unconstrained dummy variable specification or with a single time trend. The estimates in columns (3) and (4) show a marginally significant small negative time trend in median job duration for males only. In contrast, the estimates in columns (5) and (6) show a larger positive time trend in median job duration for females only. These point estimates suggest an average overall

8. Chamberlain (1994) developed this technique for estimating quantiles.

Table 5.1 Median Regression of Job Duration for Employed Individuals Aged 21–64

	All		Males		Females	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant		.450		2.63		−2.12
		(.537)		(.794)		(.460)
Year		.00240		−.0179		.0332
		(.00639)		(.00941)		(.0055)
1973	.689		1.16		.502	
	(.156)		(.230)		(.131)	
1978	.539		1.11		.342	
	(.150)		(.223)		(.123)	
1979	.731		1.55		.414	
	(.189)		(.279)		(.157)	
1981	.585		1.13		.541	
	(.146)		(.219)		(.118)	
1983	.761		1.52		.639	
	(.151)		(.226)		(.122)	
1987	.633		1.06		.794	
	(.606)		(.226)		(.122)	
1991	.606		.831		.808	
	(.154)		(.232)		(.125)	
1993	.829		.870		1.26	
	(.192)		(.289)		(.155)	
Ed < 12	−.732	−.732	−1.69	−1.70	−.852	−1.54
	(.111)	(.111)	(.159)	(.161)	(.0972)	(.266)
12 < Ed < 16	−.230	−.229	−.863	−.861	−.211	−.649
	(.100)	(.100)	(.150)	(.152)	(.0816)	(.238)
Ed ≥ 16	.570	.571	−.621	−.614	.515	.514
	(.0999)	(.0997)	(.145)	(.147)	(.0851)	(.0864)
p-Value equality of year effects	.646		.0233		<.00005	
p-Value year effects equal trend		.548		.0492		.0273
No. of cells	288	288	288	288	288	288
No. of observations	378,890	378,890	214,210	214,210	164,680	164,680
R^2	.970	.969	.964	.963	.959	.957

Notes: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantile for nine age categories, four educational categories, two sex categories (in col. 3–6), and eight years. Only observations with nonzero quantiles (employed) are included. All observations are weighted by the cell size.

decrease over the 20-year period studied of about 0.35 years in the median for men and an average overall increase of about 0.7 years in the median for women over the same period.

The estimates in table 5.2 for the .9 quantile of the employment-based distribution of job durations show a similar pattern. There is no significant relationship between year and the .9 quantile of job duration when no sex distinction is made, and there is actually a small *increase* on average in the .9 quantile for

Table 5.2 .9 Quantile Regression of Job Duration for Employed Individuals Aged 21–64

	All		Males		Females	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant		3.03		2.88		−2.76
		(.587)		(.549)		(1.02)
Year		.00839		.0138		.0734
		(.00698)		(.00651)		(.0122)
1973	3.66		3.83		3.19	
	(.172)		(.162)		(.287)	
1978	3.63		3.98		2.62	
	(.164)		(.157)		(.270)	
1979	3.63		3.93		3.12	
	(.208)		(.197)		(.344)	
1981	3.80		4.08		3.15	
	(.160)		(.154)		(.260)	
1983	3.70		4.00		3.16	
	(.165)		(.159)		(.268)	
1987	3.71		4.07		3.35	
	(.165)		(.159)		(.267)	
1991	3.85		4.17		4.21	
	(.169)		(.164)		(.274)	
1993	3.76		4.03		4.40	
	(.211)		(.204)		(.341)	
Ed < 12	−1.13	−1.13	−1.06	−1.06	−1.75	−1.74
	(.122)	(.121)	(.112)	(.111)	(.213)	(.219)
12 < Ed < 16	−.965	−.966	−1.16	−1.16	−.763	−.763
	(.110)	(.109)	(.106)	(.105)	(.179)	(.183)
Ed ≥ 16	−2.07	−2.07	−2.88	−2.88	−.564	.560
	(.110)	(.109)	(.102)	(.102)	(.187)	(.191)
p-Value equality of year effects	.873		.486		<.00005	
p-Value year effects equal trend		.945		.914		.0043
No. of cells	288	288	288	288	288	288
No. of observations	378,890	378,890	214,210	214,210	164,680	164,680
R^2	.995	.995	.996	.996	.974	.972

Notes: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantiles for nine age categories, four educational categories, two sex categories (in col. 3–6), and eight years. Only observations with nonzero quantiles (employed) are included. All observations are weighted by the cell size.

males (about 0.3 years over the 20-year period). The rate of increase in the .9 quantile of job duration for females (about 1.5 years over the 20-year period) is substantially larger than the rate of increase in the women's median.

Important differences in time trends of job duration by educational category were apparent in the figures, particularly for men, and the specification in the first two tables does not allow for these differences. In order to address this problem directly, I reestimated the models with time trends in tables 5.1 and 5.2 with the time trend interacted with the four educational categories. Table 5.3 contains estimates of the relevant parameters. These results are quite clear-cut, and they support and sharpen the visual impression from the figures. Workers with less than 12 years of education suffered a decline in median job duration of over 0.5 years on average over the 20-year period. This seems almost entirely accounted for by less educated males, who suffered a decline in median job duration of almost one full year on average over this period. Men with less than 12 years of education and men with exactly 12 years of education shared this decline. Among workers with more than a high school education, job duration increased on average. There was no significant increase in medians for more educated males (more than 12 years) on average, but the .9 quantile of the job duration distribution did increase significantly for more educated men (about 0.5 years over the 20-year period). In contrast, both quantiles increased substantially for women with at least a high school education. Depending on education level, the increase in the medians over the 20-year period range from about 0.5 years to about 1 year. The increase in the .9 quantiles for women over this period was even larger, ranging from 1.5 years to over 2 years.

Tables 5.4, 5.5, and 5.6 repeat the entire cell quantile regression analysis using population-based quantiles. Recall that these quantiles ought to be less affected by cyclical fluctuations but more affected by secular changes in labor supply. The cell quantile regression model is particularly well suited for this analysis because it allows a natural treatment of those not employed, all of whom are coded as having zero job duration. Effectively, these are censored observations, and any cell for which the particular quantile of the job duration distribution being studied is zero (i.e., is represented by a nonemployed individual) contains no information about the process that generates the cell quantiles.[9]

The results for the population-based quantiles are roughly similar to those for the employment-based quantiles, but there are some differences. Most striking is the substantial decline in the population-based median for males (about 1.6 years over the 20-year period), shown in column (4) of table 5.4. There is also a larger increase in the population-based .9 quantile for females (about 2.5 years over the 20-year period), shown in column (6) of table 5.5. The sources of these substantial trends become clearer with separate year effects by education in table 5.6. The large decrease in the median for males seems to be due almost entirely to individuals with at most a high school education. These individuals have median durations that declined by 2.2 to 3.2 years over the 20-year period. There was no significant change in median job duration for males with more than a high school education. The median job duration for

9. Chamberlain (1994) shows that it is appropriate to estimate the cell quantile regression model using only observations for which the cell quantile is not censored, and I follow this procedure.

Table 5.3 **Quantile Regression of Job Duration for Employed Individuals Aged 21–64 (year by education interaction)**

	All		Males		Females	
Variable	Median (1)	.9 Quantile (2)	Median (3)	.9 Quantile (4)	Median (5)	.9 Quantile (6)
Constant	.902	2.36	4.70	2.75	−1.70	−3.89
	(.854)	(.911)	(1.30)	(.904)	(.697)	(1.52)
Ed < 12	1.35	5.06	−1.60	1.89	.852	8.12
	(1.52)	(1.63)	(2.19)	(1.52)	(1.36)	(2.98)
12 < Ed < 16	−2.30	−2.19	−5.06	−1.87	−2.54	−2.88
	(1.40)	(1.49)	(2.10)	(1.45)	(1.17)	(2.57)
Ed ≥ 16	−.900	−2.45	−5.40	−3.94	−.117	.401
	(1.40)	(1.49)	(2.04)	(1.42)	(1.23)	(2.68)
(Ed < 12)*Year	−.0288	−.0595	−.0446	−.0210	.0072	−.0337
	(.0156)	(.0166)	(.0218)	(.0151)	(.0144)	(.0314)
(Ed = 12)*Year	−.0029	.0167	−.0428	.0155	.0283	.0873
	(.103)	(.109)	(.0157)	(.0109)	(.0084)	(.0183)
(12 < Ed < 16)*Year	.0219	.0312	.0079	.0239	.0560	.112
	(.0133)	(.0142)	(.0198)	(.0137)	(.0112)	(.0246)
(Ed ≥ 16)*Year	.0147	.0212	.0149	.0283	.0359	.0756
	(.0133)	(.0142)	(.0189)	(.0131)	(.0120)	(.0262)
p-Value equality of year effects	.0656	.0002	.0352	.0722	.0529	.0025
No. of cells	288	288	288	288	288	288
No. of observations	378,890	378,890	214,210	214,210	164,680	164,680
R^2	.970	.995	.964	.996	.958	.974

Notes: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantile for nine age categories, four educational categories, two sex categories (in cols. 3–6), and eight years. Only observations with nonzero quantiles (employed) are included. All observations are weighted by the cell size. All specifications include eight dummy variables for age categories.

Table 5.4 Median Regression of Job Duration for All Individuals Aged 21–64

Variable	All (1)	All (2)	Males (3)	Males (4)	Females (5)	Females (6)
Constant		−1.32		7.84		−2.90
		(.851)		(1.13)		(.604)
Year		.0201		−.0819		.0387
		(.0102)		(.0134)		(.0073)
1973	.337		2.02		−.0393	
	(.241)		(.325)		(.179)	
1978	.259		1.40		.206	
	(.234)		(.314)		(.155)	
1979	.352		1.92		.174	
	(.295)		(.396)		(.195)	
1981	.174		1.13		.212	
	(.232)		(.305)		(.151)	
1983	.158		.722		.223	
	(.239)		(.312)		(.155)	
1987	.409		.699		.421	
	(.239)		(.317)		(.155)	
1991	.589		.522		.643	
	(.246)		(.323)		(.158)	
1993	.792		.431		.829	
	(.302)		(.402)		(.196)	
Ed < 12	−1.28	−1.26	−2.82	−2.81	−.721	−.719
	(.174)	(.174)	(.225)	(.225)	(.444)	(.437)
12 < Ed < 16	.264	.263	−.506	−.504	.177	.176
	(.159)	(.159)	(.214)	(.215)	(.100)	(.0988)
Ed ≥ 16	1.76	1.76	.553	.552	.811	.811
	(.165)	(.165)	(.214)	(.214)	(.108)	(.107)
p-Value equality of year effects	.465		<.00005		<.00005	
p-Value year effects equal trend		.605		.269		.624
No. of cells	262	262	282	282	189	189
No. of observations	502,600	502,600	253,860	253,860	204,050	204,050
R^2	.689	.680	.849	.845	.454	.447

Notes: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantiles for nine age categories, four educational categories, two sex categories (in cols. [3]–[6]), and eight years. Only observations with nonzero quantiles (employed) are included. All observations are weighted by the cell size.

Table 5.5 .9 Quantile Regression of Job Duration for All Individuals Aged 21–64

	All		Males		Females	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant		3.98		6.11		−7.89
		(.985)		(.874)		(1.12)
Year		−.0117		−.0298		.126
		(.0117)		(.0104)		(.0134)
1973	3.17		3.90		1.69	
	(.284)		(.255)		(.318)	
1978	3.13		3.82		1.70	
	(.274)		(.246)		(.309)	
1979	3.31		3.74		2.18	
	(.344)		(.311)		(.384)	
1981	3.01		3.78		2.25	
	(.267)		(.239)		(.300)	
1983	2.80		3.50		2.29	
	(.273)		(.245)		(.308)	
1987	2.77		3.53		2.91	
	(.278)		(.249)		(.313)	
1991	3.05		3.44		3.85	
	(.284)		(.254)		(.320)	
1993	3.08		3.28		3.99	
	(.352)		(.315)		(.397)	
Ed < 12	−2.80	−2.80	−2.29	−2.29	−3.71	−3.70
	(.188)	(.188)	(.171)	(.170)	(.210)	(.211)
12 < Ed < 16	−.484	−.481	−1.11	−1.11	−.0953	−.0946
	(.188)	(.187)	(.169)	(.168)	(.211)	(.212)
Ed ≥ 16	−.710	−.710	−2.45	−2.45	1.08	1.08
	(.195)	(.195)	(.169)	(.167)	(.232)	(.233)
p-Value equality of year effects	.595		.229		<.00005	
p-Value year effects equal trend		.605		.972		.173
No. of cells	288	288	288	288	288	288
No. of observations	550,940	550,940	260,360	260,360	290,580	290,580
R^2	.981	.981	.990	.989	.941	.939

Notes: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantiles for nine age categories, four educational categories, two sex categories (in cols. [3]–[6]), and eight years. Only observations with nonzero quantiles (employed) are included. All observations are weighted by the cell size.

Table 5.6 **Quantile Regression of Job Duration for All Individuals Aged 21–64 (year by education interaction)**

	All		Males		Females	
Variable	Median (1)	.9 Quantile (2)	Median (3)	.9 Quantile (4)	Median (5)	.9 Quantile (6)
Constant	−1.12	1.47	10.7	4.70	−1.98	−10.3
	(1.35)	(1.47)	(1.83)	(1.36)	(.855)	(1.67)
Ed < 12	.164	11.1	.835	9.33	1.46	7.13
	(2.40)	(2.41)	(2.99)	(2.16)	(7.52)	(2.84)
12 < Ed < 16	−.682	−1.83	−8.14	−3.25	−.644	−1.51
	(2.22)	(2.44)	(2.92)	(2.18)	(1.44)	(2.90)
Ed ≥ 16	.883	−1.81	−7.66	−5.42	−2.31	3.53
	(2.31)	(2.56)	(2.93)	(2.18)	(1.53)	(3.21)
(Ed < 12)*Year	−.0041	−.151	−.161	−.155	.0033	.0225
	(.0246)	(.0236)	(.0292)	(.0206)	(.0856)	(.0282)
(Ed = 12)*Year	.0178	.0190	−.116	−.0124	.0277	.155
	(.0162)	(.0176)	(.0219)	(.0163)	(.0103)	(.0201)
(12 < Ed < 16)*Year	.0291	.0348	−.0238	.0131	.0376	.172
	(.0212)	(.0234)	(.0274)	(.0204)	(.0138)	(.0283)
(Ed ≥ 16)*Year	.0282	.0318	−.0169	.0231	.0650	.126
	(.0225)	(.0251)	(.0274)	(.0204)	(.0151)	(.0326)
p-Value equality of year effects	.860	<.0001	.0004	<.0001	.0058	.0005
No. of cells	262	288	282	288	189	288
No. of observations	502,600	550,940	253,860	260,360	204,050	290,580
R^2	.681	.984	.855	.991	.460	.943

Notes: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantile for nine age categories, four educational categories, two sex categories (in cols. [3]–[6]), and eight years. Only observations with nonzero quantiles (employed) are included. All observations are weighted by the cell size. All specifications include eight dummy variables for age categories.

women increases monotonically with educational category, rising from zero for women with less than a high school education to an increase of about 1.3 years over the 20-year period for women with at least 16 years of education. The large increase in the .9 quantile for women was shared across all but the lowest educational category.

Overall, the results in this section show a clear pattern. There has not been much change in the quantiles of the overall distribution of job durations that I studied. However, important changes have taken place in the distribution of job durations for particular subgroups. There are two striking changes: (1) the quantiles of the job duration distribution for the least educated workers, and especially the least educated men, have declined substantially, and (2) the quantiles of the job duration distribution for women, and especially women with more education, have increased substantially.

5.4 Changes in Probabilities of Short-Term and Long-Term Jobs, 1973–93

It is useful to examine specific points of the cumulative distribution function of job durations in order to determine if the same changes found in the quantiles can be measured there. In particular, I examine (1) the fraction of job durations less than or equal to 1 year, (2) the fraction of job durations greater than 10 years, and (3) the fraction of job durations greater than 20 years. Based on the results reported above, it is reasonable to expect that the fraction of short-term jobs (up to 1 year) has grown for the least educated workers (especially for the least educated males) and declined among females (especially those with more than a high school education). Analogously, the fraction of long-term jobs (more than 10 years and more than 20 years) has declined among the least educated male workers and increased among more highly educated females. Given the lack of a pattern in the non-sex-specific quantiles over time, no clear change in the aggregate fractions in these categories is expected.

5.4.1 Employment-Based Probabilities

Appendix tables 5A.5, 5A.6, and 5A.7 present information on the employment-based fraction of workers with job durations in the specified intervals broken down by crude age category, sex, and year. It is difficult to pick out clear trends in these data other than to note that employed females have become less likely to have been in their jobs a short time and have become more likely to have been in their jobs for a substantial length of time.

These tables also show that the probability of being in a new job and the probability of having been on the job for a substantial length of time increase with age. This is so because it is virtually impossible for very young workers to have been on their job for more than 10 or 20 years. While the logit analysis that follows includes detailed controls for age, it makes sense to (1) estimate the logit model of the probability of job duration of more than 10 years on the sample of workers who are at least 35 years old and (2) estimate the logit model of the probability of job duration of more than 20 years on the sample of workers who are at least 45 years old.

Tables 5.7, 5.8, and 5.9 contain estimates of logit models of the employment-based probabilities. The aim of this analysis is to provide summary measures of time trends in the probabilities and to examine variation in these trends across educational categories.

Table 5.7 contains estimates of logit models of the employment-based probability that a worker has been on his or her job no more than one year. The estimates in the odd-numbered columns are for models that contain a linear time trend (calendar year), eight dummy variables for age categories, four dummy variables for educational categories, and a constant. The estimates in the even-numbered columns are for models that include the same variables but allow for a separate time trend for each of the four educational categories. When no distinction is made by sex, there is a slight but significant upward trend in the probability that a job is no more than one year old. Over the 20-year period, the employment-based probability that a job is no more than one year old is predicted to have increased by about 1.3 percentage points.[10] This aggregate figure masks a larger increase for men over the 20-year period of about 3 percentage points and a small decrease for women over the 20-year period of about 1.6 percentage points.

With separate time trends by educational category, a much sharper picture emerges. The hypothesis that the time trends are the same across educational categories can be rejected in all cases. The results suggest that the overall increase in the probability of short durations is due entirely to the two lowest educational categories. The probability of a worker with less than a high school education being in a short-term job is predicted to be about 6 percentage points higher in 1993 than in 1973. This is a substantial change given that the overall probability of being in a short-term job is about .25.

An analysis of the trends separately for men and women suggests that this result is driven by a large increase in the short-term job probability for men with no more than a high school education. Men with less than a high school education have a probability of being in a short-term job that is predicted to be about 8.5 percentage points higher in 1993 than in 1973. The change is somewhat smaller but still quite substantial for men with exactly a high school education (an increase of 5 percentage points).

There has been some decrease in the short-term job probability in the higher

10. The logit coefficient of 0.0034 must be multiplied by some estimate of $p(1-p)$ when one is computing the derivative of the probability with respect to year. A reasonable mean estimate of $p(1-p)$ is 0.2. Thus, over the 20-year period, the probability that a worker was in his or her job for no more than one year is predicted to have increased by about 1.4 percentage points ($0.0034 \times 0.2 \times 20 \times 100$). The value of 0.2 for $p(1-p)$ is used in what follows to adjust the logit coefficient for the employment-based models. A cautionary note is that the underlying probabilities (and hence the appropriate $p(1-p)$) vary, and the percentage point changes mentioned in the text are, of necessity, approximations.

Table 5.7 **Logit Analysis of Probability of Job Duration One Year or Less for Employed Individuals Aged 21–64 (year by education interaction)**

	All		Males		Females	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant	−2.66	−2.91	−3.16	−3.56	−1.90	−2.22
	(.0587)	(.0910)	(.0806)	(.130)	(.0864)	(.128)
Ed < 12	.293	−.427	.340	−.339	.323	−.0849
	(.0120)	(.155)	(.0163)	(.209)	(.0183)	(.238)
12 < Ed < 16	.0686	1.07	.100	1.40	.0668	1.11
	(.0104)	(.138)	(.0147)	(.194)	(.0147)	(.198)
Ed ≥ 16	.0068	.613	.0904	1.03	−.0176	.596
	(.0107)	(.143)	(.0148)	(.196)	(.0156)	(.212)
Year	.0034		.0080		−.0043	
	(.0006)		(.00087)		(.0009)	
(Ed < 12)*Year		.0153		.0212		.0046
		(.0016)		(.0020)		(.0025)
(Ed = 12)*Year		.0065		.0128		−.0004
		(.0010)		(.0015)		(.0015)
(12 < Ed < 16)*Year		−.0053		−.0026		−.0127
		(.0013)		(.0017)		(.0018)
(Ed ≥ 16)*Year		−.0006		.0016		−.0077
		(.0013)		(.0018)		(.0020)
p-Value equality of time trends		<.0001		<.0001		<.0001
No. of observations	378,892	378,892	214,211	214,211	164,681	164,681
Log *L*	−194,019.8	−193,957.1	−102,785.3	−102,734.3	−90,374.8	−90,352.9

Notes: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling one if job duration is less than or equal to one year. All models include controls for education (three dummy variables for four categories) and age (eight dummy variables for nine categories). The analysis is weighted using CPS sampling weights. The included age range is 21–64.

educational categories. This is driven by a decrease in this probability for highly educated women of about 4 percentage points between 1973 and 1993. There was no significant change in the short-term job probability for highly educated men over this period.

Tables 5.8 and 5.9 contain estimates of logit models of the employment-based long-term employment probabilities (job durations greater than 10 or 20 years).[11] These tables show patterns generally consistent with the results for the short-term job probabilities in table 5.7.[12]

Consider first the estimates for the 10-year probabilities in table 5.8. There is no significant overall trend, but there has been a statistically significant small decrease in this probability for men (about 2.8 percentage points over the 20-year period) and a larger significant increase for women (about 6.5 percentage points over the 20-year period). As before, the change for men is concentrated in the lower educational categories, where there has been a substantial decline in the 10-year probability of about 5 percentage points over the 20-year period. And, aside from the lowest educational category, there has been an even more substantial increase in the 10-year probability for women over time (about 8 percentage points over the 20-year period).

Now consider the estimates for the 20-year probabilities in table 5.9. There is a small significant overall decrease in this probability, which once again, is driven by a decrease in the probability for males and partially offset by an increase in the probability of long-term employment for females. The increase for females (about 3 percentage points over the 20-year period) is particularly noteworthy given the fact that the sample for this analysis consists of women from less recent cohorts.

The breakdown by educational category in the 20-year probabilities is as before. The least educated men have 20-year probabilities that have declined substantially between 1973 and 1993 (by about 8 percentage points). The 20-year probabilities for highly educated women increased over the same period (by about 5 percentage points).[13]

5.4.2 Population-Based Probabilities

Appendix tables 5A.8, 5A.9, and 5A.10 contain population-based sample fractions in the various duration categories broken down by age, sex, and year. The short-term job fractions in table 5A.8 show a substantial (though nonmonotonic) increase over time for men, particularly in the older age categories.[14]

11. Recall that the sample for the 10-year probability is restricted to workers aged 35–64 and that the sample for the 20-year probability is restricted to workers aged 45–64.

12. Of course, it does not have to be the case that movements in the probability that jobs last less than one year will be reflected in concomitant movements in the probabilities of long-term job durations.

13. The latter percentage change is computed using a $p(1 - p)$ value of 0.11 rather than the 0.2 applied to all earlier estimates. This lower value is used because the fraction of females who report job durations of more than 20 years is much smaller. See table 5A.7.

14. At least part of this reflects earlier retirement behavior by men.

Table 5.8 **Logit Analysis of Probability of Job Duration More Than 10 Years for Employed Individuals Aged 35–64 (year by education interaction)**

	All		Males		Females	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant	.383	.364	1.15	1.48	−1.35	−1.39
	(.0611)	(.0967)	(.0792)	(.132)	(.101)	(.149)
Ed < 12	−.178	.658	−.303	−.0159	−.240	.930
	(.0125)	(.162)	(.0162)	(.209)	(.0212)	(.276)
12 < Ed < 16	−.0570	−.642	−.136	−1.06	−.0721	−.984
	(.0127)	(.167)	(.0169)	(.219)	(.0201)	(.275)
Ed ≥ 16	.111	−.0487	−.133	−.972	.237	.271
	(.0120)	(.159)	(.0154)	(.203)	(.0201)	(.275)
Year	−.0012		−.0069		.0161	
	(.0007)		(.0009)		(.0012)	
(Ed < 12)*Year		−.0112		−.0144		.0023
		(.0016)		(.0020)		(.0028)
(Ed = 12)*Year		.0001		−.0107		.0166
		(.0011)		(.0016)		(.0018)
(12 < Ed < 16)*Year		.0059		.00002		.0272
		(.0016)		(.0021)		(.0027)
(Ed ≥ 16) * Year		.0009		−.0008		.0162
		(.0015)		(.0018)		(.0027)
p-Value equality of time trends		<.0001		<.0001		<.0001
No. of observations	218,491	218,491	125,300	125,300	93,191	93,191
log L	−141,041.5	−141,011.1	−82,990.6	−82,969.0	−54,383.7	−54,363.2

Notes: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling one if job duration is more than 10 years. All models include controls for education (three dummy variables for four categories) and age (five dummy variables for six categories). The analysis is weighted using CPS sampling weights. The included age range is 35–64.

Table 5.9 **Logit Analysis of Probability of Job Duration More than 20 Years for Employed Individuals Aged 45–64 (year by education interacton)**

	All		Males		Females	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant	−.0733	−.0732	.407	.551	−1.96	−1.92
	(.0900)	(.132)	(.109)	(.178)	(.177)	(.263)
Ed < 12	−.144	.830	−.312	.456	−.213	1.17
	(.0176)	(.231)	(.0211)	(.277)	(.0360)	(.472)
12 < Ed < 16	−.0558	−.819	−.143	−1.45	−.0891	−.701
	(.0199)	(.258)	(.0243)	(.314)	(.0379)	(.507)
Ed ≥ 16	.103	−.345	−.194	−.682	.296	−.794
	(.0185)	(.242)	(.0221)	(.290)	(.0365)	(.490)
Year	−.0079		−.0082		.0074	
	(.0011)		(.0013)		(.0021)	
(Ed < 12)*Year		−.0199		−.0195		−.0099
		(.0022)		(.0026)		(.0048)
(Ed = 12)*Year		−.0079		−.0099		.0070
		(.0017)		(.0021)		(.0031)
(12 < Ed < 16)*Year		.0011		.0056		.0142
		(.0025)		(.0031)		(.0051)
(Ed ≥ 16)*Year		−.0026		−.0041		.0197
		(.0026)		(.0027)		(.0048)
p-Value equality of time trends		<.0001		<.0001		<.0001
No. of observations	122,849	122,849	71,409	71,409	51,440	51,440
log L	−66,675.9	−66,652.6	−43,954.4	−43,933.4	−19,432.2	−19,421.4

Notes: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling one if job duration is more than 20 years. All models include controls for education (three dummy variables for four categories) and age (three dummy variables for four categories). The analysis is weighted using CPS sampling weights. The included age range is 45–64.

The short-term job fractions for women show a dramatic decline over time, reflecting women's increased employment rates. The long-term job fractions in tables 5A.9 and 5A.10 show analogous patterns.[15] There is an aggregate increase in the 10-year probability for all but the oldest age category, but this is not reflected in the 20-year probability. Both the 10- and 20-year probabilities have declined somewhat for men. This is in contrast to the quite dramatic increase in 10-year probabilities for women, although this is somewhat weaker among women 55–64 years old. There has also been a substantial increase in the 20-year probability for women 45–54 years old, with most of this coming in the past few years. There is no strong trend apparent in the 20-year probability for women 55–64 years old.

Tables 5.10, 5.11, and 5.12 contain estimates of logit models of the population-based probabilities analogous to the employment-based estimates in tables 5.7, 5.8, and 5.9. As before, this analysis provides summary measures of time trends and examines variation in these trends across educational categories. The structure of these tables is the same as in tables 5.7, 5.8, and 5.9. They also include the same control variables.

Table 5.10 contains estimates of logit models of the population-based probability that a worker has been on his or her job no more than one year. When no distinction is made by sex, there is a slight but significant downward trend in the short-term job probability. This small aggregate figure masks large opposing movements of approximately equal magnitudes for males and females (about 8 percentage points each over this period).[16] Once again, separate time trends by educational category allow a much sharper picture to emerge.[17]

The specific results suggest that the overall increase in the probability of short durations is due entirely to the lowest educational category. The probability of a worker with less than a high school education being in a short-term job is predicted to be about 7 percentage points higher in 1993 than in 1973. The estimates show that the time trends in the three higher educational categories were significantly negative, suggesting a lower short-term job probability over time.

Examining the trends separately for men and women suggests that low-education results are driven by large increases in the short-term job probabilities for men in the two lowest educational categories. Men with less than a high school education have a probability of being in a short-term job that is predicted to be fully 16 percentage points higher in 1993 than in 1973. The

15. Remember that the 25–34 age column in table 5A.9 is not particularly relevant because many workers that young have not had time to accumulate much job tenure. Neither the 25–34 nor the 35–44 age columns in table 5A.10 are very interesting for the same reason.

16. The calculations of changes in probabilities over the 20-year period in this subsection are again calculated using a $p(1 - p)$ value of 0.2. While this is not far off on average. the same caution noted above applies. The specific percentage changes mentioned in the text are. of necessity. approximations.

17. As with the employment-based probabilities. the hypothesis that the time trends are the same across educational categories can be rejected in all cases.

Table 5.10 **Logit Analysis of Probability of Job Duration One Year or Less for All Individuals Aged 21–64 (year by education interaction)**

	All		Males		Females	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant	.487	.713	−2.00	−2.45	2.48	2.50
	(.0386)	(.0605)	(.0598)	(.101)	(.0539)	(.128)
Ed < 12	.535	−1.37	.611	−.469	.682	−1.08
	(.0077)	(.0996)	(.0119)	(.154)	(.0110)	(.144)
12 < Ed < 16	−.104	.567	.0640	1.74	−.126	.907
	(.0075)	(.0991)	(.0117)	(.156)	(.0102)	(.136)
Ed ≥ 16	−.450	−.253	−.203	1.49	−.438	.0013
	(.0080)	(.107)	(.0122)	(.163)	(.0112)	(.150)
Year	−.0023		.0206		−.0210	
	(.0004)		(.0007)		(.0006)	
(Ed < 12)*Year		.0182		.0393		.0001
		(.0010)		(.0014)		(.0015)
(Ed = 12)*Year		−.0050		.0261		−.0213
		(.0007)		(.0012)		(.0009)
(12 < Ed < 16)*Year		−.0128		.0062		−.0335
		(.0009)		(.0014)		(.0013)
(Ed ≥ 16)*Year		−.0073		.0060		−.0265
		(.0010)		(.0015)		(.0015)
p-Value equality of time trends		<.0001		<.0001		<.0001
No. of observations	550,552	550,552	260,129	260,129	290,423	290,423
log L	−362,625.5	−362,320.8	−156,831.9	−156,637.3	−189,164.4	−189,006.7

Notes: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling one if job duration is less than or equal to one year. All models include controls for education (three dummy variables for four categories) and age (eight dummy variables for nine categories). The analysis is weighted using CPS sampling weights. Not-employed workers are classified as having job duration less than one year. The included age range is 21–64.

change is somewhat smaller but still quite substantial for men with exactly a high school education (an increase of 10 percentage points). That these changes are larger than the employment-based changes reflects declines in employment rates over the 1973–93 period for less educated men.

The decrease in short-term job probabilities at higher educational levels is the result of substantial declines in these probabilities for women (a decline of 10 to 12 percentage points between 1973 and 1993). Once again, these changes are larger than those found on an employment basis, and this reflects the increased employment rates of women over the sample period.

Tables 5.11 and 5.12 contain estimates of logit models of the population-based long-term employment probabilities (job durations greater than 10 years and greater than 20 years). These tables show patterns generally consistent with the results for the short-term job probability in table 5.10.

There is a very small decrease in the both aggregate long-term job probabilities over the 1973–93 period (less than 1 percentage point overall). But, as with the short-term job probability, this apparent aggregate stability masks roughly offsetting changes for males and females of about 8 to 10 percentage points over the period. Declines in long-term job probabilities for males were offset by approximately equal increases for females. As before, the decline for men is concentrated in the lowest educational categories, where there has been a substantial decline in both long-term job probabilities of about 8 to 12 percentage points over the 20-year period. For females outside the lowest educational category, there has been an even more substantial increase in both long-term job probabilities over time (ranging from 10 to 16 percentage points for the 10-year probability and somewhat less for the 20-year probability).

Overall, the population-based estimates show the same general patterns as the employment-based estimates. The same patterns exist in both series, though they are generally more substantial in the population-based numbers. This is largely due to the fact that changes in employment rates (both supply and demand induced) that are central to the population-based numbers reinforced the changes apparent in the employment-based numbers.

5.5 Concluding Remarks

The results of my analysis are clear and consistent using several measures of job duration. Simply put, no evidence presented here supports the popular view that long-term jobs are becoming less common in the United States. It is true that long-term jobs are now allocated somewhat differently across the population than they were 20 years ago. Long-term jobs have become more scarce for the least educated (particularly men). This is consistent with other evidence that the economic position of the least educated workers has deteriorated in the past 15 to 20 years (Katz and Murphy 1992). It is worth investigating how much of this deterioration is related to job instability.

Long-term jobs used to be almost exclusively the province of men. The

Table 5.11 **Logit Analysis of Probability of Job Duration More Than 10 Years for All Individuals Aged 35–64 (year by education interaction)**

	All		Males		Females	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant	−.946	−1.20	.979	1.27	−3.56	−3.66
	(.0542)	(.0856)	(.0726)	(.121)	(.0897)	(.133)
Ed < 12	−.390	1.43	−.498	.455	−.564	1.48
	(.0108)	(.141)	(.0144)	(.188)	(.0184)	(.242)
12 < Ed < 16	.0638	−.612	−.0730	−1.25	.0469	−1.19
	(.0114)	(.150)	(.0156)	(.205)	(.0181)	(.247)
Ed > 16	.388	.319	−.0548	−1.24	.454	.187
	(.0109)	(.145)	(.0145)	(.191)	(.0181)	(.247)
Year	−.0013		−.0173		.0240	
	(.0006)		(.0008)		(.0010)	
(Ed < 12)*Year		−.0204		−.0326		.0005
		(.0014)		(.0018)		(.0025)
(Ed = 12)*Year		.0018		−.0208		.0252
		(.0010)		(.0014)		(.0016)
(12 < Ed < 16)*Year		.0097		−.0069		.0396
		(.0015)		(.0019)		(.0024)
(Ed ≥ 16)*Year		.0025		−.0055		.0283
		(.0014)		(.0017)		(.0024)
p-Value equality of time trends		<.0001		<.0001		<.0001
No. of observations	324,121	324,121	152,987	152,987	171,134	171,134
log *L*	−185,951.4	−185,817.8	−99,287.6	−99,209.7	−75,469.0	−75,400.1

Notes: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling one if job duration is more than 10 years. All models include controls for education (three dummy variables for four categories) and age (five dummy variables for six categories). The analysis is weighted using CPS sampling weights. Not-employed individuals are classified as having job duration less than one year. The included age range is 35–64.

Table 5.12 **Logit Analysis of Probability of Job Duration More Than 20 Years for All Individuals Aged 45–64 (year by education interaction)**

	All		Males		Females	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant	−.955	−1.13	.488	.614	−3.71	−3.72
	(.0845)	(.133)	(.103)	(.169)	(.169)	(.251)
Ed < 12	−.351	1.47	−.494	.833	−.539	1.64
	(.0162)	(.216)	(.0197)	(.261)	(.0338)	(.449)
12 < Ed < 16	.0468	−.894	−.100	−1.62	.0195	−1.06
	(.0186)	(.244)	(.0230)	(.300)	(.0360)	(.485)
Ed ≥ 16	.376	−.114	−.0193	−.990	.527	−.761
	(.0174)	(.229)	(.0212)	(.278)	(.0345)	(.467)
Year	−.0097		−.0185		.0139	
	(.0010)		(.0012)		(.0020)	
(Ed < 12)*Year		−.0300		−.0365		−.0126
		(.0021)		(.0025)		(.0046)
(Ed = 12)*Year		−.0076		−.0200		.0140
		(.0016)		(.0020)		(.0030)
(12 < Ed < 16)*Year		.0036		−.0020		.0266
		(.0024)		(.0029)		(.0048)
(Ed ≥ 16)*Year		−.0018		−.0085		.0290
		(.0022)		(.0026)		(.0046)
p-Value equality of time trends		<.0001		<.0001		<.0001
No. of observations	197,872	197,872	92,838	92,838	105,034	105,034
log *L*	−83,594.0	−83,523.8	−51,275.6	−51,225.1	−25,048.5	−25,022.4

Notes: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling one if job duration is more than 20 years. All models include controls for education (three dummy variables for four categories) and age (three dummy variables for four categories). The analysis is weighted using CPS sampling weights. Not-employed individuals are classified as not having job duration more than 10 years. The included age range is 45–64.

largest secular change in the data is the dramatically increased probability of long-term employment for women. However, it remains unclear whether these long-term jobs for women are of equal quality to long-term jobs held by men. It is therefore worth investigating how much of the decline in the male-female wage gap in the 1980s is related to increases in job duration (Wellington 1992). In the final analysis, to paraphrase Mark Twain, reports of the death of "the Great American Job" are greatly exaggerated.

Appendix

Table 5A.1 **Median Job Duration by Age, Year, and Sex for Employed Individuals**

	Age Category			
Year	25–34	35–44	45–54	55–64
		Employed Individuals		
1951	2.6	3.2	6.3	8.0
1963	3.0	6.0	9.0	11.8
1966	2.7	6.0	8.8	13.0
1968	2.5	5.2	8.6	12.3
1973	2.8	5.2	8.4	11.4
1978	2.5	4.9	8.3	11.1
1979	2.8	5.4	9.7	12.7
1981	3.1	5.1	9.1	12.1
1983	3.0	5.3	9.7	13.0
1987	3.0	5.6	9.2	12.2
1991	3.0	5.5	9.5	11.9
1993	3.2	5.8	9.5	12.4
		Employed Males		
1951	2.8	4.5	7.6	9.3
1963	3.5	7.6	11.4	14.7
1966	3.2	7.8	11.5	15.8
1968	2.8	6.9	10.2	14.8
1973	3.1	6.5	11.3	14.4
1978	2.8	6.8	11.1	14.6
1979	3.3	7.6	12.5	15.8
1981	3.1	7.1	11.1	15.1
1983	3.3	7.3	12.7	16.4
1987	3.2	7.1	11.8	15.1
1991	3.2	6.8	11.6	15.0
1993	3.5	6.9	11.7	14.0
		Employed Females		
1951	1.8	3.1	4.0	4.5
1963	2.0	3.6	6.1	7.8
1966	1.9	3.5	5.1	9.0
1968	1.6	2.9	5.1	8.7
1973	2.2	3.4	5.7	8.5
1978	2.0	3.3	5.8	8.6
1979	2.2	3.3	6.4	9.6
1981	3.0	4.1	6.1	10.1
1983	2.7	4.1	6.4	9.9
1987	2.6	4.4	6.9	9.9
1991	2.7	4.5	6.8	9.8
1993	3.0	5.0	7.6	10.3

Sources: Statistics for 1951–68 taken from BLS publications and based on supplements to the Current Population Survey in January of the relevant year (Bureau of the Census 1951; Department of Labor 1963, 1967, 1969). Statistics for 1973–93 based on author's calculations of weighted interpolated medians using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993.

Table 5A.2 **.9 Quantile Job Duration by Age, Year, and Sex for Employed Individuals**

	Age Category			
Year	25–34	35–44	45–54	55–64
		Employed Individuals		
1973	8.6	17.1	25.3	32.0
1978	8.7	16.4	25.6	31.5
1979	9.3	16.4	26.7	32.5
1981	9.1	16.1	26.1	33.1
1983	9.5	16.6	25.7	33.3
1987	9.7	17.0	25.2	32.8
1991	10.1	17.7	25.1	32.0
1993	9.7	17.5	25.2	31.5
		Employed Males		
1973	9.0	18.0	26.4	34.9
1978	9.4	17.8	27.4	32.9
1979	9.7	17.8	28.0	34.3
1981	10.1	18.1	28.0	35.1
1983	9.8	17.9	27.6	35.0
1987	10.0	18.1	27.0	35.0
1991	10.3	18.4	26.6	34.6
1993	10.1	18.3	26.8	34.5
		Employed Females		
1973	7.5	13.8	19.9	25.5
1978	7.8	12.4	19.0	25.5
1979	8.6	13.4	20.4	26.3
1981	9.0	14.1	20.1	26.1
1983	8.8	14.4	19.7	26.2
1987	9.1	14.9	19.8	25.4
1991	9.7	16.2	20.8	26.8
1993	9.1	16.1	22.8	25.8

Sources: Statistics for 1973–93 based on author's calculations of weighted interpolated quantiles using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993.

Table 5A.3 **Median Job Duration by Age, Year, and Sex for All Individuals**

	Age Category			
Year	25–34	35–44	45–54	55–64
		All Individuals		
1973	1.0	2.3	3.7	1.7
1978	1.0	2.4	3.7	0.7
1979	1.4	2.8	4.2	1.1
1981	2.0	3.1	4.1	0.3
1983	1.5	2.9	3.8	0.0
1987	1.7	3.4	4.4	0.2
1991	1.8	3.6	5.0	0.7
1993	2.1	3.8	5.2	1.3
		All Males		
1973	2.7	5.8	9.6	7.9
1978	2.2	5.9	9.0	6.1
1979	2.8	6.6	10.4	8.0
1981	3.1	6.1	9.1	6.1
1983	2.3	5.3	9.6	4.7
1987	2.5	5.7	8.6	4.1
1991	2.5	5.4	8.7	3.6
1993	2.8	5.4	8.2	4.6
		All Females		
1973	0.0	0.05	0.04	0.0
1978	0.2	0.4	0.4	0.0
1979	0.5	0.8	0.5	0.0
1981	0.4	0.7	0.8	0.0
1983	0.5	1.2	0.7	0.0
1987	0.9	1.8	1.8	0.0
1991	1.1	2.3	2.8	0.0
1993	1.4	2.5	3.4	0.0

Sources: Statistics for 1973–93 based on author's calculations of weighted interpolated quantiles using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993. Individuals who are not employed are counted as having zero job duration.

Table 5A.4 **.9 Quantile Job Duration by Age, Year, and Sex for All Individuals**

	Age Category			
Year	25–34	35–44	45–54	55–64
		All Individuals		
1973	7.4	15.5	22.6	26.9
1978	7.6	14.8	23.2	27.3
1979	8.3	15.2	24.8	27.9
1981	8.1	15.1	24.1	27.1
1983	8.1	15.3	23.5	27.6
1987	8.6	15.8	23.1	26.2
1991	9.3	16.7	23.8	25.9
1993	8.7	16.3	24.0	26.5
		All Males		
1973	8.8	17.7	26.0	32.2
1978	8.9	17.5	26.8	31.3
1979	9.4	17.4	27.5	32.3
1981	10.0	17.1	27.0	32.1
1983	9.2	17.2	26.8	32.4
1987	9.6	17.8	26.0	32.0
1991	10.0	18.0	25.7	30.7
1993	9.7	17.7	26.2	30.5
		All Females		
1973	5.2	9.2	14.0	16.7
1978	5.8	9.3	14.0	16.2
1979	6.5	10.7	15.3	17.5
1981	7.1	11.1	16.1	17.0
1983	7.0	11.7	15.2	16.6
1987	7.6	13.1	16.6	17.6
1991	7.8	14.3	18.9	19.7
1993	7.5	14.7	20.1	19.8

Sources: Statistics for 1973–93 based on author's calculations of weighted interpolated quantiles using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993. Individuals who are not employed are counted as having zero job duration.

Table 5A.5 **Fraction with Job Duration of One Year or Less for Employed Individuals**

Year	Age Category 25–34	35–44	45–54	55–64
	Employed Individuals			
1973	.277	.169	.112	.080
1978	.311	.203	.136	.106
1979	.345	.226	.143	.105
1981	.300	.200	.135	.101
1983	.300	.200	.130	.097
1987	.309	.206	.147	.106
1991	.303	.196	.145	.113
1993	.280	.182	.133	.100
	Employed Males			
1973	.249	.137	.097	.070
1978	.283	.166	.110	.095
1979	.309	.173	.113	.089
1981	.267	.172	.111	.094
1983	.276	.168	.112	.089
1987	.282	.174	.127	.096
1991	.280	.167	.129	.106
1993	.268	.161	.130	.099
	Employed Females			
1973	.328	.223	.137	.096
1978	.351	.259	.176	.123
1979	.398	.301	.190	.131
1981	.345	.237	.167	.112
1983	.331	.242	.155	.108
1987	.343	.245	.172	.121
1991	.331	.229	.164	.122
1993	.294	.207	.137	.100

Sources: Statistics for 1973–93 based on author's weighted counts using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993.

Table 5A.6 Fraction with Job Duration of More Than 10 Years for Employed Individuals

	Age Category			
Year	25–34	35–44	45–54	55–64
	Employed Individuals			
1973	.066	.288	.451	.546
1978	.063	.274	.443	.535
1979	.057	.284	.465	.561
1981	.076	.286	.453	.566
1983	.059	.283	.459	.562
1987	.066	.282	.438	.536
1991	.083	.297	.446	.531
1993	.074	.300	.456	.538
	Employed Males			
1973	.075	.356	.537	.603
1978	.076	.356	.532	.602
1979	.066	.363	.558	.629
1981	.090	.364	.541	.625
1983	.066	.360	.556	.637
1987	.075	.345	.523	.590
1991	.094	.346	.526	.584
1993	.084	.341	.519	.574
	Employed Females			
1973	.050	.173	.310	.451
1978	.042	.153	.307	.430
1979	.043	.171	.318	.451
1981	.057	.181	.331	.476
1983	.050	.183	.325	.458
1987	.055	.205	.329	.460
1991	.070	.239	.352	.462
1993	.062	.252	.384	.491

Sources: Statistics for 1973–93 based on author's weighted counts using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993.

Table 5A.7 **Fraction with Job Duration of More Than 20 Years for Employed Individuals**

Year	Age Category 25–34	35–44	45–54	55–64
		Employed Individuals		
1973	.001	.050	.213	.309
1978	.000	.042	.209	.314
1979	.001	.032	.218	.323
1981	.000	.043	.198	.311
1983	.000	.030	.194	.307
1987	.000	.027	.179	.282
1991	.000	.038	.193	.292
1993	.000	.036	.206	.287
		Employed Males		
1973	.001	.060	.283	.388
1978	.000	.057	.288	.398
1979	.001	.043	.296	.410
1981	.001	.058	.271	.394
1983	.001	.041	.279	.403
1987	.000	.039	.256	.365
1991	.000	.047	.268	.367
1993	.000	.041	.271	.360
		Employed Females		
1973	.000	.033	.097	.177
1978	.000	.021	.090	.183
1979	.000	.018	.097	.181
1981	.000	.022	.096	.183
1983	.000	.016	.078	.172
1987	.000	.013	.081	.164
1991	.000	.028	.106	.194
1993	.000	.030	.132	.191

Sources: Statistics for 1973–93 based on author's weighted counts using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993.

Table 5A.8 **Fraction with Job Duration of One Year or Less for All Individuals**

	Age Category			
Year	25–34	35–44	45–54	55–64
		All Individuals		
1973	.511	.409	.382	.473
1978	.502	.407	.390	.516
1979	.509	.412	.393	.511
1981	.489	.399	.380	.532
1983	.505	.408	.404	.549
1987	.478	.371	.374	.546
1991	.463	.352	.346	.530
1993	.441	.346	.336	.506
		All Males		
1973	.317	.194	.186	.291
1978	.361	.227	.213	.367
1979	.365	.235	.211	.357
1981	.364	.254	.231	.392
1983	.411	.283	.263	.426
1987	.381	.261	.254	.423
1991	.379	.259	.260	.429
1993	.364	.262	.266	.415
		All Females		
1973	.691	.607	.563	.635
1978	.635	.574	.554	.649
1979	.645	.578	.566	.654
1981	.606	.533	.518	.657
1983	.594	.525	.533	.655
1987	.570	.475	.485	.654
1991	.544	.441	.427	.620
1993	.515	.427	.401	.590

Sources: Statistics for 1973–93 based on author's weighted counts using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993. Individuals who are not employed are counted as having zero duration.

Table 5A.9 **Fraction with Job Duration of More Than 10 Years for All Individuals**

	Age Category			
Year	25–34	35–44	45–54	55–64
		All Individuals		
1973	.045	.205	.313	.312
1978	.045	.204	.313	.289
1979	.042	.216	.329	.306
1981	.056	.215	.325	.294
1983	.042	.209	.314	.281
1987	.050	.223	.321	.272
1991	.064	.239	.341	.282
1993	.058	.240	.350	.294
		All Males		
1973	.068	.332	.484	.460
1978	.068	.330	.471	.421
1979	.060	.336	.497	.444
1981	.078	.328	.469	.419
1983	.054	.310	.462	.401
1987	.064	.309	.447	.376
1991	.081	.308	.448	.373
1993	.073	.300	.438	.373
		All Females		
1973	.023	.088	.157	.182
1978	.024	.088	.166	.172
1979	.025	.103	.170	.180
1981	.035	.111	.191	.184
1983	.031	.115	.179	.177
1987	.036	.142	.204	.181
1991	.047	.173	.241	.200
1993	.043	.182	.267	.224

Sources: Statistics for 1973–93 based on author's weighted counts using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993. Individuals who are not employed are counted as having zero job duration.

Table 5A.10 Fraction with Job Duration of More Than 20 Years for All Individuals

	Age Category			
Year	25–34	35–44	45–54	55–64
		All Individuals		
1973	.000	.035	.148	.177
1978	.000	.032	.148	.170
1979	.000	.025	.154	.176
1981	.000	.032	.142	.162
1983	.000	.022	.133	.153
1987	.000	.021	.132	.143
1991	.000	.031	.148	.155
1993	.000	.029	.158	.157
		All Males		
1973	.000	.056	.256	.296
1978	.000	.053	.254	.278
1979	.000	.039	.263	.290
1981	.000	.052	.234	.265
1983	.000	.035	.231	.254
1987	.000	.035	.220	.233
1991	.000	.042	.228	.234
1993	.000	.036	.228	.234
		All Females		
1973	.000	.017	.049	.072
1978	.000	.012	.048	.073
1979	.000	.011	.052	.072
1981	.000	.013	.055	.071
1983	.000	.010	.043	.067
1987	.000	.008	.050	.065
1991	.000	.020	.073	.084
1993	.000	.022	.092	.087

Sources: Statistics for 1973–93 based on author's weighted counts using data from supplements to the Current Population Survey in January 1973, 1978, 1981, 1983, 1987, and 1991; in May 1979; and in April 1993. Individuals who are not employed are counted as having zero job duration.

References

Abraham, Katharine G., and James L. Medoff. 1984. Length of service and layoffs in union and nonunion work groups. *Industrial and Labor Relations Review* 38 (October): 87–97.

Chamberlain, Gary. 1994. Quantile regression, censoring, and the structure of wages. In *Proceedings of the Sixth World Congress of the Econometric Society,* ed. Christopher Sims. New York: Cambridge University Press.

Diebold, Francis X., David Neumark, and Daniel Polsky. 1994. Job stability in the United States. NBER Working Paper no. 4859. Cambridge, Mass.: National Bureau of Economic Research, September.

Farber, Henry S. 1993. The incidence and costs of job loss: 1982–91. *Brookings Papers on Economic Activity: Microeconomics,* 73–119.

———. 1997. The changing face of job loss in the United States, 1981–1995. *Brookings Papers on Economic Activity: Microeconomics,* 55–128.

Hall, Robert E. 1982. The importance of lifetime jobs in the U.S. economy. *American Economic Review* 72 (September): 716–24.

Katz, Lawrence F., and Kevin M. Murphy. 1992. Changes in relative wages, 1963–1987: Supply and demand factors. *Quarterly Journal of Economics* 106 (February): 35–78.

Swinnerton, Kenneth, and Howard Wial. 1995. Is job stability declining in the U.S. economy? *Industrial and Labor Relations Review* 48 (January): 293–304.

Ureta, Manuelita. 1992. The importance of lifetime jobs in the U.S. economy, revisited. *American Economic Review* 82 (March): 322–35.

U.S. Bureau of the Census. 1951. *Current population reports: Labor force,* Series P-50, no. 36. Washington, D.C.: U.S. Bureau of the Census, 5 November.

U.S. Department of Labor. Bureau of Labor Statistics. 1963. *Job tenure of American workers.* Special Labor Force Report no. 36. Washington, D.C.: Government Printing Office.

———. 1967. *Job tenure of workers.* Special Labor Force Report no. 77. Washington, D.C.: Government Printing Office.

———. 1969. *Job tenure of workers.* Special Labor Force Report no. 112. Washington, D.C.: Government Printing Office.

Wellington, Alison J. 1992. Changes in the male/female wage gap, 1976–1985. *Journal of Human Resources* 28:385–411.

Comment Derek Neal

In the introduction to this paper, the author correctly notes that recent reports by media and government either state or imply that the typical worker in the United States has recently experienced a significant loss of job security. In the conclusion, the author argues that these reports have likely overstated their

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case. In between, he uses data from the Current Population Survey to provide a careful and thorough description of changes in the distribution of existing job tenure over the period 1973–93.[1]

I want to commend the author for providing a great deal of information that speaks to an important and timely question. Further, I am inclined to agree generally with his conclusions. However, I would like to raise a few issues that I feel the author should have explored further.

My concerns arise from the fact that Current Population Survey data on job tenure do not speak directly to the issue of job security. Tenure data do not provide direct evidence about separation rates, and I will argue later that even with good information about separation rates, we cannot not make clear inferences about job security.

The results in tables 5.1 and 5.4 demonstrate that trends in median tenure among men are quite different depending on whether the estimates are employment based or population based. Both sets of analyses show that median job tenure has declined among less educated men. However, the magnitude of the decline is much greater in the population-based results. The author motivates the presentation of the population-based results by arguing that the employment-based results may be contaminated by business cycle effects because those with the least tenure are laid off during recessions. However, it is possible that male workers with little education spend more time "between jobs" than they did 20 years ago. This would explain the observed pattern of results, and it might occur either because separation rates are now higher among this group or because exit rates from unemployment are lower or both.

Without direct evidence concerning separation rates, it is hard to make strong inferences about secular changes in job security. Further, even if future studies do document how separation rates have changed or not changed within various groups, the implications for changes in job security will not be transparent. Workers leave employment matches either because they receive bad information about their current match or because they receive good information about potential alternatives. When press accounts describe workers as concerned about their job security, I interpret this as a statement that workers are worried about future separations that might arise from sudden negative changes in the expected value of their existing matches. Workers rarely lose sleep over the prospects of leaving their current jobs for better ones.

In recent work on displacement, the author notes that, within several groups, the probability of displacement by layoff or plant closing has changed substantially since the early 1980s (Farber 1993). However, the author also notes that even displacement data give an incomplete picture of job security. Workers may voluntarily leave firms that suffer adverse shocks because shocks cause them to update their forecasts of future wages. Such separations could be

1. The author not only examines various conditional quantiles of the tenure distribution, he also examines the cumulative distribution function at 1 year, 10 years, and 20 years of tenure.

traced to exogenous declines in the value of specific employment matches, but they would not appear in the data as displacements.

In short, it may be quite difficult to document trends in job security. We all have a sense of what we mean when we use the term, but we do not have a precise definition that lends itself directly to empirical measurement. In this context, it is interesting to note that both the employment-based and population-based analyses show that median job tenure has risen substantially among women. Should we interpret this as evidence that job security among women has increased over the past two decades, or could the trend in observed job tenure be driven entirely by the increased commitment of women to the labor force? Women are now less likely to leave their jobs when their children are young. This implies that in both the workforce and the population as a whole we should see an increase in median job tenure among women. We might expect that this increased attachment to market work should also increase the value of job-specific matches between women and their employers, thus making women more secure in their jobs. However, I know of no direct evidence that this is the case.

Further, changes in retirement behavior over the past 20 years also raise questions about the interpretation of changes in the distribution of job tenure. I noted earlier that, among men, population-based measures imply larger declines in median tenure than do employment-based measures. A comparison of tables 5A.1 and 5A.3 shows that the largest differences between employment- and population-based estimates of the secular changes in median job tenure come from the analyses of older men. The difference is particularly striking among men aged 55–64.

Are older men simply consuming more leisure, or are they spending more time searching for employment? Hurd (1990) reports that retirement ages have fallen significantly over the past several decades. On the other hand, the author's own work shows that, between the recessions of the early 1980s and early 1990s, displacement became more common among older men (Farber 1993).

It is likely that older workers are retiring earlier primarily because they are wealthier than previous cohorts. However, if workers become less secure in their current employment, they may become more willing to choose early retirement. Workers who view future displacement as a likely outcome may be quite willing to accept early retirement plans and then go back to work if a good opportunity comes along.

In general, it would be interesting to expand the analyses in this paper by estimating each specification separately by age group. Of particular interest is whether the decline in median tenure among less educated men is being driven by the employment patterns of the young, the old, or both. If the decline is being driven by older workers, the issue of retirement decisions becomes crucial. If older males are retiring earlier or shifting from full-time to part-time employment simply because they are wealthier than previous cohorts, the author's results may actually overstate losses of job security among less educated males.[2]

I want to end as I began by stating that this paper is basically a success. Against a backdrop of considerable discussion among both policymakers and the media about the need to address the drastic loss of job security suffered by American workers, the author presents a thorough documentation of recent changes in the distribution of job tenure. He correctly argues that since the overall distribution of tenure has been relatively stable for the past two decades, it is hard to claim that long-term jobs are becoming less common in the United States, and he clearly places the burden of proof on those who contend that declining job security is a pervasive problem.

However, I feel the paper would have been even more interesting if the author had devoted a portion of his efforts to the tasks of defining job security and discussing how one might measure it directly. Popular discussions of job security usually proceed without a clear definition of the term. Although job security is a common topic in policy debates, economists have not thought carefully about how to define it or how to measure it. These problems remain for future research.

References

Farber, Henry S. 1993. The incidence and costs of job loss: 1982–91. *Brookings Papers on Economic Activity: Microeconomics,* 73–119.

Hurd, Michael D. 1990. Research on the elderly: Economic status, retirement, and consumption, and saving. *Journal of Economic Literature* 28:565–637.

2. Furthermore, the interaction between age and educational level raises an important measurement issue. The author uses schooling as a proxy for worker skill, but it is not clear that the relationship between schooling and skill is the same across cohorts. If schooling understates the relative skill of older workers and if the decline in median tenure among the less educated has been particularly dramatic among the old, we may not want to think of the decline as primarily affecting unskilled workers. I thank Bob Topel for raising this point during the discussion.