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ARE LIFETIME JOBS DISAPPEARING? JOB DURATION IN THE UNITED STATES: 1973-1993

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ABSTRACT

The public believes that job security has deteriorated dramatically in the United States. In this study, I examine job durations from eight supplements to the Current Population Survey (CPS) administered between 1973 and 1993 in order to determine if, in fact, there has been a systematic change in the likelihood of long-term employment. In order to measure changes in the distribution of job durations, I examine changes in selected quantiles (the median and the 0.9 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 employer 1) less than one year, 2) more than ten years, and 3) more than twenty years.

The central findings are clear. By the measures I examine, there has been no systematic change in the overall distribution of job duration over the last two decades, but the distribution of long-term jobs across the population has changed in two ways. First, individuals, particularly men, with little education (less than twelve years) are substantially less likely to be in long jobs today than they were twenty years ago. Second, women with at least a high-school education are substantially more likely to be in long jobs today than they were twenty years ago.

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I. 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*, November 22, 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.

These reflect a 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 subcontracters, and on part-time workers. Potential motiviation 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 worker. Here, I examine the evidence on job durations in order to determine if, in fact, a systematic change in the likelihood of long-term employment has occurred.

The analysis in this paper is based on evidence regarding the duration of jobs in progress from supplements to the Current Population Survey (CPS) 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 0.9 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 employer 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 the next section.

The central findings, presented in Sections III and IV, are clear. No systematic change has occurred in various measures of the overall distribution of job duration over the last 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 part of the mechanism of this decline. Second, women with at least a high-school education are substantially more likely to be in long 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 1980's (Wellington, 1992).

II. DATA AND MEASUREMENT ISSUES

A. The CPS Data on Job Duration

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

Others have used these data to analyze job durations. 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 five years has a substantial probability of lasting twenty years. He also finds that a substantial fraction of workers will be on a "life-time" 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.

Two recent papers have use these data to examine changes in employment stability using data from the mobility and pension supplements to the CPS. Swinnerton and Wial (1995), using data from 1979 through 1991, analyze job retention rates

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[1]Only summary tables are available for the 1951, 1963, 1966, and 1968 surveys.

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 data 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.

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.[2] These surveys cover eight years over the twenty-year period from 1973 to 1993. One feature that will distinguish 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 (1951-1981) asked workers what year they started working for their current employer (the early question) . In later January supplements (1983-1991) and in all of the pension and benefit supplements (1979-1993), workers were asked how many years they worked for their current employer (the later question). 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

[2]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.

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.[3] 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 their 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 $D=Min(T_s-T_0,1)$. 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

[3]Ureta (1992) accounts for these spikes explicitly in her estimation procedure. Swinnerton and Wial (1993) work around these spikes in selecting intervals over which to compute retention rates.

distribution of job durations within intervals, then D would overstate D_T by one-half year on average. Now consider the later question which asked how many years workers have been with their current employer. Call this response Y. If workers have been with their employer less than one year, they are asked the number of months they have been with their 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 duration of one year. The interpretation of workers with reported durations of one year or 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 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, the distribution of responses to the 1979 question (later-style) with the 1978 and 1981 questions (early-style) do not show any systematic bias.[4] 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

[4]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 that a temporary increase in the 1979 job durations is masking the bias.

true duration in the interval $D-1 < D_T \leq D$.

B. 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 ten 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 call interpolated quantiles, defined as

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

where θ_τ is the τ^{th} 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 τ^{th} quantile is D_{k+1}, and the τ^{th} interpolated quantile is simply a weighted average of the τ^{th} quantile and the next smaller observed value of job duration. The weight, λ is

$$(2) \qquad \lambda = [\tau - P_k] \Big/ [P_{k+1} - P_k]$$

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 these interpolated quantiles under the assumption that 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.

C. Fractions of Workers in Short and Long Jobs

I also examine the fraction 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 fraction of workers who report having been with their employer 1) no more than one year, 2) more than ten years, and 3) more than twenty years. These points on the distribution give a clear picture of what has happened to the incidence of very short jobs and long 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 discuss above.

D. Employment-Based and Population-Based Distributions of Job Duration

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 counter-cyclical; tight labor markets with 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 wouild 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 these employment-based quantiles, and I call probabilities of having job duration in a particular category (≤1 year, >10 years, >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 a straightforward interpretation. 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 could be zero if less than half of the relevant group is not working. The contrast between the employment-based and the population-based probabilities are interpreted similarly. For example, the employment-based probability of being on a job more than ten years is the fraction of workers who have been on their job more than ten years. In contrast, the population-based probability of being on a job more than ten years is the fraction of all individuals (employed or not) who have been on their job more than ten years.[5]

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[5]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

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 a given group of workers holds.

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 them. 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.

III. CHANGES IN INTERPOLATED QUANTILES, 1973-1993

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 1. This figure contains plots of four weighted (by CPS sampling weights) interpolated quantiles (.25, .5, .75, .9) of the

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probability has a natural interpretation.

employment-based tenure distribution by year broken down by sex and four ten-year age categories.[6] These and succeeding figures do not show sampling errors. Sampling errors for these interpolated quantiles, calculated using the delta method, are generally in the range of 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 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 plot for females show some upward movement over time. The combined plot (no distinction by sex) looks very flat. Analogous plots of population-based quantiles are contained in figure 2. These look much like the employment-based quantiles in figure 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 A1 through A4 contain the raw data underlying the median and 0.9 quantiles for figures 1 and 2. Table A1, which contains employment-based medians,

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[6]The vertical scale of all of these 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.

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 due the fact that most workers had to "restart" after returning from World War II, long term trends using this longer time series are difficult to discern.

Figures 3 through 5 are plots of the four employment-based quantiles broken down by age and education. Figure 3 makes no distinction between sexes. It shows a substantial decline in job duration for workers in the lowest education category (<12 years). Not much change is evident in the overall quantiles in the higher education categories. Figure 4 replicates these plots for males. The substantial changes here are a decline in job durations for the least educated men and some decline for the oldest highly educated men (≥16 years). Figure 5 replicates these plots for females. It is interesting that there does not seem to be much decline in job durations for the least educated women. The plots also suggest that there is a fairly systematic increase in job durations for women in the three higher education 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 6 through 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 durations among men (figure 7). Thus, the well-known deterioration in labor market conditions for poorly-

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[7]The sources for these published tabulations are listed in the references.

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 durations for women in the top three education categories (figure 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 quantiles of the distribution of job tenures, table 1 through table 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 65), four education categories (<12 years, 12 years, >12 and <16 years, and ≥16 years), eight calendar years. I do this separately for three samples (all workers, 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 τ^{th} quantile of observations in cell j would be

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

where $\theta_{\tau j}$ is the τ^{th} 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 (1991) suggests that it may be better to use the cell sizes as weights if it is possibile that

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[8]Chamberlain (1991) developed this technique for estimating quantiles.

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 1 and 2 contains eight dummy variables for the age categories, three dummy variables for the education 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-squareds of these regressions are quite large (over 0.95).

The estimates in the first two columns of table 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 decrease over the twenty-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 2 for the 0.9 quantile of the employment-based distribution of job durations show a similar pattern. There is no significant relationship between year and the 0.9 quantile of job duration when no sex distinction is made, and there is actually a small *increase* on average in the 0.9 quantile for males (about 0.3 years over the twenty-year period). The rate of increase in the 0.9 quantile of job durations for females (about 1.5 years over the twenty year period) is substantially larger than the rate of increase of the women's median.

Important differences in time trends of job durations 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 1 and 2 with the time trend interacted with the four educational categories. Table 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 education suffered a decline in median job duration of over 1/2 year on average over the twenty-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 education and men with exactly 12 years education shared this decline. Among workers with more than a high-school education, job durations increased on average. There was no significant increase in medians for more educated males (> 12 years) on average, but the 0.9 quantile of the job duration distribution did increase significantly for more educated men (about 1/2 year over the twenty 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 twenty-year period range from about 1/2 year to about 1 year. The increase in the 0.9 quantiles for women over this period was even larger, ranging from 1.5 years to over 2 years.

Tables 4 through 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 the those not employed, all of whom are coded as having zero job durations. Effectively, these are censored observations, and any cells for which the particular quantile of the job duration distribution being studied is zero (i.e., is represented by a non-employed individual) contain 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 twenty-year period), shown in column 4 of table 4. There is also a larger increase in the

[9]Chamberlain (1991) 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.

population-based 0.9 quantile for females (about 2.5 years over the twenty-year period), shown in column 6 of table 5. The sources of these substantial trends become clearer with separate year effects by education in table 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 twenty year period. There was no significant change in median job durations for males with more than a high school education. The rate of increase of median job duration for women increases monotonically with education category, rising from zero for women with less than a high school education to an increase of about 1.3 years over the twenty-year period for women with at least 16 years education. The large increase in the 0.9 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.

IV. CHANGES IN PROBABILITIES OF SHORT AND LONG-TERM JOBS, 1973-1993

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 one year, 2) the fraction of job durations greater than ten years, and 3) the fraction of job durations greater than twenty years. Based on the results reported above, it is reasonable to expect that the fraction of short jobs (≤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 jobs (> 10 years and > 20 years) decline among the least educated male workers and has 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.

A. Employment-Based Probabilities

Appendix table A5 through A7 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 increases with age. This is so because it is virtually impossible for very young workers to have been on their job for more then ten or twenty 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 ten 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 twenty years on the sample of workers who are at least 45 years old.

Tables 7 through 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 7 contains estimates of logit models of the employment-based probability that a workers has been on his/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 education 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 twenty-year period the employment-based probability that a job is no more than one year old is predicted to have fallen by about 1.3 percentage points.[10] This aggregate figure masks a larger increase for men over the twenty-year period of about 3 percentage points and

[10]The logit coefficient of 0.0034 must be multiplied by some estimate of p(1-p) in order to compute the derivative of the probability with respect to year. A reasonable mean estimate of p(1-p) is 0.2. Thus, over the twenty-year period the probability that a work was in his/her job for no more than one year is predicted to have increased by about 1.4 percentage points (0.0034x0.2x20x100). 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)) varies, and the percentage point changes mentioned in the text are, of necessity, an approximation.

a small decrease for women over the twenty-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 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 job is in the range of 0.25.

An analysis of the trends separately for men and women suggests that this result is driven by a large increase in the short-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 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-job probability in the higher 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-job probability for highly-educated men over this period.

Tables 8 and 9 contains estimates of logit models of the employment-based long-term employment probabilities (job durations greater than ten or twenty years).[11] These tables show patterns generally consistent with the results for the short-job probability in table 7.[12]

Consider first the estimates for the ten-year probability in table 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 twenty-year period) and a larger significant increase for women (about 6.5 percentage points over the twenty-year period). As before, the change for men is concentrated in the lower educational categories, where there has been a substantial decline in the ten-year probability of about 5 percentage points over the twenty year period. And, aside from the lowest educational category, there has been an even-more-substantial increase in the ten-year probability for women over time (about 8 percentage points over the twenty-year period).

Now consider the estimates for the twenty-year probabilities in table 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 three percentage points over the twenty year period) is particularly noteworthy given the fact that the sample for this analysis consists of women from less recent cohorts.

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[11]Recall that the sample for the ten-year probability is restricted to workers aged 35 through 64 and that the sample for the twenty-year probability is restricted to workers aged 45 through 64.

[12]It does not have to be the case that movements in the probability that jobs are less than one year will be reflected in concomitant movements in the probabilities of long job durations.

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

B. Population-Based Probabilities

Tables A8 through A10 contain population-based sample fractions in the various duration categories broken down by age, sex, and year. The short-job fractions in table A8 show a substantial (though non-monotonic) increase over time for men, particularly in the older age categories.[14] The short-job fractions for women show a dramatic decline over time, reflecting women's increased employment rates. The long-job fractions in table A9 and A10 show analogous patterns.[15] There is an aggregate increase in the ten-year probability for all but the oldest age category, but this is not reflected in the twenty-year probability. Both the ten- and twenty-year probabilities have declined somewhat for men. This is in contrast to the quite dramatic increase in ten-year probabilities for women, although this is somewhat weaker among women 55-64 years old. There has also been a substantial increase in the twenty-year probability for women

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[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 is because the fraction of females who report job durations of more than twenty years is much smaller. See table A7.

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

[15]Remember that the 25-34 age column in table A9 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 columns in table A10 are very interesting for the same reason.

45-54 years old, with most of this coming in the last few years. There is no strong trend apparent in the twenty-year probability for women 55-64 years old.

Tables 10 through 12 contain estimates of logit models of the population-based probabilities analogous to the employment-based estimates in tables 7 through 9. As before, this analysis provides summary measures of time trends and to examines variation in these trends across educational categories. The structure of these tables is the same as tables 7 through 9. They also include the same control variables.

Table 10 contains estimates of logit models of the population-based probability that a workers has been on his/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-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 allows 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 is a short job is predicted to be about 7 percentage points higher in 1993 than in 1973. The estimates show that the time trends

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[16]The calculations of changes in probabilities over the twenty-year period in this sub-section 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 applied. 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.

in the three higher educational categories were significantly negative, suggesting a lower short-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-job probabilities for men in the two lowest education categories. Men with less than a high-school education have a probability of being in a short job that is predicted to be fully 16 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 10 percentage points). That these changes are larger than the employment based changes reflects declines in employment rates over the 1973-1993 period for less-educated men.

The decrease in short-job probabilities at higher education 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 11 and 12 contains estimates of logit models of the population-based long-term employment probabilities (job durations greater than ten years and greater than twenty years). These tables show patterns generally consistent with the results for the short-job probability in table 10.

There is a very small decrease in the both aggregate long-job probabilities over the 1973-1993 period (less than 1 percentage point overall). But, as with the short-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-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-job probabilities of about 8 to 12 percentage points over the twenty year period. For females outside the lowest educational category, there has been an even-more substantial increase in both long-job probabilities over time (ranging from 10 to 16 percentage points for the ten-year probability and somewhat less for the twenty-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.

V. Concluding Remarks

The results of my analysis are clear and consistent using several measures of job duration: Simply put, no evidence presented here supports to popular view that long-term jobs are becoming less common in the United States. It is true is that long-term jobs are now allocated somewhat differently across the population than they were twenty 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 last fifteen to twenty 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 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 1980's is related to increases in job duration (Wellington, 1992).

In the final analysis, to paraphrase Mark Twain (*New York Journal*, June 2, 1897) reports of the death of "the great American Job" are greatly exaggerated.

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Table 1
Median Regression of Job Duration
All Employed Individuals Aged 21-64

Variable	All (1)	All (2)	Males (3)	Males (4)	Females (5)	Females (6)
Constant		.450 (.537)		2.63 (.794)		-2.12 (.460)
Year		.00240 (.00639)		-.0179 (.00941)		.0332 (.0055)
1973	.689 (.156)		1.16 (.230)		.502 (.131)	
1978	.539 (.150)		1.11 (.223)		.342 (.123)	
1979	.731 (.189)		1.55 (.279)		.414 (.157)	
1981	.585 (.146)		1.13 (.219)		.541 (.118)	
1983	.761 (.151)		1.52 (.226)		.639 (.122)	
1987	.633 (.606)		1.06 (.226)		.794 (.122)	
1991	.606 (.154)		.831 (.232)		.808 (.125)	
1993	.829 (.192)		.870 (.289)		1.26 (.155)	
Ed < 12	-.732 (.111)	-.732 (.111)	-1.69 (.159)	-1.70 (.161)	-.852 (.0972)	-1.54 (.266)
12 < Ed < 16	-.230 (.100)	-.229 (.100)	-.863 (.150)	-.861 (.152)	-.211 (.0816)	-.649 (.238)
Ed >= 16	.570 (.0999)	.571 (.0997)	-.621 (.145)	-.614 (.147)	.515 (.0851)	.514 (.0864)
p-value equality of year effects	.646		.0233		<.00005	
p-value year effects equal trend		.548		.0492		.0273
# of Cells	288	288	288	288	288	288
# of Obs	378890	378890	214210	214210	164680	164680
R-squared	.970	.969	.964	.963	.959	.957

Note: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantile for 9 age categories, 4 education categories, 2 sex categories (in columns 3-6), and 8 years. Only observations with non-zero quantiles (employed) are included. All observations are weighted by the cell size.

Table 2
0.9 Quantile Regression of Job Duration
All Employed Individuals Aged 21-64

Variable	All (1)	All (2)	Males (3)	Males (4)	Females (5)	Females (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
# of Cells	288	288	288	288	288	288
# of Obs	378890	378890	214210	214210	164680	164680
R-squared	.995	.995	.996	.996	.974	.972

Note: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantiles for 9 age categories, 4 education categories, 2 sex categories (in columns 3-6), and 8 years. Only observations with non-zero quantiles (employed) are included. All observations are weighted by the cell size.

Table 3
Quantile Regression of Job Duration
Employed Individuals Aged 21-64
(Year by Education Interaction)

Variable	All (1) Median	(2) .9 Quantile	Males (3) Median	(4) .9 Quantile	Females (5) Median	(6) .9 Quantile
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
# of Cells	288	288	288	288	288	288
# of Obs	378890	378890	214210	214210	164680	164680
R-squared	.970	.995	.964	.996	.958	.974

Note: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantile for 9 age categories, 4 education categories, 2 sex categories (in columns 3-6), and 8 years. Only observations with non-zero quantiles (employed) are included. All observations are weighted by the cell size. All specifications include eight dummy variables for age categories.

Table 4
Median Regression of Job Duration
All individuals Aged 21-64

Variable	All (1)	(2)	Males (3)	(4)	Females (5)	(6)
Constant		-1.32 (.851)		7.84 (1.13)		-2.90 (.604)
Year		.0201 (.0102)		-.0819 (.0134)		.0387 (.0073)
1973	.337 (.241)		2.02 (.325)		-.0393 (.179)	
1978	.259 (.234)		1.40 (.314)		.206 (.155)	
1979	.352 (.295)		1.92 (.396)		.174 (.195)	
1981	.174 (.232)		1.13 (.305)		.212 (.151)	
1983	.158 (.239)		.722 (.312)		.223 (.155)	
1987	.409 (.239)		.699 (.317)		.421 (.155)	
1991	.589 (.246)		.522 (.323)		.643 (.158)	
1993	.792 (.302)		.431 (.402)		.829 (.196)	
Ed < 12	-1.28 (.174)	-1.26 (.174)	-2.82 (.225)	-2.81 (.225)	-.721 (.444)	-.719 (.437)
12 < Ed < 16	.264 (.159)	.263 (.159)	-.506 (.214)	-.504 (.215)	.177 (.100)	.176 (.0988)
Ed >= 16	1.76 (.165)	1.76 (.165)	.553 (.214)	.552 (.214)	.811 (.108)	.811 (.107)
p-value equality of year effects	.465		<.00005		<.00005	
p-value year effects equal trend		.605		.269		.624
# of Cells	262	262	282	282	189	189
# of Obs	502600	502600	253860	253860	204050	204050
R-squared	.689	.680	.849	.845	.454	.447

Note: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantiles for 9 age categories, 4 education categories, 2 sex categories (in columns 3-6), and 8 years. Only observations with non-zero quantiles (employed) are included. All observations are weighted by the cell size.

Table 5
0.9 Quantile Regression of Job Duration
All Individuals Aged 21-64

Variable	All (1)	(2)	Males (3)	(4)	Females (5)	(6)
Constant		3.98 (.985)		6.11 (.874)		-7.89 (1.12)
Year		-.0117 (.0117)		-.0298 (.0104)		.126 (.0134)
1973	3.17 (.284)		3.90 (.255)		1.69 (.318)	
1978	3.13 (.274)		3.82 (.246)		1.70 (.309)	
1979	3.31 (.344)		3.74 (.311)		2.18 (.384)	
1981	3.01 (.267)		3.78 (.239)		2.25 (.300)	
1983	2.80 (.273)		3.50 (.245)		2.29 (.308)	
1987	2.77 (.278)		3.53 (.249)		2.91 (.313)	
1991	3.05 (.284)		3.44 (.254)		3.85 (.320)	
1993	3.08 (.352)		3.28 (.315)		3.99 (.397)	
Ed < 12	-2.80 (.188)	-2.80 (.188)	-2.29 (.171)	-2.29 (.170)	-3.71 (.210)	-3.70 (.211)
12 < Ed < 16	-.484 (.188)	-.481 (.187)	-1.11 (.169)	-1.11 (.168)	-.0953 (.211)	-.0946 (.212)
Ed >= 16	-.710 (.195)	-.710 (.195)	-2.45 (.169)	-2.45 (.167)	1.08 (.232)	1.08 (.233)
p-value equality of year effects	.595		.229		<.00005	
p-value year effects equal trend		.605		.972		.173
# of Cells	288	288	288	288	288	288
# of Obs	550940	550940	260360	260360	290580	290580
R-squared	.981	.981	.990	.989	.941	.939

Note: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantiles for 9 age categories, 4 education categories, 2 sex categories (in columns 3-6), and 8 years. Only observations with non-zero quantiles (employed) are included. All observations are weighted by the cell size.

Table 6
Quantile Regression of Job Duration
All Individuals Aged 21-64
(Year by Education Interaction)

Variable	All (1) Median	All (2) .9 Quantile	Males (3) Median	Males (4) .9 Quantile	Females (5) Median	Females (6) .9 Quantile
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
# of Cells	262	288	282	288	189	288
# of Obs	502600	550940	253860	260360	204050	290580
R-squared	.681	.984	.855	.991	.460	.943

Note: Numbers in parentheses are standard errors. The dependent variable is computed as cell quantile for 9 age categories, 4 education categories, 2 sex categories (in columns 3-6), and 8 years. Only observations with non-zero quantiles (employed) are included. All observations are weighted by the cell size. All specifications include eight dummy variables for age categories.

Table 7

Logit Analysis of Probability of Job Duration One Year or Less
All Employed Individuals Aged 21-64
(Year by Education Interaction)

Variable	All (1)	All (2)	Males (3)	Males (4)	Females (5)	Females (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
# of Obs	378892	378892	214211	214211	164681	164681
Log L	-194019.8	-193957.1	-102785.3	-102734.3	-90374.8	-90352.9

Note: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling 1 if job duration 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.

Table 8

Logit Analysis of Probability of Job Duration More than Ten Years
All Employed Individuals Aged 35-64
(Year by Education Interaction)

Variable	All (1)	All (2)	Males (3)	Males (4)	Females (5)	Females (6)
Constant	.383 (.0611)	.364 (.0967)	1.15 (.0792)	1.48 (.132)	-1.35 (.101)	-1.39 (.149)
Ed < 12	-.178 (.0125)	.658 (.162)	-.303 (.0162)	-.0159 (.209)	-.240 (.0212)	.930 (.276)
12 < Ed < 16	-.0570 (.0127)	-.642 (.167)	-.136 (.0169)	-1.06 (.219)	-.0721 (.0201)	-.984 (.275)
Ed >= 16	.111 (.0120)	-.0487 (.159)	-.133 (.0154)	-.972 (.203)	.237 (.0201)	.271 (.275)
Year	-.0012 (.0007)		-.0069 (.0009)		.0161 (.0012)	
Ed < 12 *Year		-.0112 (.0016)		-.0144 (.0020)		.0023 (.0028)
Ed = 12 *Year		.0001 (.0011)		-.0107 (.0016)		.0166 (.0018)
12 < Ed < 16 *Year		.0059 (.0016)		.00002 (.0021)		.0272 (.0027)
Ed >= 16 *Year		.0009 (.0015)		-.0008 (.0018)		.0162 (.0027)
p-value equality of time trends		<.0001		<.0001		<.0001
# of Obs	218491	218491	125300	125300	93191	93191
Log L	-141041.5	-141011.1	-82990.6	-82969.0	-54383.7	-54363.2

Note: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling 1 if job duration is more than ten 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 9

Logit Analysis of Probability of Job Duration More than Twenty Years
All Employed Individuals Aged 45-64
(Year by Education Interaction)

Variable	All (1)	(2)	Males (3)	(4)	Females (5)	(6)
Constant	-.0733 (.0900)	-.0732 (.132)	.407 (.109)	.551 (.178)	-1.96 (.177)	-1.92 (.263)
Ed < 12	-.144 (.0176)	.830 (.231)	-.312 (.0211)	.456 (.277)	-.213 (.0360)	1.17 (.472)
12 < Ed < 16	-.0558 (.0199)	-.819 (.258)	-.143 (.0243)	-1.45 (.314)	-.0891 (.0379)	-.701 (.507)
Ed >= 16	.103 (.0185)	-.345 (.242)	-.194 (.0221)	-.682 (.290)	.296 (.0365)	-.794 (.490)
Year	-.0079 (.0011)		-.0082 (.0013)		.0074 (.0021)	
Ed < 12 *Year		-.0199 (.0022)		-.0195 (.0026)		-.0099 (.0048)
Ed = 12 *Year		-.0079 (.0017)		-.0099 (.0021)		.0070 (.0031)
12 < Ed < 16 *Year		.0011 (.0025)		.0056 (.0031)		.0142 (.0051)
Ed >= 16 *Year		-.0026 (.0026)		-.0041 (.0027)		.0197 (.0048)
p-value equality of time trends		<.0001		<.0001		<.0001
# of Obs	122849	122849	71409	71409	51440	51440
Log L	-66675.9	-66652.6	-43954.4	-43933.4	-19432.2	-19421.4

Note: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling 1 if job duration is more than twenty 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.

Table 10
Logit Analysis of Probability of Job Duration One Year or Less
All Individuals Aged 21-64
(Year by Education Interaction)

Variable	All (1)	All (2)	Males (3)	Males (4)	Females (5)	Females (6)
Constant	.487 (.0386)	.713 (.0605)	-2.00 (.0598)	-2.45 (.101)	2.48 (.0539)	2.50 (.128)
Ed < 12	.535 (.0077)	-1.37 (.0996)	.611 (.0119)	-.469 (.154)	.682 (.0110)	-1.08 (.144)
12 < Ed < 16	-.104 (.0075)	.567 (.0991)	.0640 (.0117)	1.74 (.156)	-.126 (.0102)	.907 (.136)
Ed >= 16	-.450 (.0080)	-.253 (.107)	-.203 (.0122)	1.49 (.163)	-.438 (.0112)	.0013 (.150)
Year	-.0023 (.0004)		.0206 (.0007)		-.0210 (.0006)	
Ed < 12 *Year		.0182 (.0010)		.0393 (.0014)		.0001 (.0015)
Ed = 12 *Year		-.0050 (.0007)		.0261 (.0012)		-.0213 (.0009)
12 < Ed < 16 *Year		-.0128 (.0009)		.0062 (.0014)		-.0335 (.0013)
Ed >= 16 *Year		-.0073 (.0010)		.0060 (.0015)		-.0265 (.0015)
p-value equality of time trends		<.0001		<.0001		<.0001
# of Obs	550552	550552	260129	260129	290423	290423
Log L	-362625.5	-362320.8	-156831.9	-156637.3	-189164.4	-189006.7

Note: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling 1 if job duration 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.

Table 11

Logit Analysis of Probability of Job Duration More than Ten Years
All Individuals Aged 35-64
(Year by Education Interaction)

Variable	All (1)	All (2)	Males (3)	Males (4)	Females (5)	Females (6)
Constant	-.946 (.0542)	-1.20 (.0856)	.979 (.0726)	1.27 (.121)	-3.56 (.0897)	-3.66 (.133)
Ed < 12	-.390 (.0108)	1.43 (.141)	-.498 (.0144)	.455 (.188)	-.564 (.0184)	1.48 (.242)
12 < Ed < 16	.0638 (.0114)	-.612 (.150)	-.0730 (.0156)	-1.25 (.205)	.0469 (.0181)	-1.19 (.247)
Ed >= 16	.388 (.0109)	.319 (.145)	-.0548 (.0145)	-1.24 (.191)	.454 (.0181)	.187 (.247)
Year	-.0013 (.0006)		-.0173 (.0008)		.0240 (.0010)	
Ed < 12 *Year		-.0204 (.0014)		-.0326 (.0018)		.0005 (.0025)
Ed = 12 *Year		.0018 (.0010)		-.0208 (.0014)		.0252 (.0016)
12 < Ed < 16 *Year		.0097 (.0015)		-.0069 (.0019)		.0396 (.0024)
Ed >= 16 *Year		.0025 (.0014)		-.0055 (.0017)		.0283 (.0024)
p-value equality of time trends		<.0001		<.0001		<.0001
# of Obs	324121	324121	152987	152987	171134	171134
Log L	-185951.4	-185817.8	-99287.6	-99209.7	-75469.0	-75400.1

Note: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling 1 if job duration is more than ten 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 not having job duration more than ten years. The included age range in 35-64.

Table 12

Logit Analysis of Probability of Job Duration More than Twenty Years
All Individuals Aged 45-64
(Year by Education Interaction)

Variable	All (1)	All (2)	Males (3)	Males (4)	Females (5)	Females (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
# of Obs	197872	197872	92838	92838	105034	105034
Log L	-83594.0	-83523.8	-51275.6	-51225.1	-25048.5	-25022.4

Note: Numbers in parentheses are asymptotic standard errors. The dependent variable is a dummy variable equaling 1 if job duration is more than twenty 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 ten years. The included age range is 45-64.

APPENDIX I

Table A1
Median Job Duration by Age, Year, and Sex
(Employed Only)

All Workers:

Year	25-34	35-44	45-54	55-64
		Age Category		
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

Male Workers:

Year	25-34	35-44	45-54	55-64
		Age Category		
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

Female Workers:

Year	25-34	35-44	45-54	55-64
		Age Category		
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

Note: The statistics in this table for 1951 through 1968 are taken from BLS publications and are based on supplements to the Current Population Survey in January of the relevant year (U.S. Bureau of Labor Statistics; 1951, 1963, 1967, 1969). The statistics for 1973 through 1993 are 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 A2
0.9 Quantile Job Duration by Age, Year, and Sex
(Employed Only)

All Workers:

Year	25-34	35-44	45-54	55-64
		Age Category		
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

Male Workers:

Year	25-34	35-44	45-54	55-64
		Age Category		
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

Female Workers:

Year	25-34	35-44	45-54	55-64
		Age Category		
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

Note: The statistics for 1973 through 1993 are 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 A3
Median Job Duration by Age, Year, and Sex
(Population Based)

All Individuals:

Year	25-34	35-44	45-54	55-64
	Age Category			
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:

Year	25-34	35-44	45-54	55-64
	Age Category			
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:

Year	25-34	35-44	45-54	55-64
	Age Category			
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

Note: The statistics for 1973 through 1993 are 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 A4
0.9 Quantile Job Duration by Age, Year, and Sex
(Population Based)

All Individuals:

Year	25-34	35-44	45-54	55-64
		Age Category		
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:

Year	25-34	35-44	45-54	55-64
		Age Category		
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:

Year	25-34	35-44	45-54	55-64
		Age Category		
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

Note: The statistics for 1973 through 1993 are 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 A5
Fraction with Job Duration of One year or Less
(Employed Only)

All Employed :

Year	Age Category 25-34	35-44	45-54	55-64
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:

Year	Age Category 25-34	35-44	45-54	55-64
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:

Year	Age Category 25-34	35-44	45-54	55-64
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

Note: The statistics for 1973 through 1993 are 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 A6
Fraction with Job Duration of More than Ten Years
(Employed Only)

All Employed :

Year	25-34	35-44	45-54	55-64
	Age Category			
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:

Year	25-34	35-44	45-54	55-64
	Age Category			
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:

Year	25-34	35-44	45-54	55-64
	Age Category			
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

Note: The statistics for 1973 through 1993 are 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 A7
Fraction with Job Duration of More Than Twenty Years
(Employed Only)

All Employed :

Year	25-34	Age Category 35-44	45-54	55-64
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:

Year	25-34	Age Category 35-44	45-54	55-64
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:

Year	25-34	Age Category 35-44	45-54	55-64
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

Note: The statistics for 1973 through 1993 are 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 A8
Fraction with Job Duration of One year or Less
(Population Based)

All Individuals:

Year	25-34	35-44	45-54	55-64
		Age Category		
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:

Year	25-34	35-44	45-54	55-64
		Age Category		
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:

Year	25-34	35-44	45-54	55-64
		Age Category		
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

Note: The statistics for 1973 through 1993 are 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 A9
Fraction with Job Duration of More than Ten Years
(Population Based)

All Individuals:

Year	25-34	35-44	45-54	55-64
		Age Category		
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:

Year	25-34	35-44	45-54	55-64
		Age Category		
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:

Year	25-34	35-44	45-54	55-64
		Age Category		
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

Note: The statistics for 1973 through 1993 are 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 A10
Fraction with Job Duration of More than Twenty Years
(Population Based)

All Indivduals:

Year	25-34	Age Category 35-44	45-54	55-64
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

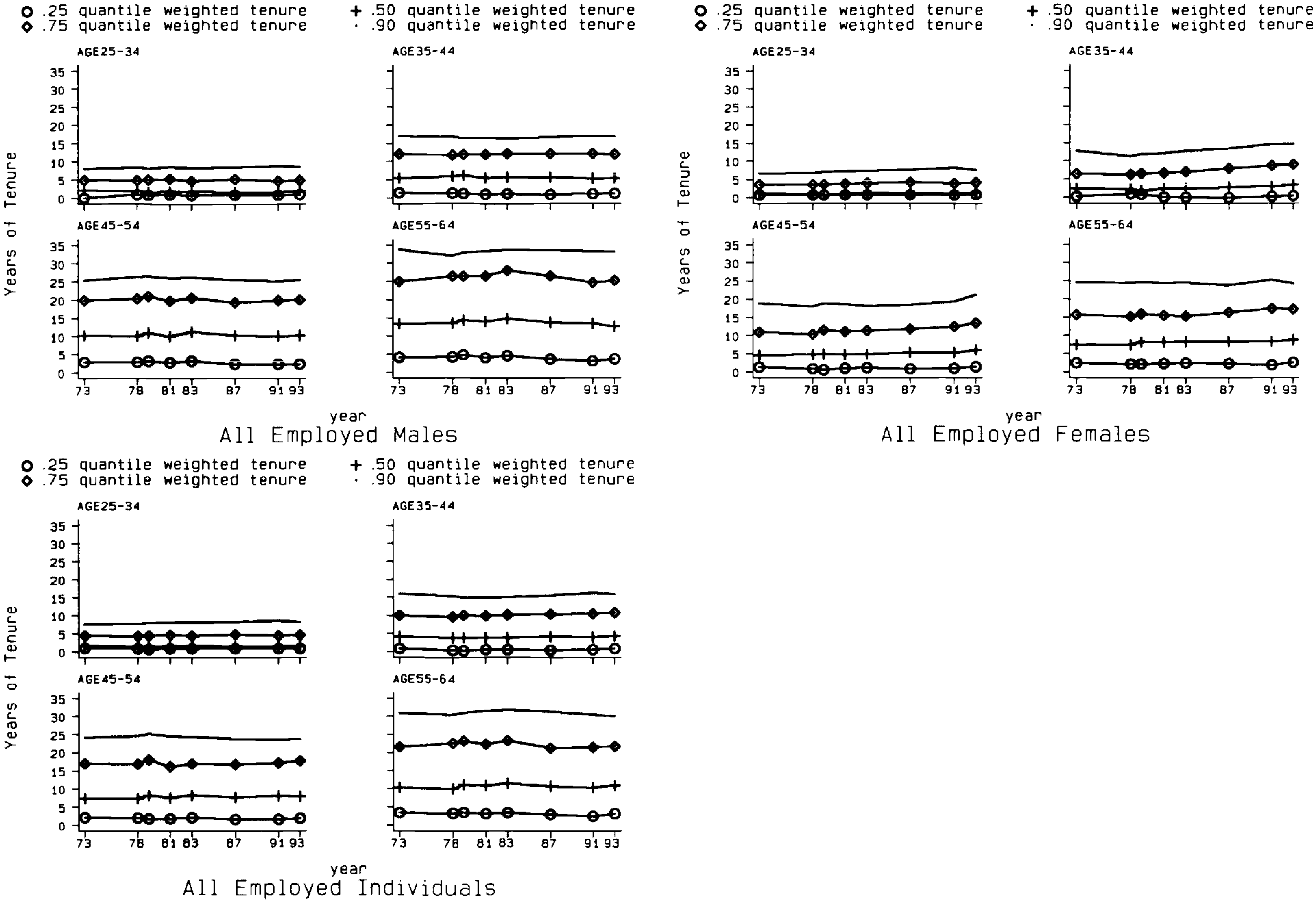
All Males:

Year	25-34	Age Category 35-44	45-54	55-64
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:

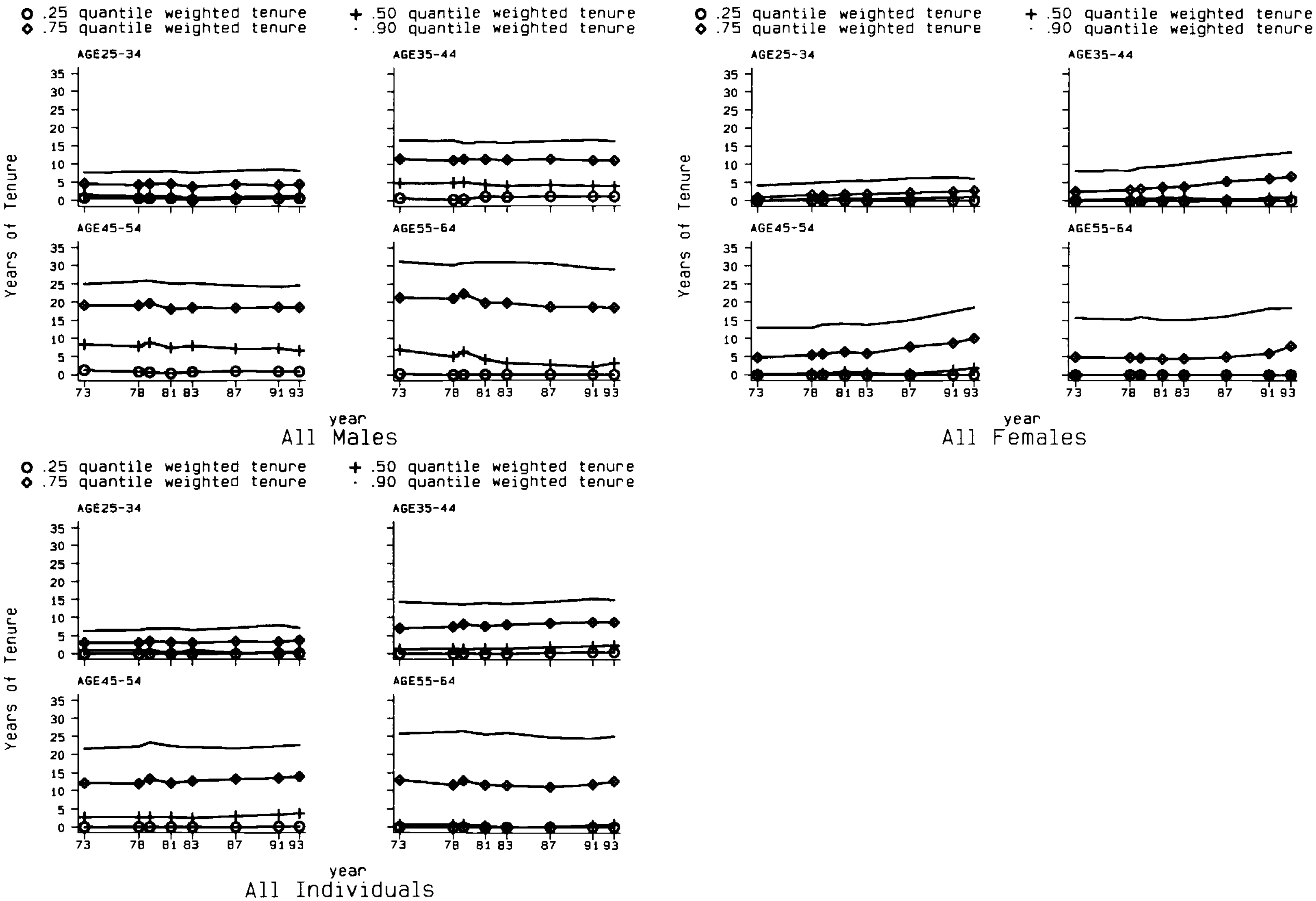
Year	25-34	Age Category 35-44	45-54	55-64
1973	.000	.017	.049	.072
1978	.000	.012	.048	.073
1979	.000	.011	.052	.072
1981	.000	.013	.055	.071
1983	.000	.010	.043	.067
1987	.000	.008	.050	.065
1991	.000	.020	.073	.084
1993	.000	.022	.092	.087

Note: The statistics for 1973 through 1993 are 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.



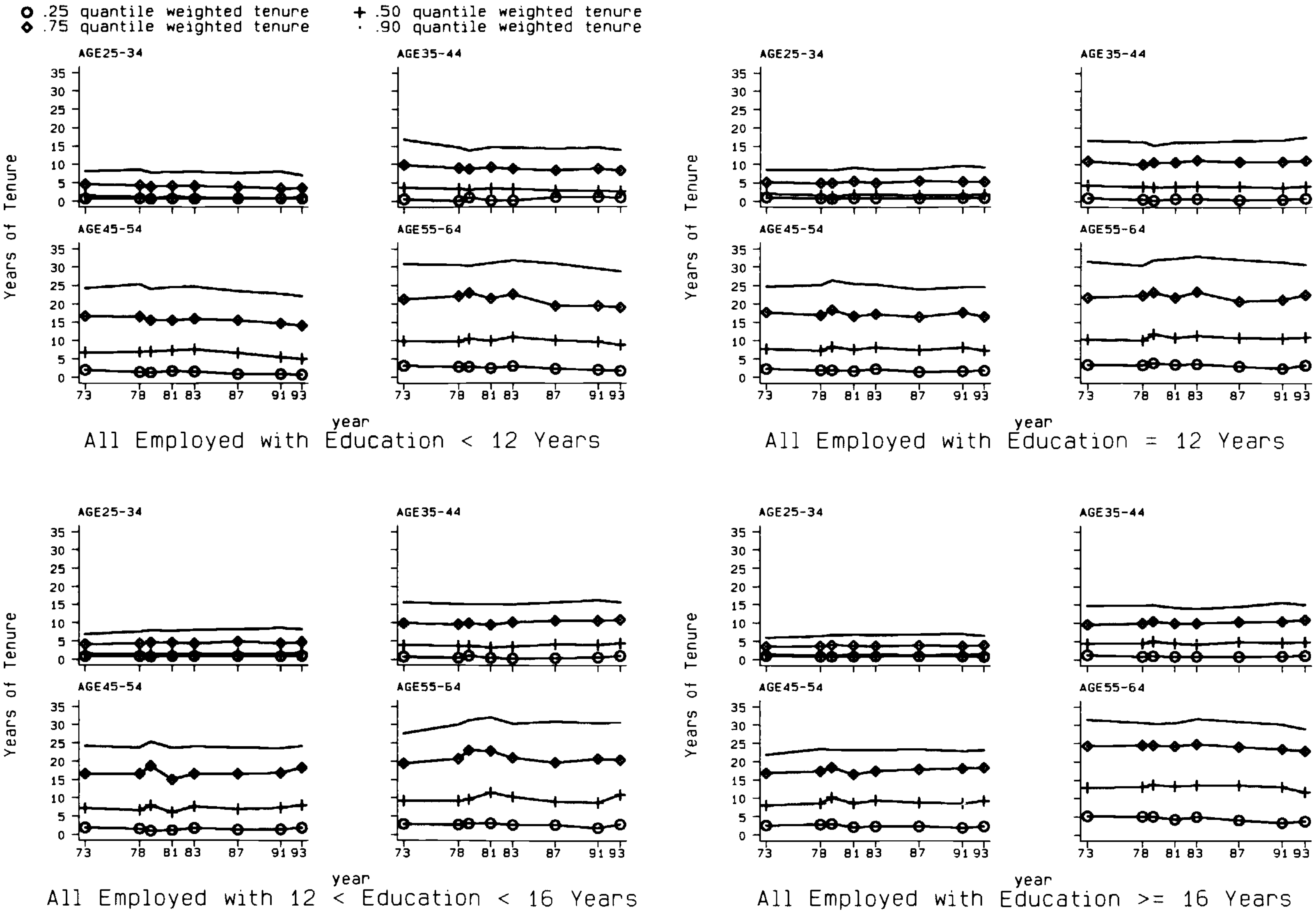
Employed by Sex

Figure 1 - Quantiles of Tenure Distribution by Year



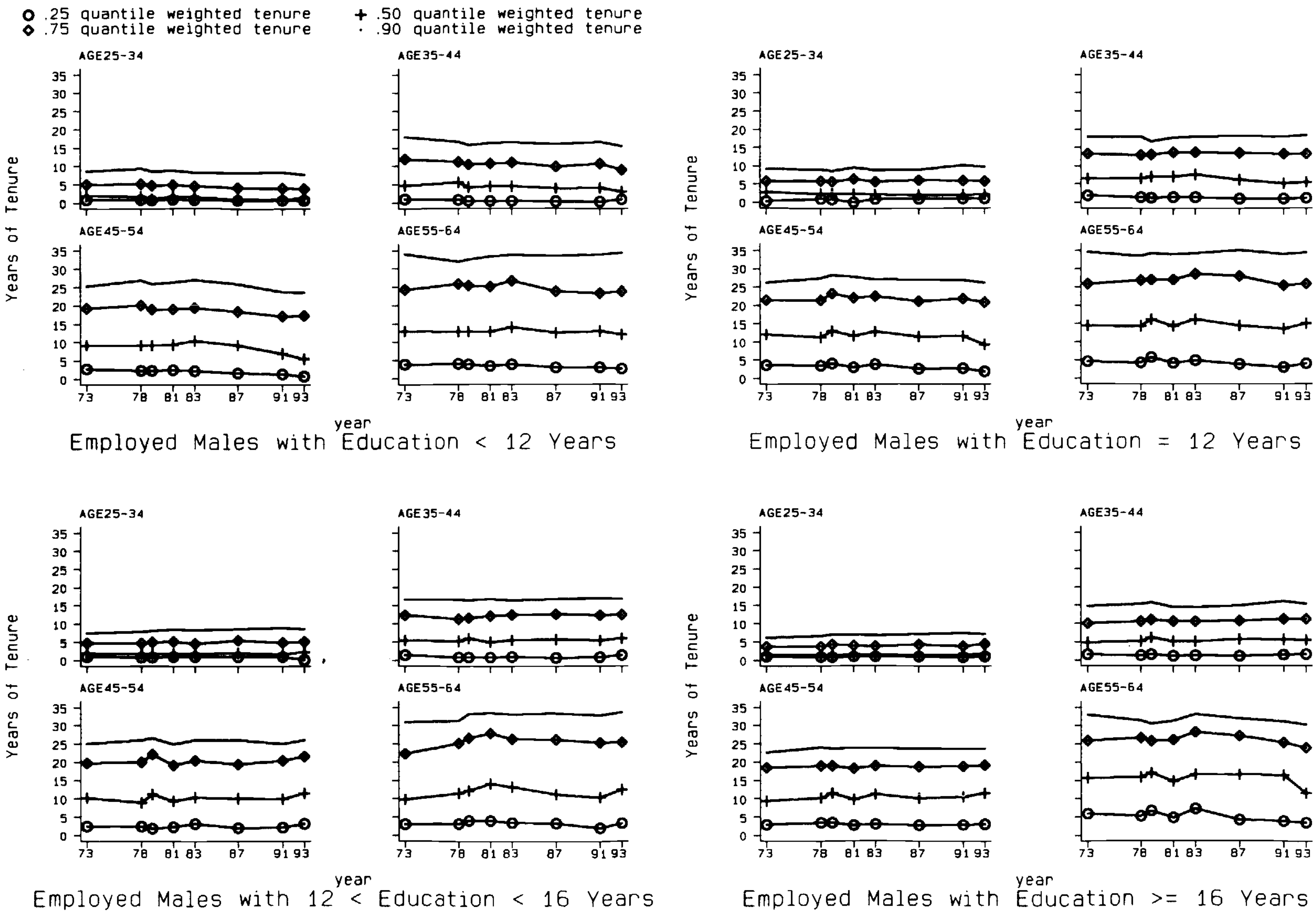
All Individuals by Sex

Figure 2 - Quantiles of Tenure Distribution by Year



All Employed Individuals by Education

Figure 3 - Quantiles of Tenure Distribution by Year



All Employed Males by Education

Figure 4 - Quantiles of Tenure Distribution by Year

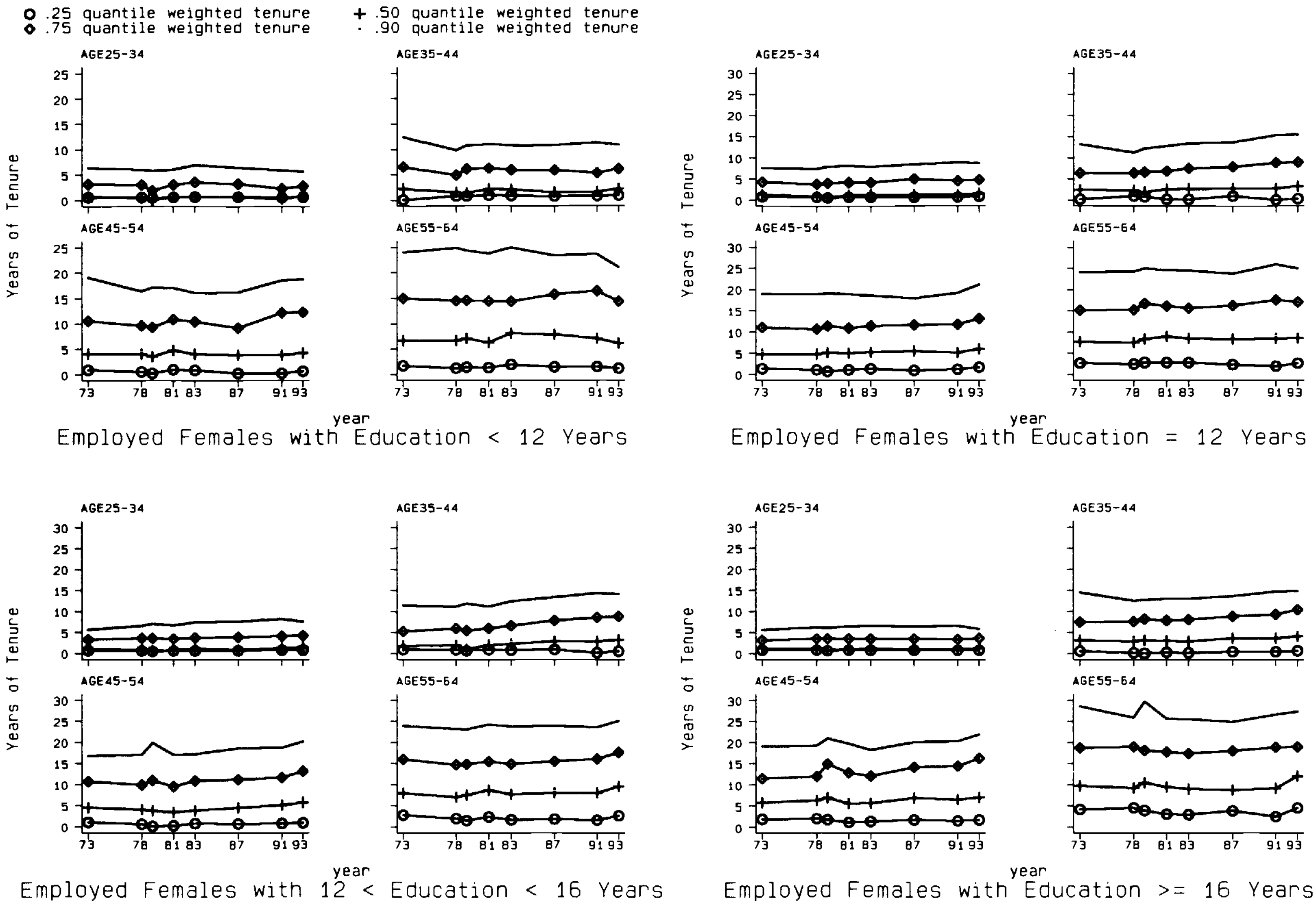
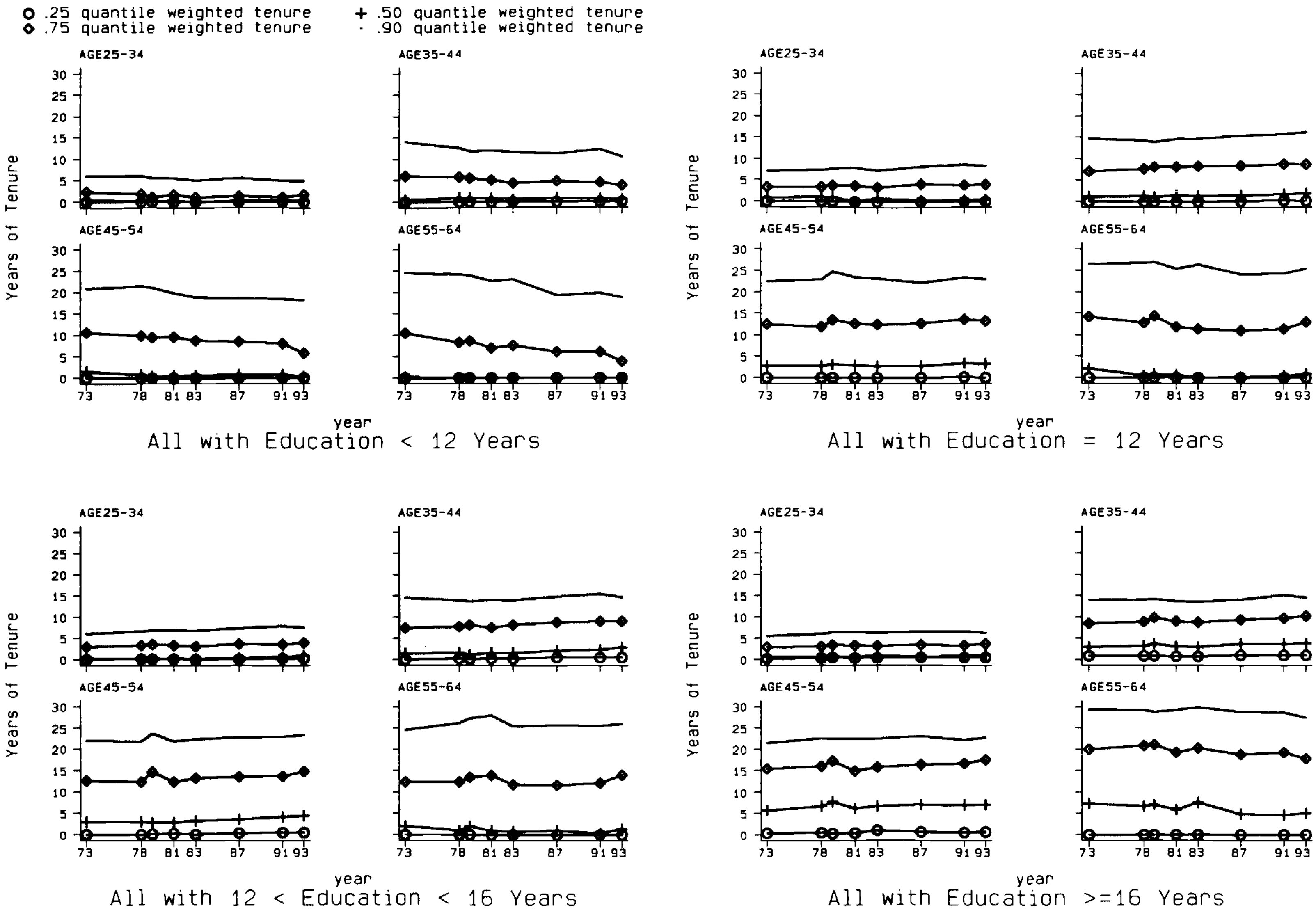
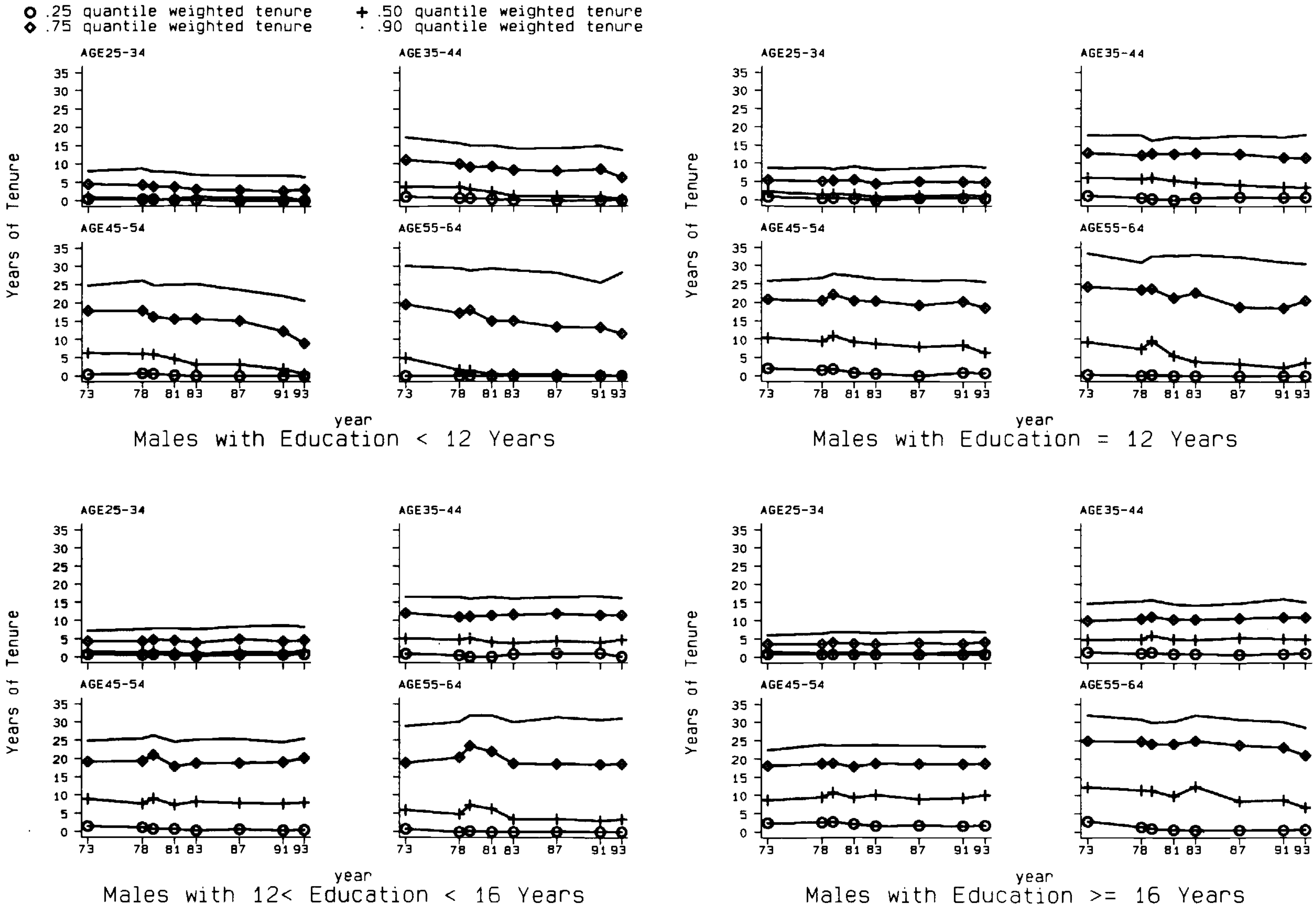


Figure 5 - Quantiles of Tenure Distribution by Year



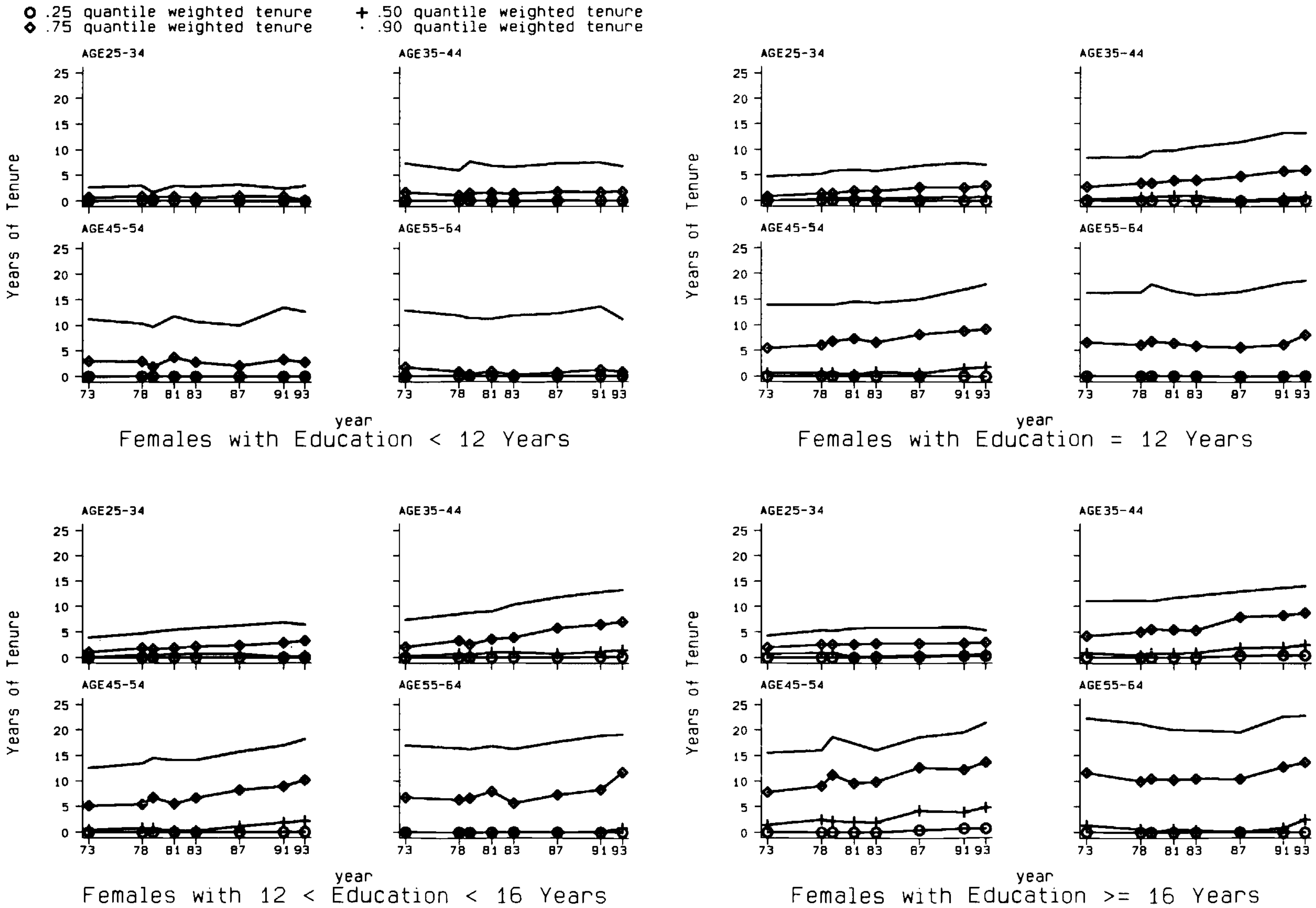
All Individuals by Education

Figure 6 - Quantiles of Tenure Distribution by Year



All Males by Education

Figure 7 - Quantiles of Tenure Distribution by Year



All Females by Education

Figure 8 - Quantiles of Tenure Distribution by Year