THE INCIDENCE AND COSTS OF JOB LOSS: 1982-1991

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ABSTRACT

I focus on two aspects of job loss. First, I examine evidence on the incidence of job loss by worker and job characteristics including age, education, race, sex, industry, and tenure over the period from 1982 to 1991. Second, I examine the cost of job loss to workers in the form of 1) lower post-displacement employment probabilities, 2) lower probabilities of full-time employment for re-employed workers, and 3) lower earnings for full-time workers.

Using data from Displaced Workers Surveys (DWS) from 1984 through 1992 to study job loss from 1982-91, I find that older workers and more educated workers are relatively more likely to suffer a job loss in the latter part of this period. Additionally, job loss became more common in some important service industries and relatively less common in manufacturing during the latter part of the ten-year period studied.

Supplementing the DWS data with data from the outgoing rotation groups of the Current Population Survey from 1982-1991, I find that displaced workers are, relative to non-displaced workers, 1) less likely to be employed and 2) more likely to be employed part-time conditional on being employed. These effects seem to decline with time since displacement. There is no systematic secular change in these costs of displacement, either in the aggregate or for particular groups. Finally, I examine the earnings losses of full-time re-employed displaced workers by comparing their earnings change with the earnings change of full-time employed workers who were not displaced. I find, consistent with what others have found, that these earnings losses are substantial.

Overall, the costs to displaced workers of job loss are substantial and come in several forms. However, the public perception that the current sluggish economy is worse than earlier downturns may reflect more who has lost jobs recently rather than either increased overall job loss or increased costs to those who are losing jobs.

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1. Introduction

There is a public perception that the nature and consequences of job loss, defined here as the involuntary (from the worker's viewpoint) termination of employment with a particular firm, have changed qualitatively in recent years. The perception is that highly-skilled white-collar workers and workers with more tenure (time with their current employer) becoming increasingly vulnerable to job loss, reduced subsequent earnings, and prolonged unemployment (N.Y. Times, November 22, 1992, Section 3, pp. 1,6). My goal in this study is to investigate whether and how the incidence and costs of job loss have changed in the last ten years, i.e., is the public perception correct.

Due to data limitations, it is difficult to get a perspective on who loses jobs prior to the 1980's. However, the Displaced Workers Surveys (DWS), which have been regular supplements to the January Current Population Survey (CPS) at two year intervals since 1984, have useful information on job loss. Specifically, these surveys ask workers if "In the past five years . . . has _____ lost or left a job because of a plant closing, an employer going out of business, a layoff from which _____ was not recalled or other similar reason." These data have much to tell us about job loss, and, augmented with data from mobility supplements the the January CPS in 1983, 1987, and 1991 and the merged outgoing rotation group CPS files from 1982-91, they form the basis of my empirical analysis.

Specifically, I focus on two aspects of job loss. First, I examine evidence on the incidence of job loss by worker and job characteristics including age, education, race, sex, industry, and tenure over the period from 1982 to 1991 in order to determine the extent to which there has, in fact, been an increase in the vulnerability of higher-skilled workers to job loss. This period covers an entire business cycle, with slack

labor markets in 1982-83 and 1990-91 and an intervening expansion from 1984 through 1989.¹ Additionally, this is a continuing period of sectoral shifts in the composition of employment away from goods-producing industries and toward service industries.

The second aspect of job loss I focus on is its cost to the workers who are displaced. I examine evidence on the post-displacement employment and earnings experience of displaced workers with various characteristics in order to measure both the costs of displacement and if and how these costs have changed.

Potential changes in the incidence and costs of job loss are of more than passing interest. Long-term employment relationships have played a central role in the U.S. economy, particularly for highly-skilled workers.² This is consistent with the existence of institutional arrangements that make younger, less-skilled, and less-senior workers bear the brunt of downward adjustments in employment levels. Declines in labor demand are generally accommodated by laying off the least senior workers.³ Historically, it has

¹I use the labor economists casual definition of a recession as periods of relatively high unemployment compared with surrounding periods. The unemployment rate was 9.5 percent in 1982 and 1983, and it is about 8 percent in 1992. The unemployment rate has, in fact, increased since 1991, so that a "recession" continues by my definition.

²Tabulations of the mobility supplement to the January 1991 Current Population Survey (CPS) show that median job tenure (time with current employer) by education level for men aged 31-60 is:

Yrs. Educ.	Med. Yrs. Tenure	Yrs. Educ.	Med. Yrs. Tenure
<12	2.4	13-15	6.2
12	6.1	≥16	8.2.

Tabulations of mobility supplements to earlier CPS's show a similar pattern by education category. Hall (1982) and Ureta (1992) present analyses of reported job tenure based on earlier CPS's that highlight the importance of long-term jobs.

³While more prevalent in the union sector, inverse seniority rules for layoff are also quite common in the nonunion sector (Abraham and Medoff, 1984).

also been the case that less-skilled (blue-collar) workers are more susceptible to layoffs than have more-skilled (white-collar) workers. These institutional regularities may be an efficient response to the desire to have more job-specific capital embodied in higher-tenure and more-skilled workers. To the extent that the job security of these workers has become more tenuous, the willingness and ability of firms and workers to invest in job-specific capital is reduced with adverse consequences for productivity.

Some amount of job change plays a positive allocative role in assuring that workers are matched with appropriate employers and that matches that become inappropriate can end. However, any job change has costs in a number of dimensions. Job-specific capital is lost, and the longer the relationship that ends, the more specific capital is likely to be lost. Additionally, workers (and perhaps firms) value stability *per se*. Finally, termination of a job may result in periods of non-employment which have both private and social costs. More generally, long-term employment relationships may be important components of productivity growth. This productivity growth may well go beyond the usual growth in individual productivity that we associate with aging and encompass the human-capital component of new technologies.

Changes in the incidence of job loss may be related to other social and economic phenomena of broader interest. First, consider the public mood, which, as we have very recently seen, can have serious political, and perhaps economic, ramifications. The relatively slow growth of the economy (perhaps even a continuing recession) in the last two years has resulted in public unhappiness with the economy that some have argued is far more extreme than warranted when measured against the "objective" economic situation in earlier recessions. For example, the so-called misery index (the sum of the unemployment and inflation rates) stood at 16.7 percent in 1983 (9.5 percent unemployment plus 7.3 percent inflation), and it is only about 11 percent in

1992 (8 percent unemployment plus 3 percent inflation). One factor that might account for this difference in mood is a difference in the incidence of job loss by age, education, and tenure groups in the recent period relative to earlier recessions. Specifically, it may be the case that because those workers who have been best protected in a "typical" recession (older workers with more education) have seen more job loss than usual and because these same workers may be more influential, the public mood seems darker.

With increased import competition signaling dramatic sectoral shifts in employment over the past ten to twenty years, it is plausible that older high-tenure workers are more likely to be at risk to lose their jobs than was previously the case and that existing institutional arrangements may be less effective at protecting older and higher-skilled workers than they were in the past.

The next section contains summary statistics of job loss over the 1982-91 period as a function of worker characteristics including age, sex, race, and education. I find that job loss in the earlier part of the period was relatively heavily weighted toward younger and less educated workers. In contrast, job loss in the last two years of the period was relatively heavily weighted toward older and more highly educated workers.

In section 3, I briefly examine the incidence of job loss by industry over the 1982-91 period. Not surprisingly, I find that job loss throughout the period was disproportionately in manufacturing. However, job loss was substantially less concentrated in manufacturing in the 1986-91 period relative to the 1982-85 period. In recent years, workers in the trade; finance, insurance, and real estate; and professional service sectors have borne significantly more job loss than was the case earlier.

Section 4 contains a multivariate analysis of job loss by tenure, education, sex, and age. As expected, the probability of job loss is

inversely related to tenure, education, and age. The finding that older and more educated workers had relatively higher probabilities of job loss in the later part of the 1982-91 period is corroborated by this analysis. However, there was no systematic change in the relationship of job loss with tenure during the decade studied. I discuss some implications of the results on the incidence of job loss in section 5.

In section 6, I use the DWS data to examine the post-displacement labor force status of workers who lost jobs in the 1982-91 period. I consider re-employment, unemployment, and out-of-the-labor-force probabilities along with the probability of full-time versus part-time employment and re-employment wages. These estimates show that re-employment probabilities are lower for workers displaced during the two recessionary periods (1982-83 and 1990-91) and that males had lower re-employment probabilities in the most recent recession relative to the earlier recession. I also find that nonwhites and workers with less education are less likely to be employed after a job loss. There is no evidence that older and/or more educated displaced workers have become more or less likely than to be re-employed than they were in the past. The estimates also suggest that the dis-employment effect of displacement declines with time since displacement so that some of the cost of displacement in this dimension is temporary.

Section 7 examines the part-time/full-time status of jobs held post-displacement. I find that displaced workers are more likely to be working part-time after displacement even controlling for part-time/full-time status prior to the job loss. Additionally, workers displaced during the slack labor market periods and subsequently re-employed are more likely to be working part-time after displacement than are those displaced during the expansion. Again, there is no difference between the two recessionary periods and no systematic secular change. The probability of post-

displacement full-time employment increases with time since displacement so that at least some of the cost of displacement from an increased incidence of part-time employment is temporary.

Finally, I estimate wage losses associated with job displacement in section 8. I find, as others have (e.g., Podgursky and Swaim, 1987), that there is a substantial wage penalty for job loss and that this penalty is more severe for older, more senior, and less educated workers. Wage losses from displacement are found to be larger for workers who lost jobs in the 1990-91 period, but there is no evidence that wage changes associated with displacement differ across the time period studied in any systematic way. Section 9 contains some concluding remarks.

2. The Incidence of Job Loss by Worker Characteristics

In this section I analyze data on individuals between the ages of twenty and sixty-four from the Displaced Worker Supplements (DWS) to the January Current Population Surveys (CPS) in 1984, 1986, 1988, 1990, and 1992 in order to investigate the incidence of job loss by individual characteristics. To my knowledge, no one has used these data for this purpose previously. All of the work of which I am aware using these data focuses on the post-displacement employment and earnings experience of displaced workers (e.g., Podgursky and Swaim, 1987; Kletzer, 1989; Gibbons and Katz, 1991).⁴

Each Displaced Workers Supplement asks workers if they were displaced from a job at any time in the preceding five-year period. For several reasons, I restrict my analysis to whether a worker reports displacement within the two-year period prior to each interview. First, if I used all

⁴I focus on this issue later in this paper.

five years, the recall periods would overlap. Second, workers may fail to recall job loss that occurred long before the interview date. Third, the survey collects and reports information on at most one job loss for each individual. For workers with more than one job loss, this information refers to the most recent job loss. Because new jobs are more likely to end than jobs that have survived for a while (Farber, 1992) and because of recall problems, it is likely that the DWS's seriously underestimate job loss that occurred long before the interview date.

The extent of this underestimation of job loss is demonstrated by a comparison of counts of overall job loss within two years of each interview (e.g., 1982-83 for the January 1984 DWS) with counts of overall job loss in the same calendar period derived from the third and fourth year prior to interviews two years later (e.g., 1982-83 for the January 1986 DWS). The weighted job loss computed for the four two-year periods from 1982 through 1989 from immediately succeeding interviews is fully 44 percent higher than the job loss computed for the same periods from interviews conducted two years later.

I proceed using the two-year recall periods. These are 1982-83 for the 1984 DWS, 1984-85 for the 1986 DWS, 1986-87 for the 1988 DWS, 1988-89 for the 1990 DWS and 1990-91 for the 1992 DWS. This is a fortuitous breakdown of the 1982-91 time period. The first two years, 1982-83, represents a fairly deep recession, with unemployment at 9.5 percent each year. The last two years, 1990-91, also represents a recession (that perhaps has not yet ended) with unemployment running over 6.5 percent by 1991. The intervening six years, 1984-90, represents a sustained expansion. I exploit this configuration of dates to compare the incidence of job loss over the cycle and in the two recessions.

A. Job Loss by Age and Sex

Table 1 presents information on the fraction of the population by sex and age category who report a displacement within two years at each interview date. The results for males, contained in the top panel, show that younger workers have higher rates of job loss than do older workers. They also show that rates of job loss were generally higher for males in the two slack labor markets (1982-83 and 1990-91). However, there is a sharp contrast between the two recessions in rates of job loss by age group. The 1982-83 recession had the expected significantly higher rate of job loss for young workers relative to the years of expansion (1984-89).⁵ The cyclical differences are less pronounced for older males, though they are still significantly different from zero. The 1990-91 recession shows a sharply different pattern. The rate of job loss for younger workers is higher than in the boom period, but not quite as high as in the 1982-83 recession. The more striking finding is that the rate of job loss among older workers in the 1990-91 recession is not only higher than in the boom period but it is also much higher than in the 1982-83 recession. This suggests that job loss in the recent period is, in fact, of a different character than earlier job loss. Older prime-age male workers (40-60) were significantly more likely to lose their jobs in the 1990-91 period than in the 1982-83 period.

The bottom panel of table 1 contains the same job-loss information for females. First note that females have lower job loss rates than men. This at least partly reflects their lower labor-force participation rates.⁶ The

⁵The standard errors on these differences in rates of job loss are all less than 0.006 so that differences of more than 0.012 are statistically significantly different at the 0.05 level.

⁶These job loss rates are computed over the entire population in each age category. It would not make sense to compute the job-loss rate over the only the employed workers in each Displaced Workers Survey because workers who

results with regard to job loss by females over time are rather different than for males. The job-loss rates for females were slightly but generally insignificantly higher for virtually all age categories in both the 1982-83 and the 1990-91 recessions than in the intervening expansion. However, a joint test of the hypothesis that that the 1982-83 period had job loss rates no different from the succeeding expansion results in clear rejection of that hypothesis ($p\text{-value} < 0.001$). Similarly, a joint test of the hypothesis that that the 1990-91 period had job loss rates no different from the preceding boom periods also results in clear rejection of that hypothesis ($p\text{-value} < 0.001$). There seems to be no overall difference in job loss rates between the two recessionary periods ($p\text{-value} = 0.550$).⁷ Thus, recessions do result in higher rates of job loss for females, but the 1990-91 period does not seem qualitatively different from the earlier recessionary period.⁸

B. Job Loss by Educational Category

Consider next how rates of job loss vary by education level. Figure 1 contains a graphic representation of two-year population-based rates of job loss by education group and age category for males in three two-year recall periods (1982-83, 1986-87, 1990-91).⁹ Rates of job loss for younger males in all educational categories are lower in the expansion (1986-87) than in

were displaced might not be employed at the time of the DWS.

⁷These tests are based on weighted grouped-logit estimations of job loss rates assuming only main age and year effects so that there are no age-year interactions.

⁸I also calculated the job-loss rates on an employment basis. I do this crudely by dividing the population-based job loss rates by the current employment rates in each group. These estimates are not presented here, but they show essentially the same patterns as the population based estimates in table 1.

⁹The other two recall periods (1984-85 and 1988-89) are not included in order to keep the graph clear. These omitted two periods, which are both part of the 1980's expansion, look very much like 1986-87.

either of the recessionary periods. The rate of job loss for older males in 1986-87 is substantially below the rates in the recessionary periods only for males with less than twelve years education. There is a striking contrast between rates of job loss by education category in the earlier periods and the latest period. Older prime-age (40-60) men in all educational categories have somewhat higher rates of job loss in the 1990-91 period relative to the 1982-83 and 1986-87 period, but the differences are substantially larger and more consistent across age categories in the groups with higher educational attainment. In particular, men 40-60 years of age with at least 16 years of education had a much higher probability of job loss in 1990-91 than men of the same description in 1982-83 or 1986-87. The average rate of job loss among males in these age and education categories was 0.0462 in 1990-91 compared with 0.0302 in 1982-83.

The results for females are generally similar to those for males, though at a lower level. Figure 2 contains a graphic representation of two-year employment-based rates of job loss by education group and age category for females. Although these results are noisier than those for males, rates of job loss for younger females are generally lower in the expansion for each education group. As with the males, older prime-age females with more education had higher rates of job loss in the 1990-91 period than in the earlier periods.

C. Multivariate Analysis of Job Loss

Table 2 contains logit estimates of the probability of job loss. In order to allow sufficient flexibility and to investigate the extent to which the incidence of job loss by age and education has changed, the model is estimated separately for each educational category and for younger (20-34 years of age) and older (35-59 years of age) individuals. These results corroborate what we see in table 1 and figures 1-2: 1) females have lower

rates of job loss than males; 2) rates of job loss are lower in the years of expansion; 3) Younger workers with <12 years of education fared substantially worse in the 1982-83 recession than in the most recent downturn; 4) older workers were more vulnerable to job loss in the recent downturn than in the base (1982-83) period. The last finding is particularly strong at the highest education levels, where the predicted job loss rate in 1990-91 is over thirty percent higher than the predicted job loss rate in the 1982-83 period.¹⁰

Another result in table 2 is worth mentioning. There seems to be no significant difference in rates of job loss by race at any education level for either age group. An additional logit analysis, not shown here, that allows separate race effects for each year, shows no significant difference in rates of job loss by race and year. Thus, there is no evidence that nonwhites fared worse than whites in terms of job over the period studied.

3. The Locus of Job Loss by Industry

It is difficult to use the data from the Displaced Workers Survey to study the relationship between job attributes and the probability of job loss because the data are not truly longitudinal. Specifically, all workers are not asked about their jobs at one point in time and then re-interviewed at a future date to find out about job loss. Rather, workers are interviewed at one point in time and asked whether they have lost a job in the recent past (the recall period). Only if workers had been displaced were they asked about the job from which they were at risk to lose. Thus, no information is available at the individual level on characteristics of jobs (e.g., industry,

¹⁰Since the probability of job loss is quite low (approximately 0.05), the logit coefficient can be interpreted as the proportional effect of the relevant variable on the probability of job loss.

occupation, tenure) held during the recall period unless the workers either were displaced or did not change jobs voluntarily.

This is a weakness of using purely cross-sectional data to study job loss, and it precludes augmenting an individual-level multivariate logit model of the sort in table 3 with industry controls. However, it is possible to construct the probability of job loss by industry by using Bayes' rule to combine 1) industry employment shares, 2) industry job-loss shares (the fraction of job loss in a given year that comes from each industry), and 3) the overall probability of displacement. On this basis, the probability that worker i in industry j is displaced in period t is

$$(3.1) \quad \Pr(D_{it}=1 | I_{it}=j) = \Pr(I_{it}=j | D_{it}=1) \cdot \Pr(D_{it}=1) / \Pr(I_{it}=j)$$

where D_{it} represents a binary variable which equals one if worker i suffers a job loss in period t and I_{it} equals j if worker i is employed in industry j in period t .

Each of the three probabilities required to calculate $\Pr(D_{it}=1 | I_{it}=j)$ are easily computed. The first piece, $\Pr(I_{it}=j | D_{it}=1)$, is computed directly as the industrial distribution of workers displaced in period t . The second piece, $\Pr(D_{it}=1)$, is computed as the unconditional probability that an individual is displaced in period t . The last piece, $\Pr(I_{it}=j)$, is computed as the industrial distribution of employment, and, for convenience, I use the industrial distribution of employment as of the January CPS containing the relevant Displaced Workers Supplement (e.g., January 1984 for the 1982-83 recall period). While this will differ somewhat from the basically unobservable (and variable) industrial distribution of employment during the recall period, the differences are likely to be minor.

Table 3 contains data for each of the five two-year periods on industry employment shares, the industrial distribution of displaced workers, and the

probability of job loss by industry as computed using equation 3.1.¹¹ It is clear from this table that the recent years are rather different from the earlier years. First, consider the share of job loss by industry in the middle panel of table 3. Throughout the 1982-91 period, manufacturing workers have disproportionately lost their jobs in that manufacturing's share of total job loss exceeds manufacturing's employment share in every period.¹² However, job loss was less substantially less concentrated in manufacturing after 1985. In the 1982-85 period, about three-eighths of all job loss was in manufacturing employment. In the 1986-91 period "only" about one-fourth of all job loss was from manufacturing. Three industry groups, trade; finance, insurance and real estate; and professional services each had an increase in the share of job loss of three to four percentage points to make up the difference.

Next examine the probability of job loss by industry in the bottom panel of table 3. While job loss is predictably higher in the two recessionary periods, the probability of job loss in manufacturing was more than two percentage points lower in 1990-91 than in 1982-83. Over the same period, the probability of job loss grew by almost two percentage points in trade and by over three percentage points in finance, insurance, and real estate. Thus, job loss has become relatively more common in important and growing non-goods producing industries.

¹¹The overall probabilities of displacement ($\Pr(D_{it}=1)$) are .0578 for 1982-83, .0478 for 1984-85, .0436 for 1986-87, .0399 for 1988-89, and .0590 for 1990-91.

¹²This is not the same thing as saying that manufacturing's share of total employment fell in every period. While this is also true, the job loss measure does not account for new hiring or for voluntary job change. Davis and Haltiwanger (1992) present an analysis of net job change in manufacturing using establishment data.

4. Job Loss by Tenure, Education, Sex, and Age: A Multivariate Analysis

A key component of the perceived change in the locus of job loss is that higher tenure individuals are now more at risk for job loss than they were. Unfortunately, as with industry, tenure during the recall period is not observed for workers who do not report that they have been displaced. I can use the same Bayes'-rule approach to recover the probability of job loss by tenure category that I used to compute the probability of job loss by industry. The appropriate relationship defining the probability that worker i in tenure category j is displaced in period t is

$$(3.2) \quad \Pr(D_{it}=1|T_{it}=j) = \Pr(T_{it}=j|D_{it}=1) \cdot \Pr(D_{it}=1) / \Pr(T_{it}=j)$$

where D_{it} equals one if worker i in period t suffers a job loss and $T_{it}=j$ if worker i is in tenure category j in period t .

The only piece of equation 3.2 that is not available directly from the DWS is $\Pr(T_{it}=j)$, the unconditional tenure distribution. However, the January CPS's in the years immediately prior to the 1984, 1988, and 1992 Displaced Workers Surveys contain mobility supplements that have information on job tenure for all workers that I use to compute unconditional tenure distributions in January 1983, 1987, and 1991. These mobility supplements are ideally timed in the sense that they fall almost precisely in the middle of the 1982-83, 1986-87, and 1990-91 recall periods associated with the 1984, 1988, and 1992 DWS's respectively. Unfortunately I have no information on the unconditional distribution of job tenure that are appropriate for the 1984-85 or 1988-89 recall periods.

Calculations identical to those performed for industry in the last section yield the unsurprising result that the probability of job loss is strongly monotonically declining in tenure. While not presented here, the probability of job loss (averaged over the six years) falls from about 10

percent for workers in their first year on the job to less than 3 percent for workers with more than ten years on the job. The results are more remarkable for what is not found. There seems to be no substantial difference in the tenure-specific probability of job loss between the two recessionary periods. The earlier finding, that older workers fared relatively worse in the recent downturn than in the 1982-83 downturn, suggested that I would find that workers with more tenure (who are also older on average) also fared worse in the more recent period. This does not seem to be the case, which suggests indirectly that the older workers with relatively low tenure conditional on their age were relatively more likely to suffer a job loss in the most recent period.

Covariation between age, education, and tenure in determining the rate of job loss, suggests that a multivariate analysis of job loss would be useful. However, without longitudinal data on all variables, a micro-level analysis is not feasible. However, I can compute cell-mean level rates of job loss using the Bayes' Rule approach. I compute job loss rates for groups defined by education category, tenure category, age category, sex, and time period. There are 1072 cells, defined by these five variables, for which I have computed job loss rates using information on individual characteristics and job displacement from the 1984, 1988, and 1992 Displaced Workers Surveys combined with the data from the mobility supplements to the January CPS's in 1983, 1987, and 1991. Unfortunately, cell sizes are too small if I expand the breakdown to include race or industry.

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

$$\begin{aligned}
 (3.3) \quad & \Pr(D_{it}=1 | T_{it}=j, S_{it}=k, A_{it}=r, E_{it}=m) \\
 &= D_{ijkrm} \\
 &= \Pr(T_{ijkrm}=j | D_{ijkrm}=1) \cdot \Pr(D_{ijkrm}=1) / \Pr(T_{ijkrm}=j)
 \end{aligned}$$

where $T_{it}=j$ if individual i is in tenure category j year t , $S_{it}=k$ if individual i is of sex k (in year t for symmetry), $A_{it}=r$ if individual i is in age category r , and $E_{it}=m$ if individual i is in educational category m in year t . The quantity $D_{ijkrm}=1$ if individual i in tenure category j , of sex k , in age category r , and with education level m in year t is displaced from his or her job. $T_{ijkrm}=j$ if individual i of sex k in age category r with education level m is in tenure category j in period t . Finally, $D_{ijkrm}=1$ if individual i of sex k with education level m in age category r is displaced from his or her job in period t . This particular representation of the conditional probability of job loss is used because it allows computation of the probability of displacement conditional on tenure and the other controls without direct information on the joint distribution of displacement and tenure.

Preliminary analysis of these data show the usual result that the intermediate time period, because it is in the middle of an expansion, has lower rates of job loss in virtually all categories. The interesting question is whether the incidence of job loss has changed systematically over the ten year period, 1982-1991. For this reason, the analysis in this sub-section focuses only on the 1982-83 and 1990-91 periods.

Table 4 contains estimates of weighted least squares (WLS) regressions of the log of the probability of job loss in a cell on the characteristics defining the cell. The weights used for each cell are estimates of the variance of the log probability of job loss for that cell. The log of the probability of job loss in equation 3.3 is a linear combination of the logs of its component probabilities. Specifically,

$$(3.4) \quad \ln(D_{ijkrt}) \\ = \ln\left(\Pr(T_{ikrt}=j|D_{it}=1)\right) + \ln\left(\Pr(D_{ikrt}=1)\right) - \ln\left(\Pr(T_{ikrt}=j)\right).$$

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

The first column of table 4 contains estimates of a model of the log probability of job loss with main effects for sex, age category, education level, tenure category, and time period. These results demonstrate the overwhelming negative relationship the probability of job loss and tenure. Workers with more than fifteen years tenure have a probability of job loss that is less than 25 percent of the rate of job loss of workers in their first year on the job. They also corroborate the earlier finding that females have much lower rates of job loss than men (about half). Once again, this is due at least in part to lower labor force participation rates for females. With regard to education, the estimates suggest that workers with at least some college education have a lower rate of job loss than do workers with less education. The rate of job loss for college graduates is only 65 percent of the rate of job loss of otherwise-equivalent high school graduates (the base group). Finally, the results indicate that the base-group rate of job loss in the 1990-91 period was not significantly different from the rate in the 1982-83 period (the base period).

The specification in the first column of table 4 contains only main effects, and, as such, it constrains the proportional effect of each factor on the probability of job loss to be independent of the other factors. Given

the findings earlier, that the differences across time periods were more likely to be important for some subgroups than others, it is important to investigate the relevant interactions. The second column of table 4 repeats to previous WLS regression including additionally a set of interaction terms between education and time period. The findings are rather interesting, particularly with regard to a comparison of high school and college graduates over time.

College graduates had an average rate of job loss in the 1982-83 period that was about 40 percent lower than the rate of job loss for otherwise-equivalent high school graduates. That gap shrunk to 30 percent in the 1990-91 period. Another way of expressing this result is that the rate of job loss for high school graduates (the base group) was not significantly different in 1990-91 than it was in 1982-83. However, the rate of job loss for college graduates in 1990-91 was about 18 percent higher than it was in 1982-83.

The third column of table 4 contains estimates of a model that includes age-year interactions along with the main effects and the education-year interactions. I present these estimates because of the finding in table 1 that older workers had higher rates of job loss in the 1990-91 period than in the 1982-83 period. The multivariate analysis strongly supports this result. Workers between 40 and 60 had a significantly higher probability of job loss in 1990-91 than otherwise-equivalent workers in 1982-83. The point estimates suggest that turnover rates among older workers were more than 25 percent higher in the recent period than in the earlier period.

Finally, the fourth column of table 4 contains estimates of a model that includes tenure-year interactions along with the main effects, the education-year interactions, and the age-year interactions. The point estimates of the tenure-year interactions are not presented, but none of them

are significantly different from zero. A test of the null hypothesis that the tenure-year interactions are all zero cannot be rejected (p -value=0.368). None of the substantive results alluded to above are altered by the inclusion of the tenure-year interactions.

5. Remarks on Changes in the Incidence of Job Loss

What can be concluded from the analysis in preceding sections? First, there is clear evidence that more skilled workers (those who are older and with more education) were relatively more likely to lose their jobs in the most recent period than in the earlier years. Second, while the probability of job loss is extremely highly negatively correlated with tenure, there is absolutely no evidence that workers with more tenure are more vulnerable to job loss in the most recent period than in the earlier recession.

This pattern of results raises as many questions as it answers. The public perception that the incidence of job loss is different in recent recession than in at least the immediately preceding recession has some basis in fact. To the extent that older workers and workers with more education have relatively more influence on the public mood, it would not be surprising for the public to be more "unhappy" with this recession than the last. This is consistent with weighted counts of the 1984 and 1992 DWS which show that about 14 percent more jobs were lost in the 1990-91 period than the 1982-83 period (8.6 million versus 7.5 million). On the other hand, the unemployment rate was much higher (9.5 percent versus 6.0 percent) and the rate of employment growth was much lower (-1.0 percent versus 0.0 percent) in the 1982-83 period than in the 1990-91 period (Economic Report of the President, 1992, p.344).

The finding that job loss in the 1982-83 period was relatively concentrated in manufacturing while the job loss in the 1990-91 period was

relatively concentrated in a collection of service industries suggests that adjustments have been concentrated in different sectors in the two periods. Employment in goods-producing industries fell by about 9 percent between 1978 and 1983 and by only 3.5 percent between 1987 and 1991 (Economic Report of the President, 1992, p.344). Employment in service producing industries grew by about 9.5 percent between 1978 and 1983 and by 9.1 percent between 1987 and 1991 (Economic Report of the President, 1992, p.345). Clearly, the adjustment in the goods producing sector was much more severe in the earlier period, and, given the skill composition of employment by industry, its effects were disproportionately borne by less-skilled workers.

The finding that the rate of job loss by tenure category was no different in the two recessions suggest that plant shutdowns are not contributing more to job loss than in the past. Some direct evidence is available from the DWS that is consistent with this. The DWS asks displaced workers the cause of their job loss. Plant closings accounted for 30.7 percent of job loss in 1982-83, approximately 34 percent from 1984-89, and 29.4 percent in 1990-91. There is no significant difference between the two recessions in the fraction of job loss that was due to plant closings ($p\text{-value} > 0.3$). It is interesting to note that the fraction of job loss due to plant closings is significantly higher in the periods of expansion than in either recession. However, because total job loss is higher in the recessionary periods, the absolute number of jobs lost to plant closings is higher in the recessionary periods than in the expansion.

6. The Costs of Job Loss I: Post-Displacement Labor Force Status

Earlier work using the Displaced Workers Surveys concludes that displaced workers, on average suffer earnings losses (e.g., Podgursky and Swaim, 1987). These losses come in the form of both non-employment and lower

wages upon reemployment. In this section, I examine the labor force status at the survey date of over 420,000 individuals between the ages of 20 and 64 who were part of one of the five January CPS's containing the Displaced Worker Surveys from 1984 to 1992, whether they were displaced or not. Approximately 20,000 of these workers reported being displaced within two years of one of the surveys from 1984 to 1992.¹³ I use these data to examine how displacement is related to labor force status by comparing the distribution of labor force states among workers who reported being displaced to the distribution of labor force states among workers who did not report being displaced.¹⁴

Comparing the probabilities of displaced workers being in the various labor force states with the same probabilities for individuals who were not displaced provides a useful basis for judging the effects of displacement. However, there is an important difficulty with this approach. Workers who were displaced had to have been employed at some point within the last two years while those individuals who were not displaced may be consistently out of the labor force. This will tend to bias the comparison toward finding a smaller adverse employment effects of displacement and the results must be interpreted with this in mind.

Table 5 contains tabulations of labor force status at the time of the Displaced Workers Surveys for all individuals broken down by sex, survey

¹³I delete a small number of workers who report being displaced in the month of the survey because they have had little time to find a new job.

¹⁴Another outcome that would be interesting to investigate is time spent unemployed after being displaced. Unfortunately, a change in the relevant question asked in the DWS between the 1986 and 1988 surveys makes a meaningful comparison impossible. In the 1984 and 1986 DWS's, displaced workers were asked how much time they had spent unemployed since being displaced. In the 1988, 1990, and 1992 DWS's, displaced workers were asked how much time they were unemployed until they started another job.

year, and whether the individual reported being displaced within two years of the survey date. Notwithstanding the potential bias noted above, I find dramatically lower employment probabilities in every year for displaced males compared with non-displaced males. The difference in employment probabilities (displaced - non-displaced) for females is also negative, though by a smaller amount. Unemployment probabilities for both sexes in all years are almost an order of magnitude higher for displaced workers than for non-displaced workers. Non-displaced workers of both sexes are more likely to be out of the labor force than displaced workers, and this is almost certainly due to the bias induced by not controlling for past employment among the non-displaced workers.

The tabulations in table 5 also clearly show that the probability of employment (controlling for displacement status) for both males and females is significantly higher in the survey years following periods of expansion (1986, 1988, 1990) than in the survey years following recessionary periods (1984, 1992). Not surprisingly, the probability of unemployment moves inversely with the probability of employment. The probability of being out of the labor force is rather small for men, and it does not vary much over time. Displaced women are much more likely to be out of the labor force than men, but this probability is falling over time. This is consistent with an increasing commitment of women to the labor force.

A comparison of the displacement - non-displacement difference in the probabilities of employment and unemployment at the 1984 and 1992 interview dates shows that males' penalty for displacement in this dimension is larger in 1992 than it was in 1984. The male 1992-1984 difference in the displacement - non-displacement difference in employment probabilities is -0.026 (p-value = 0.007). The male 1992-1984 difference in the displacement - non-displacement difference in unemployment probabilities precisely offsets

this at 0.026 (p-value < 0.001). Thus, displacement is estimated to result in lower employment probabilities and higher unemployment probabilities for males in the recent recession relative to the 1982-83 period. None of these differences in differences are significantly different from zero for females so that the penalty for job loss is no different for females in the recent recession relative to the 1982-83 period.¹⁵

While a multi-nomial logit specification would normally be appropriate to analyze labor force states because of the unordered discrete outcomes being considered (employment, unemployment, out of the labor force), my preliminary analyses suggests that the relevant variation in the data is summarized efficiently by a simple logit on the probability of employment. Because these estimates are easier to interpret than multi-nomial estimates, the remainder of the analysis in this section focuses on the probability of employment using the simple logit model.

Table 6 contains estimates of simple logit models of the probability of employment separately for males and females. The sample includes all individuals between the ages of 20 and 64, whether displaced or not. The model includes main effects for race, education, age, year along with interactions of these variables with whether the individual suffered a job loss. By fully interacting the worker characteristics with the job loss variable, I get the same fit as if I estimated separate logit models for displaced and non-displaced workers. The advantage to the combined estimation is that estimates are produced directly of the difference between displaced and non-displaced workers in the relationship of the probability of employment with each of these characteristics. The "main effect" columns

¹⁵In fact, the probability of employment for displaced women is lower in 1991-92 than in 1982-83, but the decline is smaller than the decline in employment probabilities for non-displaced women over the same period.

contain the estimates of the relationship for non-displaced workers of worker characteristics with the probability of employment. the "interact w/loss" columns contain estimates of the difference between displaced and non-displaced workers in these relationships. Division of the appropriate logit coefficient by five results in a good estimate of the effect of the variable on the probability of employment.¹⁶

The discussion will focus on the interaction estimates because they represent the relevant difference in employment probabilities between displaced and non-displaced workers. The multivariate analysis show a somewhat different picture of the effect of displacement on the probability of employment than was evident in the simple tabulations in table 5. Relative to non-displaced workers, the baseline probability of employment is substantially lower for individuals who were displaced in the year prior to the survey date. This difference is about -25 percentage points for men and about 12 percentage points for women. However, the apparent penalty for displacement is virtually zero for workers who were displaced in the second year prior to the survey date. The difference is reduced to about -4 percentage points for men and is actually positive (though insignificant) for women.

This is strongly consistent with workers who were displaced more than a year in the past having had more time to find a new job than recently displaced workers. This finding is unlikely to be due to the observation that the probability of an individual recalling a job loss declines with the time since displacement. It is natural to expect individuals who become employed to be less likely to recall the job loss than individuals who remain

¹⁶The derivative of the probability of full-time employment in the logit model is $\beta P(1-P)$. The sample average value of P is approximately 0.75 so that $P(1-P)$ is approximately 0.20.

not-employed. This would bias the results toward finding larger negative employment effects of displacement as the time since job loss increases.

Nonwhites and particularly nonwhite females suffer larger negative employment effects of job loss than do whites. The difference is relatively small for men (about 2 percentage points), but is larger (about 12 percentage points) for women.

The data clearly show that individuals with more education are more likely to be working. The interactions with job loss indicate that, relative to displaced workers with 12 years of education, displaced workers with other levels of education have higher probabilities of employment. The finding for the lowest education category is likely due to the selection induced by the fact that displaced workers had to be working at some point in the past. The finding for the higher education categories may reflect a greater ability by highly educated workers to find jobs after displacement. Adding the main effects for education with their respective interactions with job loss verifies that the probability of re-employment for displaced workers is monotonically increasing in education.

One of the goals of this study is to investigate how the costs of job loss have changed over time. The estimates of the main year effects and their interactions with job loss show that the displacement penalty in employment probabilities is significantly lower in the expansionary period (the 1986, 1988, 1990 surveys) than in the two recessions. This difference is large, averaging about six percentage points for men and about four percentage points for women. There is no significant difference in the displacement penalty in reemployment probabilities between the 1982-83 downturn and the most recent recessionary period.

Given the findings that older and more educated workers are more likely to suffer a job loss in the 1990-91 period than in the 1982-83 period, it is

interesting to investigate whether there are changes in the distribution of post-displacement labor force states over time. I investigated this by re-estimating the logit model of the probability of employment separately for each education category and separately for each age group. These estimates, while not presented here, show no evidence of any systematic change in the displacement penalty in employment probabilities other than the cyclical effects noted above. Specifically, there is no evidence that the displacement - non-displacement difference in employment probabilities for older or more educated workers has gotten larger (negatively).

Overall, there is clear evidence that workers who were displaced within one year of the survey date are less likely to be employed at the survey date than either workers who were not displaced or were displaced more than one year before the survey date. I find 1) that nonwhite females suffer a larger displacement penalty in employment probabilities than do white females. The race difference is smaller for males. I further find that the displacement penalty in employment probabilities is inversely related to education level. There is no evidence of secular changes displacement non-displacement difference in employment probabilities, even for particular age and education groups.

7. The Costs of Job Loss II: Part-Time Employment among Displaced Workers

While most workers who are displaced have found new employment, it may be that the quality of the job found is lower than the quality of the job lost. Of course, some workers may find better jobs, but the earlier literature on the consequences of job loss suggest that displaced workers, on average, find jobs that are inferior to the jobs they lost (e.g., Podgursky and Swaim, 1987). I consider the obvious job attribute, wages, in the next

section. Another important dimension of job quality is hours worked.¹⁷ The Displaced Worker Surveys do not have detailed information on hours on the jobs workers were displaced from, but they do ask if the job was part time (less than 35 hours per week). Hours on the job held at the survey date are available, and I use this information to determine if the displaced workers are currently working full time or part time. Table 7 contains calculations of the fraction of displaced workers who are currently working full time by year and by whether the worker was displaced from a full-time or a part-time job.

A pair of interesting facts are apparent from examination of this table. First, among those displaced workers who were working at the survey date, those displaced from full-time jobs were substantially more likely to be working on a full-time job at the survey date than were those who were displaced from part-time jobs. This may reflect persistent individual differences in labor supply preferences. Second, workers displaced in the two recessionary periods who were subsequently employed at the survey date had a significantly lower probability of working in full-time jobs at the survey date. This is true regardless of whether the lost job was full time or part time.

A weakness of this analysis of variation in the probability of full-time work among displaced workers is that there is no comparison group of workers who were not displaced. As with the analysis of employment probabilities, a natural comparison group consists of the non-displaced workers in the January CPS's that contain the Displaced Worker Supplements. I constructed a sample of currently employed workers from the five January

¹⁷See Altonji and Paxson (1988) for an interesting analysis of hours constraints, wages, and labor supply.

CPS's with the DWS's, and I generated a measure of full-time work (≥ 35 hours per week) based on reported hours in the week preceding the interview. Table 8 contains tabulations by year of the fraction of re-employed displaced workers and non-displaced workers who are working full-time at each survey date by this definition. These tabulations clearly show that re-employed displaced workers have a lower probability of working full-time than do workers who were not displaced.

This finding is not due to part-time workers having a disproportionately high probability of being displaced. Approximately 88 percent of displaced workers at each DWS survey date were displaced from full-time jobs. Note further in table 8 that the lower probability realized by displaced workers in recessionary periods of being re-employed full-time is *not* mirrored among workers who were not displaced. The probability of full-time employment among non-displaced workers is not significantly lower in the two recessionary periods than in the periods of expansion.

These results suggest that an important cost of job displacement is that workers are less likely to find or take full-time jobs after displacement and that this problem is larger in recessions than in expansions. Additionally, since the wage rate on part-time jobs is substantially less than the wage rate on full-time jobs, other things equal, the increased incidence of part-time employment among displaced workers in the recessionary periods is an important component of the costs of job loss.

Table 9 contains estimates of a logit model of the probability of full-time employment among those displaced workers who are employed at the survey date. These estimates are derived using the sample of displaced workers who were re-employed at the survey date, and the results must be interpreted conditional on re-employment. This model includes a measure of whether or not the lost job was part-time or full-time in order to account

for gross and persistent differences in hours preferences. As shown in table 7, workers who lost part-time jobs were less likely to be employed in full-time jobs at the survey date. Division of the appropriate logit coefficient by five results in a good estimate of the effect of the variable on the probability of full-time employment.¹⁸ Thus, workers who lost part-time jobs were about 25 percentage points less likely to hold full-time jobs at the survey date than were workers who lost full-time jobs.

The logit estimates also show that females are about 11 percentage points less likely than otherwise-equivalent men to hold a full-time job after displacement. Note that this result is not due to the fact that women are more likely to hold part-time jobs generally because the 11 percentage point estimate is derived controlling for part-time status of the lost job. Nonwhites are about 5 percentage points less likely to hold a full-time job after displacement than otherwise-equivalent whites.

There is a strong positive correlation between the probability of full-time employment after displacement and education. For example, college graduates are about 10 percentage points more likely than high school graduates to be employed full time after displacement. Workers with less than a high school education are about 7 percentage points less likely than high school graduates to be employed full time after displacement.

Older workers (aged 55 to 64) are significantly less likely than younger workers to be employed full time after displacement. For example, employed workers 60-64 who are displaced are about 17 percentage points less likely than workers twenty years younger to be working full time after displacement. This suggests that older displaced workers may "partially"

¹⁸The derivative of the probability of full-time employment in the logit model is $\beta P(1-P)$. The sample average value of P is approximately 0.75 so that $P(1-P)$ is approximately 0.20.

retire. These workers may have a pension available to them on displacement.

With regard to calendar year, workers in the periods of expansion (1986, 1988, 1990) are more likely (about 5 percentage points) to be employed full time than are workers in either recessionary period (1984 and 1992). There seems to be no significant difference between 1984 and 1992, suggesting that full-time employment was not harder to find after displacement in the recent recession relative to the earlier recession. Estimates of a model that allows for different education and age effects by year (not shown), provide no evidence (other than the cyclical factors) that the relationship of the probability of full-time employment with age and education varies systematically over time. In particular, the two recessionary periods are not significantly different.

The final variables included in the model crudely measure the length of time since displacement. The base group consists of those workers who were displaced in January of the survey year (immediately prior to the survey). It is interesting that the probability of full-time employment rises with time since displacement. This is consistent with displaced workers accepting part-time employment relatively quickly while continuing to search for a full-time job. It is also consistent with the role that time since displacement played in determining the probability of employment. The passage of time provides an opportunity not only to find a job but also to find a good (i.e., full-time) job.

Overall, the results in this section demonstrate that re-employed displaced workers are less likely to be employed full-time after displacement than they were before displacement and that there are important differences across re-employed displaced workers in their probability of full-time employment. Most of these differences are not likely to be due to differences in desired hours as the estimates control for the full-time /

part-time status of the lost job. Workers displaced during the recessionary periods have a lower probability of full-time employment than workers displaced during the expansion, but there is no evidence that the relationship between the probability of full-time employment and worker characteristics varies systematically over time. Finally, the reduction in the probability of full-time employment declines with time since displacement as displaced workers have sufficient time to find full-time jobs.

8. The Costs of Job Loss III: Earnings Losses of Re-Employed Workers

The final piece of the analysis of the costs of job loss relates to the wage changes associated with job loss. Earlier work using the DWS documents that displaced workers suffer earnings losses on average (Podgursky and Swaim, 1987; Kletzer, 1989; de la Rica, 1992). The DWS contains information only on usual weekly earnings on the lost job. No hours information is collected other than the full-time / part-time distinction so that it is not possible to compute an hourly wage rate. The same information is available for the current job along with usual hours so that an hourly wage can be calculated for the current job. However because I am interested in wage *changes*, I am forced to use the weekly earnings measure controlling for full-time / part-time status on both the old and new jobs.

Table 10 contains weighted average log real wage changes (changes in reported weekly earnings) by year for displaced workers who are employed as of the date of the survey. The log real wage change is computed as the difference between reported weekly earnings as of the survey date and the reported weekly earnings on the lost job as of the date of displacement. Wages are deflated by the 1982-84=100 consumer price index (CPI). The CPI in the reported year of displacement is used to deflate earnings on the old job. Because the DWS is conducted in January, the CPI in the year *prior* to the

survey date is used to deflate current earnings in this table.

The first column of table 10 contains these averages for all employed displaced workers and they show an average loss in weekly earnings of about 15 percent. However, the analysis in the preceding sub-section demonstrated that re-employed job losers are less likely to hold a full-time job after displacement than they were before displacement. Thus, the remaining four columns of table 10 contain the mean log wage changes for the four combinations of full-time / part-time status on the old and new jobs.

Workers who lost a full-time job but found a new full-time job suffered wage losses that average about 9 percent. Workers who lost a part-time job and found a new part-time job generally suffered wage losses insignificantly different from zero on average. However, there is more scope for hours variation within part-time jobs (1-34 hours per week), and these comparisons suffer from unmeasured hours variation. The two other categories show the not-surprising result that displaced workers who move from full-time to part-time suffer huge losses in weekly earnings while those who move from part-time to full-time enjoy huge gains (of approximately equal magnitude) in weekly earnings. These reflect changes in labor supply and are not directly informative about the costs of displacement other than observing the fact that among displaced workers full-time to part-time transitions are over three times more common than part-time to full time transitions.

Table 11 contains estimates of log wage change regressions that control explicitly for the full-time / part-time status of both the old and the new jobs. The R-squared is 0.168, which is quite high for a wage-change regression, but most of the explanatory power is coming from the part-time-status variables which represent labor supply factors more than

wage-change factors.¹⁹

The results do suggest that older workers have lower wage changes than younger workers. Each ten years of age represents about 3 percent lower wage growth. Tenure on the previous job is also significantly related to wage changes with more senior workers having more negative wage changes. Each year of previous tenure is related to about a 1 percent lower wage change. This may be related to unobservable worker ability that is correlated with job tenure so that high tenure workers who are displaced are of relatively high ability.²⁰ It is important to keep in mind that there is overwhelming evidence that age-log earnings profiles are concave so that wage changes are decreasing in age for workers generally. Similarly, there is evidence that tenure-log earnings profiles are concave so that wage changes are decreasing in tenure for workers generally.²¹

The estimates suggest that workers who were displaced in either of the two years prior to the survey have about a 15 percentage point lower wage change than do those workers who were displaced in the month of the survey (the base group). This is likely due to selection in that, of the workers who were displaced in the month of the survey, only the most able were likely to have found a new job before the survey was taken. Additionally, any reasonable job search model with a declining reservation wage would have the property that searches that end quickly will yield higher wages.

¹⁹The R-squared for the regression without the part-time status variables is about 0.02.

²⁰Abraham and Farber (1987) make this point and present some consistent evidence from the Panel Study of Income Dynamics. Kletzer (1989) examines the relationship between previous tenure and wage change using some of the early Displaced Worker Surveys.

²¹See Topel (1991) for an investigation of the returns to tenure.

There is no significant difference by calendar year in wage changes among displaced workers. Additional regressions, not presented here, that include interactions of year with education and age do not find any significant interactions. Thus, workers who are displaced in recessions do not seem to suffer larger wage declines (or smaller wage increases) than do workers who are displaced in expansions. One caveat to this conclusion is that workers displaced in recessions are more likely to be re-employed in part-time jobs which pay substantially lower wages generally.²²

An important weakness of this analysis is that wage change for displaced workers ought to be benchmarked against wage change for workers who were not displaced (or against the population as a whole since displaced workers represent a relatively small fraction of the workforce). Only by differencing the wage changes of non-displaced workers with the wage changes of displaced workers can a proper measure of the wage penalty associated with job loss be made. This is related to the point made above that simply finding smaller wage changes for older and/or higher tenure workers is not sufficient to conclude that these workers suffer disproportionately from job loss. The wage changes of displaced workers must be compared to the wage changes of similar workers who were not displaced (a control group).

I generate a control group using a five-percent random sample from the merged outgoing rotation group (OGRG) files of the CPS for the two calendar years prior to each DWS together with all non-displaced workers from the outgoing rotation groups of the January CPS's containing the DWS's.²³ The

²²Remember that the estimates here cannot be used to draw this conclusion because the part-time measures are dominated by hours differences and greatly exaggerate and wage penalty to part-time work.

²³The OGRG data are used by de la Rica (1992) as a control group for earnings losses suffered by workers who report being displaced in the 1986 DWS. Jacobson, Lalonde, and Sullivan (1991) also use a control group to estimate

OGRG contain information on weekly earnings and hours for a random sample of the workforce each year. The two year period prior to each DWS covers the entire period of displacement for workers, and, as such, provides base-period weekly earnings for workers who were not displaced during this period.²⁴ Weekly earnings as of the "current" period (the date of each DWS) for non-displaced workers are derived from the outgoing rotation groups of the relevant January CPS.

Since these data are neither longitudinal nor retrospective, I cannot compute wage changes for individual workers. However, I can measure whether or not the observation represents a base-period or a current period wage. This allows me to estimate the average changes in weekly earnings of non-displaced workers.²⁵ My sample includes full-time workers (≥ 35 hours per week) between the ages of 20 and 64. A comparable sample of displaced workers is derived using the workers who report having been displaced within two years of each DWS survey date.

These data are used to estimate an earning function of the form

$$(8.1) \quad \ln E_j = \gamma_0 + X_j\beta + \gamma_1 D_j + \gamma_2 C_j + \gamma_3 D_j C_j + \epsilon_j,$$

where

E_j = real earnings for cell j ,

X_j = dummy variables for main effects for worker characteristics

D_j = a dummy variable for job displacement,

earnings losses of displaced workers.

²⁴In fact, workers who are or will be displaced are included in the merged OGRG files, but, since only a small fraction of workers are displaced in any year (less than 5 percent, I ignore this slight distortion.

²⁵The OGRG files contain enough information to compute hourly earnings, but, because this DWS's do not contain this information, the analysis here uses on weekly earnings for full-time workers.

C_j = a dummy variable for current period, and

ε_j = an error term.

The vector β and the γ 's are parameters to be estimated. The X vector contains controls for sex, race, four education categories, and nine age categories.

This specification allows separate intercepts for four groups: 1) initial wages of non-displaced workers (γ_0), 2) initial wages of displaced workers ($\gamma_0 + \gamma_1$), 3) current wages of non-displaced workers ($\gamma_0 + \gamma_2$), and 4) current wages of displaced workers ($\gamma_0 + \gamma_1 + \gamma_2 + \gamma_3$). The usual estimate of the wage effect of displacement is the wage change of displaced workers ($\gamma_2 + \gamma_3$). The difference-in-difference estimate subtracts the "normal" wage change of non-displaced workers over the same period (γ_2) from the wage change of displaced workers. The resulting estimate is simply γ_3 . To the extent that the wages of non-displaced workers are growing, the difference-in-difference estimate will be smaller (more negative) than the simple wage change of displaced workers.

Age categories are defined as of the current date rather than the base period date in order to ensure that calculations of wage changes between the base and current period use a consistent group of workers. For example, non-displaced workers 20-24 years old in the January 1984 CPS contribute current earnings observations. The base period earnings for these workers are drawn from the OGRG files using workers 18-22 in 1982 and 19-23 in 1983.

A brief word on the deflator used for earnings in this analysis is required. The annual average CPI for each year was used to deflate base period earnings (the year of the OGRG file for non-displaced workers and the year of job loss for displaced workers). The annual average CPI was also used to deflate current earnings, but this is not strictly correct given the fact that current earnings are always computed in January. Thus, current

earnings are "over-deflated" so that wage changes for both non-displaced and displaced workers are estimated to be smaller than they ought to be by approximately one half the annual rate of inflation. A further complication is that the average rate of inflation for 1992 is not available so that the average CPI for 1991 was used to deflate January 1992 earnings. Thus, wage changes between 1990 or 1991 and 1992 are "under-deflated" so that wage changes are estimated to be larger than they ought to be by about one half the 1991 rate of inflation.

This problem of over- or under-deflation of earnings is not particularly serious for estimation of the difference-in-difference estimator of the effect of displacement on earnings (γ_3). The estimate of γ_3 is not affected by the choice of timing of the deflator, because the CPI is differenced away in log terms. However, interpretations of simple time differences in earnings (i.e., γ_2 and $\gamma_2 + \gamma_3$) are affected by the timing.

Table 12 contains ordinary least squares estimates of equation 8.1 using the data on full-time real weekly earnings. While only the estimates of the relevant differences (γ 's) are shown, all regressions also include main effects for sex, race, education category, and age category. Note that the base period is (randomly) up to two years prior to the current period. Since the average time difference between the base and current periods is one year, it is reasonable to interpret the wage changes as annual rates of change.

The estimates of γ_1 show that base-period earnings are significantly lower for displaced workers relative to non-displaced workers. This is consistent with existing evidence from other sources that displaced workers suffer earnings losses even before they lose their jobs. Most convincing on this point is the work by Jacobson, Lalonde, and Sullivan (1991) using

administrative data with workers' quarterly earnings histories.²⁶ There are at least two other potential explanations for this finding. First, the finding may result from displaced workers having less tenure than non-displaced workers. Second, displaced workers may simply be of lower quality than non-displaced workers.²⁷

The estimates of γ_2 show that real wages of non-displaced workers did not change very much on average over the sample period. However, there is considerable year-to-year variation with real wages falling by as much as five percent in some years and growing by as much as four percent in others. Some of this variation is probably due to the over- and under-deflation of current earnings. Wage changes are understated in 1984 through 1990, years that generally have negative estimated wage changes. In contrast, wage changes are overstated in 1992, the only year with a substantial positive wage change.

The estimates of the earnings change for displaced workers ($\gamma_2 + \gamma_3$) show that their wages fell by over ten percent on average. Again, these estimates are subject to the deflation problem.

The difference-in-difference estimates (γ_3) show a large and significant negative effect of displacement of about fifteen percent, and these estimates are free of the over- or under-deflation problem. Thus displaced workers suffer a substantial real wage decline relative to non-displaced workers. With regard to changes over time in the earnings loss associated with displacement, the point estimate of γ_3 does vary somewhat

²⁶Blanchflower (1991) using data from the U.K. and de la Rica (1992) using the 1986 DWS also find that displaced workers suffer earnings losses prior to displacement.

²⁷Being of lower tenure and of lower ability may be related if part of what makes a worker high-ability is stability. See Abraham and Farber (1987).

from year to year, but only the estimate for 1990 is significantly different from the other years. There seems to be no systematic pattern to movements in γ_3 .

While not presented here, the earnings function was re-estimated separately for each of the four education categories. Estimated differences in γ_3 by education category were generally not statistically significant. The analogous exercise was carried out by age category, and differences in γ_3 by age category were generally not statistically significant. There was no consistent pattern over time in the estimated displacement effects by either education category or age category. Note specifically that older and/or more educated displaced (and full-time re-employed) workers did not suffer larger earnings losses relative to younger/ less educated displaced workers in the most recent period relative to earlier periods.

Overall, the results in the section document the substantial earnings losses (on the order of ten percentage points) suffered by displaced workers relative to non-displaced workers. These losses are found to be greater for workers displaced in the 1990-91 period. There is no evidence that the age difference in earnings losses suffered by displaced workers has grown over time or that earnings losses are falling relatively more heavily on highly educated workers.

9. Concluding Remarks

It is clear that the costs of job loss work through several channels. First, displaced workers are less likely to be employed than otherwise equivalent non-displaced workers. Second, displaced workers who are re-employed are less likely to be employed full time after displacement relative either to otherwise equivalent non-displaced workers or to their own pre-displacement hours. Third, even conditioning on re-employment on a

full-time job, displaced workers earn substantially less either than otherwise equivalent non-displaced workers or than they themselves earned prior to displacement.

There is evidence that some of the costs of displacement are temporary. The probability-of-employment penalty and the part-time-employment penalty for displacement seem to decline with time since displacement. Thus, the unconditional expected earnings penalty declines with time since displacement. However, there was no evidence that the earnings gap between full-time re-employed displaced workers and full-time employed non-displaced workers narrows.

More formally, the expected current period earnings of workers is

$$(9.1) \quad E(W) = \Pr(\text{EMP}) \cdot \left(\Pr(\text{FT}|\text{EMP}) \cdot E(W|\text{EMP}, \text{FT}) + \Pr(\text{PT}|\text{EMP}) \cdot E(W|\text{EMP}, \text{PT}) \right)$$

where

$\Pr(\text{EMP})$ = probability of re-employment,

$\Pr(\text{FT}|\text{EMP})$ = conditional probability of full-time employment,

$\Pr(\text{PT}|\text{EMP})$ = conditional probability of part-time employment,

$E(W|\text{EMP}, \text{FT})$ = conditional expectation of full-time earnings, and

$E(W|\text{EMP}, \text{PT})$ = conditional expectation of part-time earnings.

Job loss adversely affects all terms in this expectation.²⁸ Additionally, the earnings (and wage rates) of full-time workers are higher than those of part-time workers so that a shift toward part-time employment adversely affects expected earnings. The effects of displacement on the probabilities of employment and full-time employment are temporary at least in part. The full-time wage effect is likely to be persistent, at least in part because the wage loss is directly related to tenure and lost tenure is never

²⁸In fact, little evidence is presented regarding the effect of displacement on part-time earnings. The simple statistics in table 10 suggest that displacement has only a small effect on this value.

recovered.

With regard to the perception that the nature and consequences of job loss are different than they used to be, this seems generally correct. It is indeed the case that older and more educated workers were more vulnerable to job loss in the 1990-91 period than they were in the 1982-83 period. Additionally, job loss was relatively more common in important service industries and relatively less common in manufacturing in 1990-91 compared with 1982-83. There is also some weaker evidence that the costs to workers who lose their jobs has increased. Specifically, the wage loss associated with displacement for full-time re-employed workers was substantially larger for workers displaced in 1990-91 than it was for workers displaced earlier. This is an aggregate phenomenon, not confined to particular education or age groups.

While the probabilities of post-displacement re-employment and full-time re-employment are strongly pro-cyclical, there is no evidence for any secular change in these probabilities either in the aggregate or for particular groups.

Overall, the costs to displaced workers of job loss are substantial and come in several forms. However, the public perception that the current sluggish economy is worse than earlier downturns may reflect more who has lost jobs recently rather than either increased overall job loss or increased costs to those who are losing jobs.

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Table 1

Fraction Reporting Job Displacement in Two-Year Interval

Population Based

Males

Age Cat	Years				
	1982-83	1984-85	1986-87	1988-89	1990-91
20-24	0.101	0.0814	0.0651	0.0578	0.0973
25-29	0.108	0.0768	0.0693	0.0627	0.0990
30-34	0.0904	0.0688	0.0642	0.0562	0.0782
35-39	0.0777	0.0640	0.0578	0.0525	0.0784
40-44	0.0556	0.0579	0.0510	0.0472	0.0693
45-49	0.0614	0.0436	0.0487	0.0347	0.0625
50-54	0.0486	0.0543	0.0411	0.0312	0.0635
55-59	0.0467	0.0407	0.0349	0.0324	0.0588
60-64	0.0363	0.0272	0.0250	0.0219	0.0198

Females

Age Cat	Years				
	1982-83	1984-85	1986-87	1988-89	1990-91
20-24	0.0588	0.0512	0.0439	0.0416	0.0542
25-29	0.0495	0.0431	0.0437	0.0360	0.0550
30-34	0.0476	0.0391	0.0363	0.0378	0.0555
35-39	0.0402	0.0355	0.0357	0.0372	0.0426
40-44	0.0422	0.0294	0.0365	0.0376	0.0443
45-49	0.0339	0.0344	0.0278	0.0320	0.0472
50-54	0.0320	0.0248	0.0270	0.0241	0.0357
55-59	0.0268	0.0218	0.0236	0.0230	0.0281
60-64	0.0182	0.0210	0.0148	0.0142	0.0198

Source: From authors tabulations of Displaced Workers Supplements to January Current Population Surveys in 1984, 1986, 1988, 1990, and 1992.

Table 2
Logit Estimates of Probability of Job Loss - DWS

Variable	Educ < 12 age		Educ = 12 age		Educ 13-15 age		Educ >= 16 age	
	20-34	35-59	20-34	35-59	20-34	35-59	20-34	35-59
Constant	-1.80 (.0612)	-2.25 (.0683)	-1.95 (.0365)	-2.47 (.0473)	-2.60 (.0528)	-2.56 (.0678)	-3.04 (.0927)	-3.11 (.0734)
female	-.834 (.0536)	-.669 (.0494)	-.613 (.0295)	-.499 (.0335)	-.457 (.0411)	-.492 (.0472)	-.339 (.0558)	-.448 (.0547)
nonwhite	-.0211 (.0649)	.0058 (.0602)	.105 (.0396)	-.0216 (.0519)	.0772 (.0572)	-.0597 (.0740)	-.0649 (.0896)	-.104 (.0904)
1984-85	-.218 (.0722)	-.150 (.0688)	-.229 (.0422)	-.121 (.0529)	-.347 (.0639)	-.195 (.0812)	-.206 (.0842)	-.0699 (.0883)
1986-87	-.453 (.0778)	-.293 (.0731)	-.382 (.0446)	-.194 (.0536)	-.274 (.0636)	-.0735 (.0781)	-.325 (.0873)	-.122 (.0878)
1988-89	-.534 (.0802)	-.356 (.0767)	-.481 (.0467)	-.229 (.0537)	-.452 (.0670)	-.268 (.0798)	-.352 (.0894)	-.141 (.0864)
1990-91	-.175 (.0764)	.0259 (.0709)	.0066 (.0428)	.127 (.0499)	-.0335 (.0580)	.177 (.0703)	.124 (.0800)	.326 (.0788)
age25-29	-.115 (.0764)	---	-.136 (.0344)	---	-.336 (.0490)	---	.206 (.0851)	---
age30-34	-.0621 (.0619)	---	-.280 (.0360)	---	-.205 (.0502)	---	.140 (.0845)	---
age40-44	---	-.0726 (.0731)	---	-.141 (.0451)	---	-.0929 (.0597)	---	-.105 (.0658)
age45-49	---	-.226 (.0745)	---	-.333 (.0494)	---	-.144 (.0677)	---	-.172 (.0755)
age50-54	---	-.414 (.0753)	---	-.450 (.0536)	---	-.230 (.0769)	---	-.313 (.0871)
age55-59	---	-.543 (.0746)	---	-.550 (.0560)	---	-.516 (.0904)	---	-.489 (.100)
Log L	-5915.5	-7272.7	-18514.2	-15187.1	-9805.5	-7726.8	-5756.9	-6842.0
N	22187	37306	73119	83489	46624	40102	35175	49037
Mean dep var	.0780	.0502	.0717	.0453	.0550	.0489	.0391	.0318

Note: The numbers in parentheses are asymptotic standard errors. Data are for workers aged 20-64 from the Displaced Worker Surveys from 1984 through 1992. The base group consists of white males in 1982-83 who, at the interview date were 20-24 years old for the younger groups and 35-39 years old for the older groups. The dependent variable equals one if the individual suffered a job loss in the relevant two-year period.

Table 3
Incidence of Job Loss by Industry

Employment Shares by Industry (January CPS)					
Industry	1984	1986	1988	1990	1992
manufacturing	.2010	.1982	.1852	.1836	.1745
trans,comm,pu	.0728	.0733	.0751	.0728	.0736
trade	.1891	.1858	.1875	.1871	.1871
finsre	.0651	.0686	.0714	.0693	.0693
nonprof service	.0976	.1027	.1085	.1104	.0951
prof. service	.2181	.2189	.2251	.2251	.2607
other	.1563	.1525	.1471	.1491	.1396

Share of Job Loss by Industry
(January CPS -DWS)

Industry	1982-83	1984-85	1986-87	1988-89	1990-91
manufacturing	.3663	.3561	.2457	.2525	.2485
trans,comm,pu	.0692	.0817	.0647	.0609	.0569
trade	.1692	.1615	.2345	.2330	.2194
finsre	.0242	.0290	.0488	.0566	.0626
nonprof service	.1051	.1157	.1103	.1146	.1098
prof. service	.0602	.0636	.0933	.0977	.1075
other	.2057	.1924	.2027	.1848	.1954

Probability of Job Loss by Industry
(Equation 3.1)

Industry	1982-83	1984-85	1986-87	1988-89	1990-91
manufacturing	.1053	.0859	.0578	.0549	.0840
trans,comm,pu	.0549	.0533	.0376	.0334	.0456
trade	.0517	.0415	.0545	.0497	.0692
finsre	.0215	.0202	.0298	.0326	.0533
nonprof service	.0622	.0539	.0443	.0414	.0681
prof. service	.0160	.0139	.0181	.0173	.0243
other	.0761	.0603	.0601	.0495	.0826

Note: Based on tabulations of Displaced Worker Surveys. See text for details.

Table 4
WLS Estimates of Log Probability of Job Loss
Sex, Age, Tenure, Education, Year Cells
Displaced Workers Supplements and January Mobility Supplements to CPS
1982-83 and 1990-91

Variable	Weighted Mean	(1)	(2)	(3)	(4)
Constant	1.0	-1.86 (.0326)	-1.84 (.0347)	-1.82 (.0396)	-1.83 (.0408)
1990-91	.504	.0300 (.0229)	-.0004 (.0345)	-.0505 (.0557)	-.0361 (.0585)
Education < 12	.219	.0280 (.0322)	.0388 (.0424)	.0525 (.0425)	.0549 (.0425)
Educ 13-15	.241	-.143 (.0291)	-.165 (.0424)	-.170 (.0423)	-.169 (.0423)
Education >=16	.203	-.426 (.0335)	-.514 (.0491)	-.512 (.0492)	-.511 (.0492)
Education < 12 *1990-91	.0987	---	-.0323 (.0644)	-.0541 (.0646)	-.0581 (.0649)
Educ 13-15 *1990-91	.130	---	.0432 (.0582)	.0464 (.0581)	.0458 (.0581)
Education >=16 *1990-91	.108	---	.164 (.0668)	.154 (.0670)	.155 (.0671)
Age 25-29	.128	.0973 (.0367)	.102 (.0366)	.115 (.0493)	.110 (.0496)
Age 30-34	.132	.130 (.0390)	.135 (.0389)	.132 (.0529)	.120 (.0538)
Age 35-39	.126	.0992 (.0421)	.105 (.0420)	.0966 (.0587)	.0799 (.0597)
Age 40-44	.118	.111 (.0460)	.112 (.0459)	.0207 (.0666)	-.00544 (.0681)
Age 45-49	.111	.188 (.0503)	.189 (.0502)	.0950 (.0721)	.0604 (.0740)
Age 50-54	.0986	.156 (.0556)	.160 (.0555)	.0589 (.0771)	.0225 (.0793)
Age 55-59	.0917	.0805 (.0595)	.0837 (.0594)	-.0253 (.0796)	-.0614 (.0825)
Age 60-64	.0772	-.177 (.0688)	-.172 (.0686)	-.1893 (.0902)	-.232 (.0930)

(continued on next page)

Table 4 (continued)
WLS Estimates of Log Probability of Job Loss

Variable	Weighted Mean	(1)	(2)	(3)	(4)
Age 25-29 *1990-91	.0619	---	---	-.0247 (.0724)	-.0134 (.0734)
Age 30-34 *1990-91	.0652	---	---	.0134 (.0756)	.0358 (.0778)
Age 35-39 *1990-91	.0647	---	---	.0248 (.0812)	.0561 (.0840)
Age 40-44 *1990-91	.0645	---	---	.168 (.0886)	.217 (.0924)
Age 45-49 *1990-91	.0601	---	---	.175 (.0959)	.236 (.101)
Age 50-54 *1990-91	.0509	---	---	.198 (.105)	.263 (.111)
Age 55-59 1990-91	.0453	---	---	.225 (.111)	.293 (.119)
Age 60-64 1990-91	.0366	---	---	.0340 (.130)	.111 (.137)
1 ≤ Tenure < 2	.146	-.400 (.0355)	-.399 (.0353)	-.399 (.0353)	-.403 (.0508)
2 ≤ Tenure < 3	.130	-.343 (.0392)	-.344 (.0391)	-.343 (.0390)	-.271 (.0544)
3 ≤ Tenure < 5	.142	-.636 (.0374)	-.637 (.0373)	-.638 (.0372)	-.627 (.0522)
5 ≤ Tenure < 10	.140	-.951 (.0388)	-.956 (.0387)	-.957 (.0386)	-.935 (.0544)
10 ≤ Tenure < 15	.0821	-1.22 (.0550)	-1.22 (.0549)	-1.22 (.0548)	-1.15 (.0785)
15 ≤ Tenure < 20	.0538	-1.378 (.0751)	-1.38 (.0749)	-1.37 (.0747)	-1.26 (.104)
20 ≤ Tenure < 25	.0377	-1.45 (.0969)	-1.45 (.0966)	-1.44 (.0964)	-1.26 (.139)
Tenure ≥ 25	.0359	-1.60 (.0902)	-1.60 (.0900)	-1.59 (.0898)	-1.54 (.122)
Female	.448	-.620 (.0231)	-.668 (.0267)	-.619 (.0230)	-.617 (.0230)

(continued on next page)

Table 4 (continued)
WLS Estimates of Log Probability of Job Loss

Variable	Weighted Mean	(1)	(2)	(3)	(4)
Tenure*Year Interactions		No	No	No	Yes
R-squared		.694	.696	.700	.703
SEE		.370	.369	.368	.368

Note: The numbers in parentheses are standard errors. The sample consists of 1072 sex-education-age-tenure-year cells derived using data on individuals aged 20-64 from the 1984 and 1992 Displaced Workers Surveys. The base cell contains males 20 to 24 years old with less than one year of tenure and twelve years of education in 1982-83. The means are weighted by the square root of the inverse variance of the log probability of job loss in each cell. This is the estimation weight as described in the text. All variables are dummy variables. The weighted mean of the dependent variable is -2.84.

Table 5
Labor Force Status by Displacement, Sex, and Year
As of January of Survey Year
(weighted column percentage)

Males										
LF State	Displaced Within two Years of January?									
	1984		1986		1988		1990		1992	
	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
employed	.828	.577	.841	.599	.843	.646	.843	.665	.826	.550
unemployed	.052	.358	.043	.322	.038	.286	.039	.269	.048	.380
Not in LF	.121	.065	.117	.079	.119	.068	.118	.066	.126	.070
N	37519	3028	37901	2443	37594	2135	37763	1855	36704	2918

Females										
LF State	Displaced Within two Years of January?									
	1984		1986		1988		1990		1992	
	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
employed	.593	.486	.619	.550	.644	.590	.664	.595	.660	.531
unemployed	.037	.303	.034	.238	.030	.223	.028	.221	.032	.293
Not in LF	.370	.210	.347	.213	.326	.188	.308	.184	.308	.176
N	42527	1806	42608	1493	42391	1447	42115	1410	41522	1879

Source: From authors tabulations of Displaced Workers Supplements to January Current Population Surveys in 1984, 1986, 1988, 1990, and 1992. The CPS final sampling weights are used as weights.

Table 6

Logit Estimates of Probability of Employment

Workers from January CPS
1984, 1986, 1988, 1990, 1992

Effect of Job Loss

Variable	Males			Females		
	Weighted Mean	Main Effect	Interact w/loss	Weighted Mean	Main Effect	Interact w/loss
Constant	1.0	1.23 (.0217)	-1.31 (.0683)	1.0	.288 (.165)	-.616 (.0818)
Displaced 2nd year prior to Survey	---	---	1.12 (.0430)	---	---	.720 (.0508)
Nonwhite	.134	-.648 (.0172)	-.106 (.0606)	.152	.817 (.0716)	-.637 (.0801)
Education < 12	.176	-.719 (.0170)	.395 (.0556)	.170	.485 (.0788)	.282 (.0740)
Educ 13-15	.215	-.0699 (.0179)	.392 (.0563)	.219	-.476 (.0716)	.216 (.0639)
Education >=16	.238	.631 (.0211)	.101 (.0679)	.183	-.766 (.0916)	.203 (.0287)
Age 25-29	.150	.913 (.0239)	-.731 (.0701)	.147	.123 (.0181)	-.105 (.0856)
Age 30-34	.150	1.13 (.0208)	-1.10 (.0729)	.147	.0980 (.0181)	-.219 (.0856)
Age 35-39	.134	1.23 (.0271)	-1.18 (.0778)	.132	.235 (.0189)	-.317 (.0913)
Age 40-44	.112	1.20 (.0285)	-1.30 (.0847)	.111	.355 (.0200)	-.319 (.0965)
Age 45-49	.091	1.11 (.0294)	-1.16 (.0931)	.091	.287 (.0209)	-.418 (.106)
Age 50-54	.079	.787 (.0282)	-1.11 (.0981)	.080	.0038 (.0210)	-.559 (.120)
Age 55-59	.076	.180 (.0251)	-.724 (.104)	.079	-.405 (.0207)	-.182 (.126)
Age 60-64	.073	-.901 (.0231)	-.341 (.125)	.079	-1.15 (.0214)	-.0227 (.152)

(continued on next page)

Table 6 continued

Logit Estimates Probability of Employment

Workers from January CPS
1984, 1986, 1988, 1990, 1992

Variable	Males			Females		
	Weighted Mean	Main Effect	Interact w/loss	Weighted Mean	Main Effect	Interact w/loss
1986	.196	.0792 (.0214)	.105 (.0657)	.197	.0926 (.0152)	.219 (.0801)
1988	.201	.0554 (.0212)	.297 (.0676)	.201	.186 (.0153)	.233 (.0809)
1990	.204	.0304 (.0212)	.486 (.0710)	.204	.249 (.0153)	.211 (.0816)
1992	.211	-.134 (.0208)	.0664 (.0614)	.209	.195 (.0153)	-.0700 (.0751)
Sample Size	214767			8229		
Log Likelihood	-130301.6			-7606.6		

Note: The numbers in parentheses are asymptotic standard errors. All data are weighted by the CPS final sampling weights. The weighted means of the dependent variable are given in table 5. The base group consists of white individuals from the January 1984 CPS who are 20 to 24 years old, have twelve years of education, and who had not been displaced within two years of the survey date. The sample includes all individuals aged 20-64. 714 Individuals who reported a job loss in the survey month were deleted. This was approximately 3 percent of displaced workers and less than 0.2 percent of the total sample.

Table 7
Hours Transitions for Re-Employed Displaced Workers
by Year

Fraction Full-Time			
Year	All	FT on Old Job	PT on Old Job
1984	.729 (.0113)	.766 (.0115)	.450 (.0377)
1986	.755 (.0103)	.788 (.0104)	.485 (.0360)
1988	.770 (.0100)	.796 (.0102)	.562 (.0340)
1990	.769 (.0106)	.804 (.0108)	.474 (.0371)
1992	.724 (.0093)	.760 (.0094)	.434 (.0320)
N	8899	7882	1004

Note: The numbers in parentheses are standard deviations of the means. All averages are weighted by the CPS final sampling weights. The two sub group sample sizes do not sum to the sample size in the "all" column because of missing data on part-time / full-time status in the old job. Part-time is defined as less than 35 hours per week.

Table 8

Fraction Working Full-Time by Whether Displaced Within Two Years

by Year

January Current Population Surveys

Fraction Full-Time at Survey Date

Year	All	Not Displaced	Displaced
1984	.787 (.0017)	.790 (.0018)	.721 (.0089)
1986	.804 (.0017)	.806 (.0017)	.754 (.0095)
1988	.792 (.0017)	.794 (.0017)	.763 (.0093)
1990	.806 (.0016)	.807 (.0017)	.758 (.0097)
1992	.790 (.0016)	.794 (.0017)	.716 (.0084)
N	292671	281187	11504

Note: The numbers in parentheses are standard deviations of the means. All averages are weighted by the CPS final sampling weights. The displaced sample does not match the sample used in table 7 because it is based on the main CPS current hours information rather than the information on hours in the DWS. Full-time employment is defined as ≥ 35 hours per week. The sample consists of individuals aged 20-64 employed at the interview date.

Table 9

Logit Estimates of Probability of Full-Time Employment
Displaced Workers from Displaced Workers Supplements to CPS
Re-Employed at the Survey Date

1984-91

Variable	Weighted Mean	Coefficient Estimate	Variable	Weighted Mean	Coefficient Estimate
Constant	1.0	.411 (.163)	Age 45-49	.0754	.255 (.113)
Female	.382	-.552 (.0530)	Age 50-54	.0490	.0831 (.130)
Nonwhite	.116	-.243 (.0766)	Age 55-59	.0381	-.354 (.134)
Part-time on Old Job	.110	-1.20 (.0739)	Age 60-64	.0183	-.625 (.180)
Education < 12	.151	-.370 (.0731)	1986	.197	.200 (.0849)
Educ 13-15	.248	.164 (.0651)	1988	.209	.263 (.0846)
Education >=16	.175	.510 (.0797)	1990	.186	.264 (.0871)
Age 25-29	.209	.0959 (.0819)	1992	.243	-.0412 (.0804)
Age 30-34	.176	.137 (.0863)	Displaced in Yr Prior to Survey	.496	.696 (.144)
Age 35-39	.149	.0985 (.0899)	Displaced 2nd Yr Prior to Survey	.479	1.07 (.144)
Age 40-44	.109	.224 (.101)			

Sample Size: 8886

Log Likelihood: -4641.9

Note: Asymptotic Standard Errors are in parentheses. The data are weighted by the CPS final sampling weights. The weighted mean of the dependent variable is 0.749. The base group consists of white individuals from the 1984 DWS who are 20 to 24 years old, have twelve years of education, and who were displaced in January of the survey year.

Table 10

Average Log Wage Change For Displaced Workers
Re-Employed at the Survey Date

by Year

Year	All	FT-FT	FT-PT	PT-FT	PT-PT
1984	-.155 (.0162)	-.0887 (.0152)	-.532 (.0430)	.556 (.0795)	-.0383 (.0790)
1986	-.145 (.0147)	-.0653 (.0136)	-.600 (.0411)	.436 (.0731)	-.127 (.0780)
1988	-.150 (.0143)	-.116 (.0132)	-.545 (.0408)	.591 (.0641)	-.0266 (.0799)
1990	-.125 (.0151)	-.0815 (.0139)	-.476 (.0439)	.405 (.0762)	-.0192 (.0800)
1992	-.154 (.0133)	-.0911 (.0125)	-.571 (.0348)	.557 (.0686)	.0364 (.0661)
N	9144	6117	1765	463	541

Note: The numbers in parentheses are standard deviations of the means. All averages are weighted by the CPS final sampling weights. The four sub-group sample sizes do not sum to the sample size in the "all" column because of missing data on part-time / full-time status. Part-time is defined as less than 35 hours per week.

Table 11

OLS Regression Estimates of Log Wage Change
 Displaced Workers from Displaced Workers Supplements to CPS
 Re-Employed at the Survey Date

1984-91

Variable	Weighted Mean	Coefficient Estimate	Variable	Weighted Mean	Coefficient Estimate
Constant	1.0	.211 (.0460)	Current Job Part-Time	.251	-.463 (.0158)
Female	.382	-.0176 (.0129)	Both Jobs Part-Time	.0571	-.0775 (.0396)
Nonwhite	.116	.0230 (.0189)	1986	.197	-.0044 (.0202)
Education < 12	.151	-.0322 (.0183)	1988	.209	-.0081 (.0199)
Educ 13-15	.248	-.0118 (.153)	1990	.186	.0134 (.0205)
Education >=16	.175	.0167 (.0173)	1992	.243	.0017 (.0194)
Age	34.6	-.00315 (.00065)	Displaced in Yr Prior to Survey	.496	-.147 (.0389)
Tenure on Lost Job	3.87	-.0104 (.0013)	Displaced 2nd Yr Prior to Survey	.479	-.138 (.0391)
Previous Job Part-time	.110	.575 (.0276)			

Sample Size: 8886

Log Likelihood: R-squared: .168

Standard Error of Estimation: .569

Note: Standard Errors are in parentheses. The data are weighted by the CPS final sampling weights. The weighted mean of the dependent variable is -.143. The base group consists of white males from the 1984 DWS who have twelve years of education, whose previous job and current job were full time, and who were displaced in January of the survey year.

Table 12

Estimates of Effect of Displacement on Log Full-time Weekly Earnings

	Year of Displaced Workers Survey					
	Pooled	1984	1986	1988	1990	1992
γ_1	-.0399 (.0050)	-.0399 (.0102)	-.0287 (.0111)	-.0051 (.0111)	-.0668 (.0121)	-.0669 (.0113)
γ_2	-.0097 (.0026)	-.0315 (.0059)	.0140 (.0059)	-.0288 (.0059)	-.0482 (.0057)	.0429 (.0059)
$\gamma_2 + \gamma_3$	-.0988 (.0070)	-.0986 (.0149)	-.0834 (.0151)	-.141 (.0153)	-.0854 (.0166)	-.0831 (.0156)
γ_3	-.0891 (.0074)	-.0671 (.0160)	-.0974 (.0162)	-.112 (.0164)	-.0372 (.0176)	-.126 (.0167)
$\gamma_1 + \gamma_3$	-.129 (.0055)	-.107 (.0125)	-.126 (.0112)	-.118 (.0121)	-.104 (.0128)	-.193 (.0124)
# obs	128714	23630	24581	24801	26943	28759
R ²	.334	.337	.347	.344	.333	.329

Note: Estimates are from the regression

$$(8.1) \quad \ln E_j = X_j \beta + \gamma_1 L_j + \gamma_2 C_j + \gamma_3 L_j C_j + \varepsilon_j$$

γ_1	Prior Earnings (Displaced - Not Displaced)
γ_2	(Current Earnings - Prior Earnings) Not Displaced
$\gamma_2 + \gamma_3$	(Current Earnings - Prior Earnings) Displaced
γ_3	Difference-in-Difference Estimate of Displacement Effect
$\gamma_1 + \gamma_3$	Current Earnings (Displaced - Not Displaced)

The numbers in parentheses are standard errors. These estimates are derived from regressions of log real weekly earnings on dummy variables for sex, race, 9 age categories, 4 educational categories, whether the worker was displaced from his/her job, whether the observation represents "current" earnings (at the January survey date including a Displaced Workers Survey), and the interaction of the displacement dummy and the current observation indicator. The "current" observations for non-displaced workers (n=50090) are from the outgoing rotation groups of the January CPS's that include Displaced Workers Surveys. The "prior" observations for non-displaced workers (n=62578) are a five percent random sample from the merged outgoing rotation group CPS files for the two years preceding each DWS survey. All observations for displaced workers (n=8839 "prior", n=7207 "current") are from the DWS's.

Figure 1

Fraction With Job Loss in Last 2 Years - Males

Figure 2

Fraction With Job Loss in Last 2 Years - Females