Are Workers Permanently Scarred by Job Displacements?

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Abstract:

This paper investigates whether workers suffer lasting "scars" following job displacements. Using David T. Ellwood's (1982) terminology, "scars" represent persistent effects, whereas "blemishes" are transitory adjustments which dissipate over time. More precisely, dislocated individuals are defined as scarred if they continue to earn less or to be unemployed more than their nondisplaced counterparts, even after the conclusion of a several-year adjustment period.

Article:

Displacements may have a transitory impact if workers initially obtain unstable positions but later move to more secure jobs or if low wages are received during short- lasting probationary or training periods. These temporary effects, although important, cause less concern than persistent scarring. It is therefore somewhat surprising that, despite the proliferation of recent research studying the consequences of permanent layoffs, relatively little is known about the duration of the associated adjustment problems.[1]

Data for this study were obtained for heads of households from the 1969-1982 waves (survey reporting years) of the Michigan *Panel Study of Income Dynamics* (PSID). Displacement status was ascertained for the five base years 1971-1975, with respondents defined as permanently displaced if they terminated jobs as the result of plant closings or layoffs (excluding departures from temporary or seasonal jobs) and failed to return to the original employer by the end of the second full calendar year following the layoff. Data were collected on unemployment and weekly wages for the three years preceding and four years following the base period. The base year is hereafter referred to as time t, the following four years are referred to as periods $t + 1$ through $t + 4$, and numbers subtracted from t refer to periods preceding the base year. Information was also assembled indicating whether respondents were displaced in year $t + 5$.

The sample includes household heads between the ages of 21 and 65 at time t who participated in the labor force during some part of each of years t through t +5 and received positive earnings at some time during period t - 2 or t - 3. The analysis is therefore restricted to individuals with fairly strong attachments to the labor force.[2]

***1.** See the survey article by Daniel S. Hamermesh (1989) and the edited volume by John T. Addison (1990) for examples of recent work.*

***2.** The PSID oversamples low-income individuals and has information on relatively few permanent job changers (even with pooling). This raises concern that the results of this study might differ from those obtained using more-representative data. For this reason, regression estimates of postdisplacement unemployment on a commonly defined set of control variables were compared for the PSID and for household heads in the Displaced Workers Survey, a supplement to the January 1984 Current Population Survey. Coefficients from the two sets of regressions were similar, suggesting that the findings of this paper are reasonably robust.*

To reduce the number of multiple observations for given individuals, data were used only for base years in which the individual was displaced and for a maximum of one randomly chosen base year in which no displacement occurred.[3] The sample contains 3,813 person-year observations, 800 for workers losing jobs at time 0 and 3,013 for individuals not displaced during that year.

I. ESTIMATION TECHNIQUE

The effect of displacements occurring at time *t* on employment conditions at *t* + *n* can be estimated from regressions of

$$Y_{i,t+n} = X_{i,t}\alpha + D_{i,t}\beta + \mu_{i,t+n} \tag{1}$$

where *i* is an individual subscript, *Y* is the dependent variable (either weeks of unemployment or the natural log of weekly wages), **X** is a vector of observable characteristics, *D* is a dummy variable indicating whether a displacement occurs at time *t*, and μ is the regression disturbance term. These estimates will be biased, however, if the vector of independent variables fails to control fully for all types of heterogeneity that jointly influence the dependent variable and the probability of displacement. For this reason, two additional specifications were estimated.

In the first, predisplacement wages and predisplacement unemployment ($Y_{i,t-m}$) were included to control for heterogeneity not captured by **X**, yielding the regression equation

$$\begin{aligned} Y_{i,t+n} = X_{i,t}\alpha &+ D_{i,t}\beta \\ &+ Y_{i,t-m}\gamma + \mu_{i,t+n}. \end{aligned} \tag{2}$$

The second specification included persons involuntarily terminating jobs in year 5 (after the end of the observation period) as a comparison group. These equations were of the form

$$Y_{i,t+n} = X_{i,t}\alpha + D_{i,t}\beta + D_{i,t+5}\delta + \mu_{i,t+n} \tag{3}$$

with the bias introduced by unobserved heterogeneity captured by the $\hat{\delta}$. The net displacement effect was then calculated as ß - $\hat{\bar{\delta}}$, where $\hat{\bar{\delta}}$ is the average value of $\hat{\delta}$ for periods *t* through *t* +2.[4] The earnings regressions were estimated using ordinary least squares, and the unemployment equations were estimated as TOBIT models.

3. *The precise selection criteria were as follows: a random variable V was created which, for each individual, had an equal probability of taking integer values 1 ≤ V ≤ 5; the base year was then included in the sample if: 1) YEAR — 1970 — V = 0; or 2) if a permanent displacement occurred either at t or t +5.*

4. *As discussed in the next paragraph,* $\hat{\bar{\delta}}$ *will include a portion of the displacement effect in later years, if adjustments begin prior to the job termination. Jacob Mincer (1986) and I (Ruhm, 1990) have previously used subsequent job losers as a comparison group to account for unobserved differences when analyzing the effects of job mobility.*

There is some ambiguity as to the appropriate choice of lagged dependent variables to be used in equation (2). If a period shortly before the involuntary termination is chosen (e.g., Y_{t-1}) and employment conditions begin to deteriorate prior to permanent separations (as firms institute temporary layoffs or obtain wage concessions), the coefficient on the lagged dependent variable will include a portion of the displacement effect, and $\widehat{ß}$ will understate the impact of permanent layoffs. Conversely, the use of a longer lag period will cause $\widehat{ß}$ to overstate the actual displacement effect if firms strategically terminate workers whose performance has been deteriorating or who previously received high pay, relative to their productivity. This occurs because the displacement coefficient captures the effects of changes that would have been expected even in the absence of mobility.[5]

Preliminary analysis of the comparison group of workers displaced in year t +5 revealed that wage and employment adjustments begin two calendar years before permanent layoffs. For this reason, equation (2) was estimated with the dependent variable alternatively lagged one and three periods prior to the base year. Regressions with Y_{t-1} (Y_{t-3}) included are likely to provide lower (upper) bounds on the actual displacement effect.

TABLE 1—POSTDISPLACEMENT WAGE AND UNEMPLOYMENT DIFFERENTIALS

	Unemployment				Weekly wages			
Time period	a	b	c	d	a	b	c	d
t	8.35	8.43	7.10	8.73	−0.1058	−0.1054	−0.0645	−0.1149
	(27.28)	(27.22)	(25.94)	(26.99)	(5.39)	(6.49)	(4.64)	(5.73)
$t+1$	4.32	4.18	3.69	4.64	−0.1751	−0.1664	−0.1365	−0.1820
	(17.27)	(17.15)	(15.53)	(19.98)	(8.73)	(9.89)	(9.09)	(8.88)
$t+2$	2.08	2.11	1.79	2.24	−0.1623	−0.1585	−0.1188	−0.1690
	(9.45)	(9.21)	(8.29)	(9.65)	(7.77)	(8.83)	(7.25)	(7.91)
$t+3$	1.45	1.40	1.16	1.77	−0.1486	−0.1435	−0.1067	−0.1627
	(7.00)	(6.83)	(5.84)	(8.07)	(7.67)	(8.03)	(6.43)	(7.17)
$t+4$	1.27	1.12	0.88	1.85	−0.1473	−0.1402	−0.1117	−0.1698
	(5.81)	(5.62)	(4.60)	(8.56)	(6.84)	(7.34)	(6.33)	(7.73)
Control for heterogeneity:	none	Y_{t-3}	Y_{t-1}	D_{t+5}	none	Y_{t-3}	Y_{t-1}	D_{t+5}
				0.86				−0.0434
				(2.61)				(1.77)

Notes: See text for explanation of columns a–d. Unemployment equations are estimated using maximum-likelihood TOBIT methods; the wage equations are estimated by ordinary least squares. The unemployment coefficients show the impact of marginal changes in the regressors on actual joblessness and are obtained by evaluating the TOBIT regressions at the independent variable means. The wage coefficients show the impact of base-year displacements on the natural log of weekly wages. Numbers in parentheses are t statistics. Regressions include controls for experience, education, marital status, race, sex, city size, tenure, the survey year, industry, occupation, and age (>55 years old). Coefficient on D_{t+5} is the average obtained from regressions for years t through $t+2$.

5. *Closely analogous concerns are raised in the econometric analysis of training programs, where reductions in employment and earnings are typically observed shortly prior to the beginning of the training period. See Orley Ashenfelter (1978), Ashenfelter and David Card (1985), and Card and Daniel Sullivan (1988) for discussions of these issues.*

II. WAGE SCARS AND UNEMPLOYMENT BLEMISHES

Permanent layoffs are associated with substantially elevated initial unemployment. Between years t and $t + 2$, dislocated workers were three times more likely than their counterparts to experience some unemployment (83.2 percent vs. 26.3 percent), averaged six times as many weeks out of work (23.9 vs. 3.8 weeks), were eight times more probable to be unemployed for more than six months (35.8 percent vs. 4.5 percent), and were jobless for more than one year 16 times as frequently (12.9 percent vs. 0.8 percent).

Dislocated workers were also more than twice as likely to experience wage reductions exceeding 25 percent between $t - 2$ and $t + 2$ (28.6 percent vs. 14.1 percent), and lost more than 10 percent of previous wages 1.6 times as often (40.1 percent vs. 25.0 percent). Although displaced individuals averaged only a 1.5-percent earnings reduction over the four years, they missed out entirely on the 8.4-percent real wage gain obtained by the control group over the same period.

I now turn to considering whether displacements leave lasting scars. Table 1 displays regression results of unemployment and wage equations estimated for periods t through $t + 4$. For each dependent variable, coefficients in column a are obtained from estimating equation (1), which includes no controls for unobserved heterogeneity. Coefficient in columns b and c are from equation (2) with Y_{t-3} and Y_{t-1}, respectively, incorporated.[6] Estimates in column d are from equation (3), with the final entry in the column indicating $\hat{\delta}$. In this case, the effects of base-year layoffs are calculated by subtracting $\hat{\delta}$ from the estimated coefficients on D_t. Displacement impacts in the early years (t through $t + 2$) indicate adjustment costs in the periods immediately following permanent job loss. The major focus of this paper, however, is on longer-run effects, as measured by the coefficients for years $t + 3$ and $t + 4$.

Permanent layoffs lead to substantial temporary unemployment blemishes but far more enduring earnings scars. Displaced workers were out of work eight weeks more than their observably similar counterparts in the year of the separation, four additional weeks in period $t + 1$, and two extra weeks at $t + 2$. By year $t + 3$ they were jobless only 1.5 weeks more than the peer group, and the $t + 4$ increase was just six days (see Table 1, column a). Thus, at least 85 percent of the initial rise in joblessness dissipated prior to year $t + 4$.[7]

In contrast, wage effects are large and lasting. The estimates in column a of Table 1 imply that base-year weekly earnings of job losers were 10.0 percent below those of their nondisplaced peers. This understates the reduction caused by dislocations, however, since period t wages partially reflect the (typically higher) pay received on the preseparation job.[8] It is therefore not surprising that the period $t + 1$ wage differential was 1.6 times greater (16.1 percent) than the base-year effect.

6. *If data for period t -3 were unavailable (in Table 1, column b), corresponding information for period t -2 was used.*

7. *Ellwood (1982) also finds that current unemployment has only a small effect on future joblessness. His results are for teenagers.*

8. *For example, individuals displaced in July worked more than half the base year in the predisplacement position.*

More significantly, almost none of the *t* + 1 wage reduction dissipated with time. The earnings gap remained at 13.8 percent and 13.7 percent, respectively, in years *t* + 3 and *t* + 4, which was 85 percent of the disparity predicted for year *t* + 1.[9]

The inclusion of controls for unobserved heterogeneity only slightly reduces the measured impact of base-year terminations. As expected, the estimated displacement effect was somewhat larger when Y_{t-3} was added to the regressions (Table 1, column b) than when Y_{t-1} was included (column c). Even in the latter case, however, unemployment increased temporarily (by 7.1 weeks in year *t* and by 3.7 weeks in year *t* + 1), and the wage loss remained above 10 percent in periods *t* + 2 through *t* + 4. If anything, addition of the lagged dependent variables strengthens the earlier finding of unemployment blemishes and permanent wage scars. This is seen by noting, that although the unemployment effect fell to around one week by period 4, the upper (lower) bound on the corresponding wage loss remained at 13.1 (10.6) percent.

The coefficient on D_{t+5}, in column d of Table 1, indicates that unobserved heterogeneity accounted for a 0.9-week disparity in annual unemployment and a 4-percent wage differential between displaced and nondisplaced workers. Subtracting this from the coefficients on D_t, the net increase in unemployment was 7.9 weeks at time 0, 3.8 weeks in period *t* + 1, but less than 1 week annually in periods *t* + 3 and *t* + 4. Conversely, the estimated wage loss exceeded 11 percent in each of years *t* + 2 through *t* + 4, which was almost 90 percent as large as the maximum reduction observed in period *t* + 1. Interestingly, these displacement effects are quite close to the lower-bound estimates presented in column c of Table 1 and again indicate transitory employment shocks but lasting earnings changes.

III. SUMMARY

Although permanent loss of jobs leads to unemployment distinguished by its long duration, there is no evidence that these initial difficulties translate into lasting scars. Conversely, involuntary terminations typically result in a significant loss of long-term earnings potential. Four years after displacement, job losers are out of work only one week more than their nondisplaced counterparts but continue to earn 10-13 percent less. The lasting wage reductions suggest significant worker attachments to specific jobs. Future research is needed to investigate the sources of these attachments.

APPENDIX

Data Set and Variable Construction

Data were taken from the 1969-1982 waves of the *Panel Study of Income Dynamics* (PSID). The PSID was obtained from the Inter-University Consortium for Political and Social Research (ICPSR); requests for descriptions of the data set and for the raw data should be directed to the ICPSR.

9. *Evidence of persistent wage losses has also recently been obtained by Robert Topel (1989). The wage-equation model in column a of Table 1 was also estimated using the standard two-stage correction for sample- selection bias. Except for period t, for which the estimates indicated that workers with low potential wages were more likely to be reemployed, the ordinary and two-stage least-squares estimates were virtually identical. The latter, if anything, showed greater persistence of wage losses.*

Analysis was restricted to household heads between the ages of 21 and 65 during the 1971-1975 base years who: 1) participated in the labor force during some part of the base year (t) and the following five years; 2) received earnings during at least one of the years *t* - 3 or *t* - 2; and 3) fulfilled the sampling criteria specified in footnote 3. Respondents were excluded if the head of household had changed during the relevant sample period or if they retired during the five years subsequent to the base year.

All data manipulations and analyses were performed using SAS (SAS Institute, 1985) on the Boston University mainframe computer, with the exception of the TOBIT regressions which were estimated using LIMDEP (William H. Green, 1986).

Individuals were defined to be displaced if they involuntarily terminated jobs in the base period [or in year 5 for the comparison group described in equation (3)] and failed to return to the old employer by the end of the second full calendar year following the separation. Dependent variables used in the analysis were weeks of calendar-year unemployment and the log of weekly wages. The latter were calculated by taking the natural logarithm of real annual earnings divided by number of weeks worked, with annual earnings adjusted to 1972 prices using the GNP deflator. Missing values were assigned for periods in which the natural log of weekly wage was less than 3 (this corresponds to weekly wages of less than $20).

Except for the lagged dependent variables, the regressors are all dummy variables and refer to individual or job characteristics in the base year. They equal 1 if the following are true (and 0 otherwise).

Experience**:** Labor market experience is less than 10 years. Data on actual labor market experience is available and used beginning in 1974. For earlier years, experience is calculated as 1974 experience less the difference between 1974 and the survey date (e.g., in 1972, experience was set equal to 1974 experience less 2 years).
School: Education is less than or equal to 12 years.
Married: Respondent is married or permanently cohabiting.
Female: Respondent is female.
Nonwhite: Respondent is black or Spanish- American.
Age: Age is greater than 55 years old.
City1: City size is greater than 100,000.
City2: City size is less than 25,000.
Blue Collar: One-digit predisplacement occupation code is 5, 6, or 7.
Manufacturing: One-digit predisplacement industry code is 3 or 4.
Professional: One-digit predisplacement occupation code is 1 or 2.
Tenure1: Predisplacement seniority is 1-3 years.
Tenure2: Predisplacement seniority is 4-9 years.
Tenure3: Predisplacement seniority is 10-19 years.
Tenure4: Predisplacement seniority is at least 20 years.
Survey Year: Four dummy variables were included indicating the survey years 1972-1975 (1971 is the excluded category.)

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