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Job Displacement, Relative Wage Changes, and Duration of Unemployment

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Using special CPS data on displaced workers, this article investigates the wage consequences of job displacement in a framework that emphasizes the effects of past job duration(s) and unemployment duration(s) on postdisplacement wages. Our model also attempts to take account of the simultaneity between unemployment duration and the postdisplacement wage. It is found that duration strongly reduces subsequent earnings and that considerable overstatement of the loss in firm-specific training investments is implied by conventional routes to measuring wage losses.

I. Introduction

This article investigates the economic implications of involuntary job loss occasioned by plant closings and employment cutbacks for a large, nationally representative sample of male workers displaced between January 1979 and January 1984. The goal is to identify the determinants of changes in earnings following displacement. To this end, we employ three different specifications for the postdisplacement earnings function. An important feature of the analysis is the justification and use of a flexible specification

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for the postdisplacement wage equation that allows for the effects of past job duration and unemployment duration; factors that have been ignored in the small but growing displacement literature (e.g., Shapiro and Sandell 1985; Podgursky and Swaim 1987).

Our principal findings are as follows. First, the returns to tenure observed on the predisplacement job do not vanish on displacement; that is, previous tenure has a positive impact on postdisplacement wages. Thus, the negative coefficient on previous tenure widely reported in the wage-loss literature should be read as simply indicating that tenure on the lost job raises wages on that job by more than it does on the "second" job. We identify the degree of inflation of the wage-loss estimate produced by uncritical use of the tenure argument: for the worker with median tenure the degree of overstatement is around 45%.

Second, the length of the intervening spell of unemployment emerges as a potent source of reduced earnings on the postdisplacement job. Our simple ordinary least squares (OLS) results suggest that an increase in unemployment duration of 10% will lower wages on the subsequent job by .6%–.8%. However, we also report evidence of simultaneity between the postdisplacement wage and unemployment duration. Allowing for simultaneity actually strengthens the negative impact of duration: a corresponding increase in duration reducing wages between .8% and 1.4%.

Third, the wage consequences of a change in industry and occupation, although not location, are profound. Depending on the specification of the postdisplacement wage equation, our preferred estimates suggest that industry shifts lower the postdisplacement wage in the range 16.1%–19.8% and that occupational shifts are associated with lower wages of between 5.4% and 13.9%.

II. An Accounting Framework

In seeking to discover the effects of job displacement on earnings, we shall argue that the overriding priority is to work with a flexible earnings function so as to avoid the pitfalls of the conventional displacement literature while incorporating insights into the role of individual and job match heterogeneity offered by the modern earnings-function literature. The former literature typically employs the standard Mincerian earnings function to inquire into the sources of earnings loss and in so doing places very strong restrictions on the parameters of the model. The latter literature emphasizes the biases on the tenure (and experience) coefficient(s) introduced by omitted variables that are positively related to tenure and earnings (Topel 1986; Abraham and Farber 1987; Altonji and Shakotko 1987).

Consider the following representation of the earnings function, which, for expositional convenience, considers only tenure and unemployment arguments and neglects individual and time subscripts. The natural logarithm of individual earnings in job j may be expressed

$$\ln W_j = \sum_{s=1}^{j-1} \alpha_s \text{TENURE}_s + (\alpha_j + \beta_j)\text{TENURE}_j + \sum_{s=1}^{j} \gamma_s \text{SLU}_s + u_j, \tag{1}$$

where TENURE_s is completed job duration on job s, TENURE_j is current tenure, SLU_s is completed spell length of unemployment, or search, prior to job s, α is the transferable, general-training component of the return to TENURE, β is the nontransferable, specific-training component, and u is a disturbance term.

Note that in equation (1) we allow investments in general training to differ from job to job, so that $\alpha_1 \neq \alpha_2 \neq \ldots \alpha_j$. The α_s coefficients will provide us with information on the past investment profile. The TENURE_s sequence enriches the model by incorporating information on prior mobility (Mincer and Jovanovic 1981).

Turning to the unemployment duration argument, it is possible that unemployed workers may suffer some depreciation of their human capital (Lazear 1976; Kiker and Roberts 1984) that, coupled with stigma effects (Heckman and Borjas 1980), may be expected to lower wages. A declining reservation wage will also yield a negative association between wages and duration.[1] In addition, less efficient searchers may find themselves unemployed for longer durations and obtain lower wages on reemployment (Mincer 1986). Each mechanism runs counter to the standard productive-search model wherein, given a constant reservation wage assumption, longer spell duration tends to yield higher postunemployment wages (Stigler 1962; Lippman and McCall 1976).

Simple manipulation of equation (1) yields

$$\begin{aligned} \ln W_j = \alpha_1 \sum_{s=1}^{j} \text{TENURE}_s + \sum_{s=2}^{j-1} (\alpha_s - \alpha_1)\text{TENURE}_s \\ + (\alpha_j + \beta_j - \alpha_1)\text{TENURE}_j + \sum_{s=1}^{j} \gamma_s \text{SLU}_s + u_j. \end{aligned} \tag{2}$$

Letting $\alpha_1 = \alpha_2 = \ldots = \alpha_j$ and $\gamma_s = 0$ $(s = 1, \ldots, j)$, we arrive at the more familiar wage equation

$$\ln W_j = \alpha_1 \text{EXPERIENCE} + (\alpha_j + \beta_j - \alpha_1)\text{TENURE}_j + u_j. \tag{3}$$

In other words, specification (3) compounds the effects of previous job durations and previous unemployment durations in a single coefficient;

[1] Theoretical reasons for declining reservation wages include finite lives, income constraints, and exhaustion of benefits.

and the current tenure coefficient is then to be interpreted as the effect of tenure on wage growth over and above that resulting from experience, α_1. Note that the specific, nontransferable component, β_j, of the tenure coefficient is *not* identified unless one imposes the restriction that $\alpha_j = \alpha_1$. Furthermore, if the experience variable is subject to measurement error, the tenure coefficient will also be affected. The estimate of α_1 will be biased downward, thus biasing upward the tenure coefficient. One possible source of measurement error arises from the difficulty of distinguishing between previous job durations and unemployment durations. Finally, note that *if the worker is increasingly investing in general training* (e.g., younger workers or those who, being aware of their impending displacement, switch from specific- into general-training investments) so that $\alpha_j > \alpha_1$, then the tenure coefficient will be increased even though no additional investment in specific training is observed.

The general approach implied by equation (1) is thus likely to provide richer insights into worker investment profiles, job-matching strategies, job-search behavior, and unemployment effects on wages. Moreover, its formulation allows one to test the assumptions implicit in equation (3). The empirical application of equation (1) is remitted to Section IV.

III. The Data

Our data are drawn from a special survey conducted in January 1984 as a supplement to the Current Population Survey (CPS). The purpose of this Displaced Worker Survey (DWS) was to obtain information on the number and characteristics of all adult workers displaced from their jobs between January 1979 and January 1984. In order to identify such workers, CPS respondents were first asked whether they or any other household member had lost a job over the 5-year sample period "because of a plant closing, an employer going out of business, a layoff from which [he/she] was not recalled or other similar reasons." If the answer to this question was in the affirmative, the respondent was then asked to identify which among the following categories best fitted the reason for job loss: (i) plant or company closed down or moved; (ii) plant or company was still operating but job lost because of (*a*) slack work, (*b*) position or shift abolished, or (*c*) seasonal job was completed; (iii) self-employment business failed; and, (iv) other reasons.

A series of questions were then asked of respondents as to the nature of the lost job, including year of loss, years of tenure, and earnings.[2] Another set of questions sought to determine what transpired after job loss. These included the length of the unemployment spell, unemployment insurance benefit status, exhaustion of benefits, and worker relocation. Finally, if the

[2] In the event of more than one displacement, the data refer to the job with the longest duration.

worker was reemployed at the time of the interview, follow-up questions were deployed so as to determine the current (January 1984) level of earnings, occupation, and industry attachment among other things.

The value of the DWS is precisely that it is a large, nationally representative microdata set. In weighted terms the survey identifies some 13.9 million displaced workers. Previous studies of displaced workers have either been of the case-study type (reviewed by Gordus, Jarley, and Ferman [1981]) or have investigated small samples of laid-off workers derived from the Panel Study of Income Dynamics (PSID) or the National Longitudinal Survey (NLS) (Shapiro and Sandell 1985; Maxwell 1986; Hamermesh 1987). Other, broader-based econometric studies of trade-displaced workers have the disadvantage that many of those identified as displaced returned to their old employers (Neumann 1978; Corson and Nicholson 1981).

All of this is not to suggest that the DWS is without its own difficulties. First, the rotating nature of the CPS makes longitudinal use difficult. Our response is to focus on relative wage changes in the wake of displacement, namely the difference between pre- and postdisplacement weekly earnings, suitably adjusted for interim price movements. Second, reliance on retrospective information going back up to 5 years raises the problem of recall bias. Fortunately, there are no signs that this form of measurement error disturbs our findings with respect to the contribution of tenure on the lost job to the postdisplacement wage.[3]

Third, there is some ambiguity as to the reported unemployment duration measure. The duration data supplied in the DWS do not necessarily refer to a single spell of unemployment. They may include a period in which the respondent was out of the labor force and even more than one spell of unemployment if the respondent links such spells to the initial job loss. The possibility of unidentified, intruded spells of temporary employment means that we cannot, in practice, identify current tenure on the postdisplacement job as the difference between January 1984 and the date of job loss plus reported spell length of unemployment. In this article, we simply use year-of-displacement dummies as a crude proxy for tenure on the postdisplacement job.

Despite this fuzziness, one major advantage of the DWS is that it contains information on completed spells of unemployment. But there remain two sources of right censoring in the duration data that will need ultimately to be accommodated in our analysis. Continuous duration data are supplied for those whose jobless spells were still in progress as of January 1984 but only up to 99 weeks of joblessness. To those who are currently in an

[3] When we estimated separate equations for specific years, the predisplacement tenure coefficient in the postdisplacement wage equations was larger for the most recent years. This would seem to imply that, if anything, recall errors understate the results reported below with respect to the impact of previous tenure.

ongoing spell of unemployment for less than the 99-week cut-off, therefore, must be added those workers whose joblessness either has exceeded or exceeds 99 weeks by an unknown margin. Both sources of sample truncation will be factored into our analysis.

A final consideration is that job tenure on the predisplacement job could be spuriously productive of income in the postdisplacement job if some of the sample of displaced workers returned to their previous employer. Bias from this source is unlikely to be profound because we are dealing with subcategories of the *permanently* displaced, a large proportion of whom lost jobs by reason of plant closure or relocation. Accordingly, deciding whether or not the worker was merely temporarily separated from his employer is much less problematic here than in other large data sets (e.g., the Social Security Administration's Longitudinal Employee-Employer Data). Moreover, our exclusion of those workers displaced because of seasonal factors and for "other reasons" should serve further to reduce problems of data contamination. For these reasons, our finding that tenure on the predisplacement job is productive of income in the postdisplacement job is unlikely to be a chimera produced by the displaced worker returning to his previous employer.[4]

In addition to excluding groups of workers by source of displacement, we confined our sample to male workers between 20 and 65 years of age who were displaced from full-time nonagricultural jobs over the full 5 years covered by the DWS and who were economically active as of January 1984. These restrictions produced an unweighted sample of 3,223 workers of whom 2,007 were reemployed in either full-time or part-time jobs by January 1984 and who reported a wage for the new job. Descriptive data on the full sample and the subsample of the reemployed are given in Appendix A.

IV. Empirical Implementation of the Model

In specifying equation (1), no account was taken of the possible effects of unobserved job match or individual heterogeneity. Although our data do not allow us adequately to address the job-match heterogeneity argument (most notably because of missing data on completed tenure in the postdisplacement job),[5] we will attempt to control for individual heterogeneity by conditioning the postdisplacement wage equation on the predisplacement wage (Topel 1986). In implementing that specification we initially allow the coefficient on the predisplacement wage to be freely estimated (as do Kiefer and Neumann [1979]) but subsequently impose

[4] In support of this contention, we note that restricting the sample to those who moved industry—a crude proxy for *not* returning to the previous employer—actually served to increase the return to predisplacement tenure on the postdisplacement job.

[5] For an explicit empirical test of the job matching model, see Flinn (1986).

the restriction that the coefficient equals unity, thereby providing a better control for permanent, individual unobserved heterogeneity (in the manner of Bartel and Borjas [1981]).

We restrict our attention to just two job pairs, namely the predisplacement job, $j - 1$, and the postdisplacement job, j. Given the nature of the data set, we perforce ignore the effects of unemployment duration prior to $j - 1$, and, accordingly, the coefficients reported below are to be interpreted with some caution since we are now working with a tentative approximation to the model.

Accordingly, we specify the following pre- and postdisplacement wage equations for individual i (where X_i is a matrix of individual and demographic variables):

$$\ln W_{i,j-1} = \alpha_1 \text{EXPERIENCE}_{i,j-1} + (\alpha_{j-1} + \beta_{j-1} - \alpha_1)\text{TENURE}_{i,j-1} + X_{i,j-1}\Omega + u_{i,j-1}, \tag{4}$$

$$\ln W_{ij} = \alpha_1 \text{EXPERIENCE}_{ij} + (\alpha_{j-1} - \alpha_1)\text{TENURE}_{i,j-1} + (\alpha_j + \beta_j - \alpha_1)\text{TENURE}_{ij} + \gamma_j \text{SLU}_{ij} + X_{ij}\Omega + u_{ij}, \tag{5}$$

$$\begin{aligned} \ln W_{ij} = {} & \delta \ln W_{i,j-1} + (1 - \delta)\alpha_1 \text{EXPERIENCE}_{ij} \\ & + [(1 - \delta)(\alpha_{j-1} - \alpha_1) - \delta\beta_{j-1}]\text{TENURE}_{i,j-1} \\ & + [\alpha_j + \beta_j - \alpha_1(1 - \delta)]\text{TENURE}_{ij} + \gamma_j \text{SLU}_{ij} \\ & + (X_{ij} - \delta X_{i,j-1})\Omega + (u_{ij} - \delta u_{i,j-1}), \end{aligned} \tag{6}$$

and, finally,

$$\ln W_{ij} = \ln W_{i,j-1} + (\alpha_j + \beta_j)\text{TENURE}_{ij} - \beta_{j-1}\text{TENURE}_{i,j-1} + \gamma_j \text{SLU}_{ij} + (X_{ij} - X_{i,j-1})\Omega + (u_{ij} - u_{i,j-1}). \tag{7}$$

Our focus will be upon the estimates of the tenure and unemployment duration coefficients. In particular, equation (4) should yield the return to tenure over and above that resulting from experience (i.e., $[\alpha_{j-1} + \beta_{j-1} - \alpha_1]$); equation (5) should give us an estimate of the transferable component of this return above α_1 (i.e., $[\alpha_{j-1} - \alpha_1]$); and from equation (7) we have an estimate of the loss in specific-training investments ($-\beta_{j-1}$). Note that interpretation of the parameters of equation (6), which has been widely used in the literature, is in fact clouded because of the presence of a constant of proportionality that affects most of the coefficients.[6]

[6] Other problems that attach to the inclusion of the predisplacement wage as a regressor have to do with its endogeneity and possible association with unemployment duration as a result of the procedure used to deflate the predisplacement wage. In our case, the latter correlation is negligible.

In practice, we substitute age (EXPERIENCE + SCHOOL + 6) for experience and hence also include a schooling variable. The age variable is not entered in continuous form but, rather, via three linear splines.[7] Our basic estimating equation for the predisplacement wage includes—in addition to the age, tenure, and schooling arguments—dummies for race, skill level, city, and region. The postdisplacement wage equations supplement these variables not only with the duration of unemployment but also with dummies for changes in residence, industry, occupation, and full-time job status. As noted in Section III, year dummies picking up the time of displacement are substituted for an unreported current tenure argument. In estimating equation (7) we included the full set of X_{ij} variables to investigate additional sources of earnings loss.

Equations (4)–(7) were first estimated by OLS. Equation (4) is presented in Appendix B. The predisplacement wage function is conventional enough, and all coefficients are significant at the .05 level or better and of the expected sign. The statistical significance of the age splines indicates major differences between the three age profiles. Between the ages of 20 and 35 earnings increase, ceteris paribus, at 1.9% a year; after 35, earnings continue to rise at a reduced rate of .2% a year until 50 years of age. Thereafter, earnings decline at some 2.1 percent a year.

Ordinary least squares (OLS) estimates of equations (5)–(7) are given in columns 1, 3, and 5 of table 1. Focusing first on the basic postdisplacement wage equation, an interesting finding is the significantly positive coefficient of TENURE1. It will be recalled that the initial tenure argument was retained to test the hypothesis that the general-training component of the return to tenure on the first job is not fully captured by the age (or experience) variable(s).

The schooling, age, race, location, and year dummies are all significant and have the same signs as in the predisplacement wage equation.[8] These results, together with the obvious effect on (weekly) earnings of changing from full- to part-time job status, need not detain us further. More notable findings are the strongly negative effects of changes in industry and

[7] We also experimented with a quadratic spline specification. This exercise did not produce improved results and, for expositional convenience, only results for the linear spline formulation are given here.

[8] It will be recalled that the year dummies proxy tenure on the postdisplacement job. Their coefficients tend to decline the closer displacement is to the end of survey period, although the decline is not monotonic not only because the proxy picks up the effect of macroeconomic fluctuations but also because those displaced in the most recent interval are less likely to be reemployed and hence report earnings. Note that we would anyway be reluctant to place undue emphasis on even a correctly identified current tenure variable precisely because of the censored design of the sample; that is to say, the maximum possible tenure on the postdisplacement job is only 5 years in this data set.

Table 1
Postdisplacement Wage Equations

Variable	Specification (1)	(2)	(3)	(4)	(5)	(6)
CONSTANT	5.040*	5.138*	2.521*	2.294*	.088	−.029
	(59.42)	(45.35)	(17.93)	(13.83)	(1.05)	(.27)
SCHOOL	.045*	.036*	.027*	.021*	.009**	.008
	(9.31)	(6.34)	(6.06)	(4.01)	(1.96)	(1.49)
AGE:						
20–35	.015*	.019*	.004	.007**	−.006***	−.002
	(4.65)	(5.36)	(1.44)	(2.23)	(1.91)	(.70)
36–50	−.012**	−.013**	−.002	−.003	.008	.006
	(2.13)	(2.22)	(.37)	(.51)	(1.41)	(1.00)
51–65	−.025*	−.032*	−.018**	−.025*	−.011	−.020**
	(2.95)	(3.40)	(2.34)	(2.99)	(1.31)	(2.19)
TENURE1	.013**	.014*	−.002	−.002	−.016*	−.015*
	(2.53)	(2.72)	(.48)	(.37)	(3.33)	(2.91)
TENURE1SQ	−.00045**	−.0005**	−.0001	−.0001	.00025	.00021
	(2.39)	(2.53)	(.56)	(.62)	(1.32)	(1.07)
WHITE	.183*	.118**	.090**	.047	.000	−.011
	(4.40)	(2.49)	(2.39)	(1.09)	(.01)	(.23)
UNSKILLED	−.153*	−.143*	−.128*	−.116*	−.104*	−.094*
	(6.16)	(5.44)	(5.71)	(4.87)	(4.21)	(3.67)
SMSA	.076*	.105*	.043**	.066*	.011	.033
	(3.36)	(4.36)	(2.10)	(3.02)	(.50)	(1.43)
SOUTH	−.073**	−.114*	−.009	−.028	.053	.043
	(2.19)	(3.13)	(.29)	(.84)	(1.62)	(1.23)
MOVEREG	.036	...	−.007	...	−.048***	...
	(1.34)		(.30)		(1.83)	
MOVEIND	−.200*	...	−.175*	...	−.150*	...
	(8.06)		(7.80)		(6.10)	
MOVEOOC	−.145*	...	−.096*	...	−.049**	...
	(5.95)		(4.35)		(2.02)	
PART TIME	−.850*	...	−.802*	...	−.756*	...
	(12.72)		(13.30)		(11.39)	
YEAR:						
1979	.148*	...	.123*	...	.098**	...
	(3.62)		(3.32)		(2.42)	
1980	.087**	...	.066**	...	.046	...
	(2.35)		(1.98)		(1.24)	
1981	.109*	...	.078**	...	.049	...
	(3.20)		(2.54)		(1.44)	
1982	.096*	...	.067**	...	.040	...
	(2.98)		(2.31)		(1.24)	
lnWAGE1	0	0	.509*	.550*	1	1
			(21.35)	(21.77)		
lnDURATION	−.059*	−.067*	−.066*	−.074*	−.073*	−.080*
	(9.01)	(9.75)	(11.12)	(11.98)	(11.11)	(12.03)
LAMBDA	...	−.256*	...	−.137*	...	−.040
		(4.50)		(2.65)		(.73)
R^2	.310	.196	.439	.350	.193	.108

NOTE.—Absolute t-values are given in parentheses.
* $p \leq .01$.
** $.01 < p \leq .05$.
*** $.05 < p \leq .10$.

occupation (although not location) and the length of the spell of unemployment on the postdisplacement wage. Changes in industry and occupation following job loss are associated with reductions in the postdisplacement wage of 18.1% and 13.5%, respectively. For the reemployed worker, the estimated elasticity of the postdisplacement wage with respect to duration is −.06. The clear suggestion of this latter result is that, on balance, declining reservation wages, depreciation, and stigma effects associated with longer spell lengths dominate productive search outcomes. We return to a fuller discussion of the unemployment duration results below.

The wage functions reported in columns 3 and 5 of table 1 both suggest that higher education levels reduce any loss in earnings attendant on displacement and that unskilled workers experience significantly higher earnings losses than their semiskilled and skilled counterparts. In both specifications, the strongly negative effects on earnings of unemployment duration and changes in industry and occupation are again evident.

Greater reliance is to be placed upon the wage-difference model (eq. [7]) because of its ability to control for permanent, unobserved individual heterogeneity. Here the results in column 5 strongly confirm that a substantial overstatement of the earnings loss is implied by conventional routes to the measurement of firm-specific training investments. Although tenure is negatively and significantly associated with the change in wages, the coefficient is markedly lower than that implied by the predisplacement equation. Note, too, that its value is almost exactly the difference between the corresponding coefficients reported in Appendix B and column 1 of table 1. The same is true of the other coefficients significantly associated with wage change; namely, schooling, age (the first age spline), and skill level. The elasticity of the postdisplacement wage with respect to duration strengthens to −.07.

Note that as the coefficient on the predisplacement wage rises from 0 to .5 to unity (specifications 1, 3, and 5, respectively) all coefficients of the variables that influence the *postdisplacement wage but not the predisplacement wage* steadily decline in magnitude (with the exception of unemployment duration), probably reflecting the presence of unobserved individual heterogeneity.

Finally, table 1 also contains selectivity-adjusted results for the three postdisplacement wage equations. As noted earlier, our sample is truncated since a considerable fraction of the displaced worker sample is not employed full-time or part-time as of January 1984. For such workers the effects of the determinants of postdisplacement earnings may differ systematically from those of employed workers. Typically, it is assumed that if a worker is reemployed, the offered wage is either equal to or greater than the reservation wage, and conversely for those currently unemployed. In order to account for the possibility of censoring bias—namely, that the sample of reemployed workers is not in fact a random sample drawn from the (stock of the) unemployed—the conventional two-step selectivity adjustment procedure suggested by Heckman (1979) was implemented. At this stage a number of the variables contained in the basic postdisplacement wage equation were excluded from the (probit) analysis because they are observed only in the event of reemployment (changes in industry, occupation, and job status). Also, because they are intrinsically associated with the censoring mechanism, the year-of-displacement dummies were dropped from the wage equations. The underlying reemployment equation (probit) is reported in Appendix C and the selectivity-adjusted estimates appear alongside their corresponding specifications in table 1.

The standard search model with a constant reservation wage and no human capital depreciation or stigma effects would suggest that the subsample of still-unemployed workers would have a higher postdisplacement wage due to higher reservation wages (search efficiency). The presence of such effects, however, might justify at least in part the use of unemployment duration as an additional regressor. If that variable cannot capture all the above effects or if longer unemployment duration picks up job search inefficiencies (Mincer 1986), then the association between duration and the probability of being in the reemployed sample will make it difficult to disentangle the various effects and the expected sign of the lambda coefficient is no longer unambiguous. The lambda coefficients reported in the table should be interpreted with these observations in mind.

Both the postdisplacement wage equations displayed in columns 2 and 4 of table 1 point to negative selection. That is to say, holding constant unemployment duration, the currently unemployed are predicted to have somewhat higher postdisplacement wages than their equivalent reemployed counterparts. (Inclusion of the selectivity argument does not alter the other coefficients substantively.) Note, however, that the lambda coefficient, although still negative, is statistically insignificant in the wage-difference equation (col. 6). Also note that in all three specifications the negative effect of unemployment duration on postdisplacement wages is strengthened, the elasticities now ranging from −.07 to −.08.

A correct reading of the search-theoretic literature would suggest that the postdisplacement wage and the duration of unemployment are *jointly* determined. Not only does duration have a direct impact on reservation and offered wages but the latter will also feed back into duration. Unfortunately, only a few studies have explicitly sought to recoup the structural parameters of the wage and duration equations (Kiefer and Neumann 1979; Lancaster and Chester 1984). When self-selection into reemployment status can be ignored, the two-stage least squares (2SLS) estimator can be employed. With censoring, simultaneous equation models with truncated dependent variables have to be considered (Amemiya 1974; Lee, Maddala, and Trost 1980; Lee 1982).

Table 2
Simultaneous Determination of the Postdisplacement Wage and Unemployment Duration Equations, Two-Stage Least Squares (2SLS) and Selectivity Adjusted Simultaneous Equations (SASE)

Equation	2SLS	SASE	2SLS	SASE	2SLS	SASE
lnWAGE2:						
lnDURATION	−.022	−.044***	−.133*	−.117*	−.234*	−.174*
	(.83)	(1.98)	(5.31)	(5.71)	(7.85)	(8.08)
lnWAGE1	0	0	.523*	.559*	1	1
			(20.83)	(21.28)		
LAMBDA	...	−.248*	...	−.148*	...	−.070
		(4.24)		(2.78)		(1.22)
lnDURATION:						
lnWAGE2	−.793*	−1.460**	−.793*	−1.460**	...	...
	(4.10)	(2.15)	(4.10)	(2.15)		
lnDIFF	...	...	...	...	−.793*	−1.460**
					(4.10)	(2.15)
lnWAGE1	.621*	.911**	.621*	.911**	−.173	−.548
	(4.58)	(2.50)	(4.58)	(2.50)	(1.34)	(1.62)
LAMBDA	...	−1.305**	...	−1.305**	...	−1.305**
		(2.23)		(2.23)		(2.23)

NOTE.—The estimates shown for the duration equation are the same across specifications because the set of regressors in the reduced-form wage equations are identical.
* $p \leq .01$.
** $.01 < p \leq .05$.
*** $.05 < p \leq .10$.

The results of a crude test of simultaneity using both methods are reported in table 2. As in any exercise of this type, our results depend crucially on the adequacy of the identification constraints. Variables excluded from the wage equation include reasons for displacement, whether or not the worker received advance notification of impending redundancy, marital status, and industry and regional dummies applicable to the job at displacement. For the duration equation, the variables excluded are tenure and its square (note, too, that age now enters linearly rather than in spline form). Reasons for displacement, advance notification, marital status, and even industry of job loss are likely to be reservation-wage specific.

The results in table 2 point to evidence of simultaneity bias in equations that incorporate unemployment duration. Looking first at the structural form postdisplacement wage equation, it can be seen that increases in duration reduce postdisplacement wages across all specifications (although quite clearly the effect is measured with imprecision), the negative selectivity coefficient suggesting that yet-to-be-reemployed workers have relatively higher predicted earnings than their currently employed counterparts, controlling for duration effects. As in the OLS case, however, selectivity bias is not indicated for the wage-difference equation shown in the last column. For the duration equation, an increase of 10% in offered wages decreases unemployment duration by between 7.9% and 14.6%. On the other hand, a 10% increase in the predisplacement wage increases duration by between 6.2% and 9.1%. The selectivity argument now has a very simple interpretation: the currently unemployed have longer durations of unemployment than earlier experienced by the currently employed. The basic inferences from the duration equation are as follows. First, predisplacement wages likely raise reservation wages and thus increase duration. Second, the higher postdisplacement wages, given the predisplacement wage, the shorter is the unemployment spell. This outcome may be the result of either a higher arrival rate of job offers or simply higher offered wages. Returning to the line of causation running from duration to postdisplacement wages, the suggestion would appear to be that longer durations associated with lower reservation wages, human capital depreciation, and stigma effects dominate productive search.

Although the results provided in table 2 do indicate that simultaneous equations bias attaches to OLS estimates such as those presented in table 1, it is unfortunate that the coefficients on unemployment duration vary so widely across specification (namely from −.02 to −.23). This outcome might reflect either the poor predictive power of the duration equation or the presence of censoring bias and with it the inadequacy of the two-step selectivity correction estimator employed, or both.

Since information on the duration of yet-to-be-completed spells of unemployment is available in our data set, the duration equation can be estimated more efficiently using this additional duration information via a maximum-likelihood method that explicitly incorporates the stochastic nature of censoring. We employ an accelerated failure time model (Kalbfleisch and Prentice 1980), which in this application is equivalent to a Tobit model in which the truncation point is variable. Using this methodology, we duly estimated a reduced-form duration equation. The predicted duration variable was then used as a regressor in the postdisplacement wage equation(s) (Lee 1981). Our model thus conforms to an instrumental variables approach. Maximum-likelihood estimates of the duration equation are given in Appendix D, and instrumental variables estimates of the postdisplacement wage equations containing a predicted-duration variable are presented in table 3. All the results in the reduced-form unemployment-duration equation seem sensible and are more robust than their probit counterparts. In particular, the source of job-loss dummies and advance notification significantly reduce the duration of unemployment. Note also that previous job duration positively influences the duration of unemployment. All the other results are conventional enough with the exception of the statistical insignificance of the coefficient on the predisplacement wage. Note, however, that we are now dealing with a reduced-form equation so that the coefficient on the previous wage compounds its effect on the reservation wage with the effect of its worker quality content on the offered wage.

Table 3
Postdisplacement Wage Equations Using Predicted Values for Unemployment Duration from Maximum-Likelihood Estimates

	Specification					
Variable	(1)	(2)	(3)	(4)	(5)	(6)
CONSTANT	5.358*	5.150*	2.911*	2.406*	.433*	.132
	(43.59)	(39.44)	(17.77)	(13.65)	(3.51)	(1.04)
SCHOOL	.037*	.036*	.019*	.018*	.001	.004
	(6.79)	(5.97)	(3.79)	(3.38)	(.16)	(.73)
AGE:						
20–35	.015*	.019*	.004	.008**	−.007**	−.002
	(4.43)	(5.27)	(1.31)	(2.26)	(2.05)	(.59)
36–50	−.011**	−.013**	−.001	−.002	.009	.006
	(9.97)	(2.05)	(.24)	(.40)	(1.54)	(1.07)
51–65	−.023*	−.034*	−.016**	−.028*	−.009	−.022**
	(2.63)	(3.59)	(2.00)	(3.20)	(1.00)	(2.40)
TENURE1	.013*	.014*	−.001	−.005	−.016*	−.013**
	(2.64)	(2.61)	(.23)	(.10)	(3.09)	(2.40)
TENURE1SQ	−.00046**	−.0005**	−.0001	−.00013	.00025	.00016
	(2.37)	(2.35)	(.60)	(.70)	(1.30)	(.80)
WHITE	.117*	.108**	.023	.025	−.072	−.043
	(2.58)	(2.17)	(.55)	(.56)	(1.59)	(.90)
UNSKILLED	−.127*	−.138*	−.102*	−.101*	−.077*	−.069**
	(4.78)	(4.95)	(4.22)	(3.96)	(2.88)	(2.55)
SMSA	.083*	.111*	.051**	.077*	.018	.049**
	(3.59)	(4.43)	(2.40)	(3.40)	(.78)	(2.03)
SOUTH	−.107*	−.124*	−.047	−.043	.014	.023
	(3.13)	(3.31)	(1.50)	(.127)	(.41)	(.64)
MOVEREG	.011	...	−.034	...	−.078*	...
	(.40)		(1.38)		(2.92)	
MOVEIND	−.220*	...	−.198*	...	−.175*	...
	(8.79)		(8.67)		(6.97)	
MOVEOCC	−.150*	...	−.103*	...	−.055**	...
	(6.05)		(4.53)		(2.20)	
PART TIME	−.839*	...	−.793*	...	−.745*	...
	(12.38)		(12.84)		(10.96)	
YEAR:						
1979	.102**	...	.073***	...	.044**	...
	(2.43)		(1.93)		(1.06)	
1980	.043	...	.018	...	−.007	...
	(1.14)		(.53)		(.19)	
1981	.079**	...	.045	...	.011	...
	(2.30)		(1.44)		(.31)	
1982	.078**	...	.046	...	.014	...
	(2.36)		(1.54)		(.42)	
lnWAGE1	0	0	.497*	.547*	1	1
			(20.42)	(21.01)		
lnDURATION	−.089*	−.075*	−.094*	−.109*	−.099*	−.138*
	(4.87)	(3.49)	(5.68)	(5.62)	(5.45)	(6.66)
LAMBDA	...	−.113***	...	.064	...	.211*
		(1.65)		(1.03)		(3.21)
R^2	.291	.163	.414	.314	.156	.064

* $p \leq .01$.
** $.01 < p \leq .05$.
*** $.05\ p \leq .10$.

The most notable feature of the fitted equations in table 3 is the consistency and strength of the unemployment-duration variable. The direct effect of duration on the postdisplacement wage and the change in wages is strongly negative: a 10% increase in unemployment duration decreases the accepted wage in the ranges .8%–1.4% and .9%–1.0% for the estimates with and without selectivity correction, respectively. Also of interest is the change in the lambda coefficient. Its sign is now positive for the specifications reported in columns 4 and 6 of table 3 and is statistically significant in the wage-difference equation, suggesting that, controlling for the direct effect of duration on wages and individual heterogeneity, the currently unemployed have *greater* wage losses than their employed counterparts. This is in one sense a more realistic result that presumably reflects the likelihood that *both* the duration and the selectivity arguments are picking up the effects of a declining reservation wage on postdisplacement wages. As before, the coefficients of the variables that influence the postdisplacement wage but not the predisplacement wage steadily decline in magnitude (again with the exception of unemployment duration) as the coefficient on the predisplacement wage rises from zero to unity.

One other feature of the wage-difference equation (col. 6 of the table) that merits some attention is the emergence of age-related wage losses over and above that effect produced by the shape of the age-earnings function. The wages of the oldest group fall by 2.8% a year in the postdisplacement job (col. 2 of table 3). The wage difference amounts to a statistically significant 1.8% a year (col. 6 of table 3). These results are largely due to the fact that the industry and occupational dummies are dropped from the equation. In other words, older workers are less likely to locate or be offered jobs in their old industry or occupation.

We turn finally to the overstatement of the loss in earnings resulting from the assumption that the returns to tenure on the predisplacement job are sacrificed in their entirety on displacement. As we have shown, this overstatement is caused by uncritically identifying the tenure coefficient with the return to firm-specific training investments. The exaggeration of firm-specific training losses can be approached in a number of ways. Here, we rely on two measures based on a comparison of the tenure-earnings profile on the old job with the TENURE1-earnings relation constructed for the new job. Our first measure compares the height of the two profiles at various tenure intervals and is the ratio of the computed returns to TENURE1 in the postdisplacement job to the corresponding returns to that tenure on the predisplacement job. Our second measure focuses on the size of the gap between the two functions and thus has a basis in the estimated "wage-change" functions. It is computed as the ratio of the "returns" to TENURE1 in the wage-change equation to the returns to that variable in the predisplacement job. Table 4 reports the results of this exercise for the three specifications of the postdisplacement wage and for

Table 4
Measures of the Overstatement of the Earnings Loss from Assuming *Returns to Tenure on the First Job Are Lost on* Displacement, Selectivity-adjusted Estimates

	Tenure Interval			
Specification	Median (TENURE1 = 3)	Mean (TENURE1 = 4.564)	Ninetieth Percentile (TENURE1 = 12)	Ninety-fifth Percentile (TENURE1 = 17)
Ordinary least squares:				
Ratio *a*	.473	.464	.404	.343
Ratio *b*	.527	.538	.606	.678
Two-stage least squares:				
Ratio *a*	.434	.425	.371	.315
Ratio *b*	.366	.380	.470	.565
Instrumental variables approach:				
Ratio *a*	.473	.465	.413	.359
Ratio *b*	.450	.461	.527	.595

NOTE.—Ratio *a* is obtained by dividing the computed return to TENURE1 in the postdisplacement wage equation by the return to tenure in the predisplacement wage equation. Ratio *b* is obtained by dividing the computed loss to TENURE1 in the wage-change equation by the return to tenure in the predisplacement wage equation.

four tenure intervals. Thus, for a worker with the mean tenure of around 4.5 years, TENURE1 yields a return on the postdisplacement job that ranges between 42.5% and 46.5% of the return received on that tenure on the predisplacement job. Alternatively, using the second ratio, the size of the gap between the TENURE1-earnings profiles in the two jobs amounts to between 38.0% and 53.8% of the returns to tenure on the predisplacement job. Evidently, considerable overstatement of wage loss is involved in conventional estimates (e.g., Hamermesh 1987). The degree of overstatement falls as we read across the table from left to right, but remains sizable.

V. Conclusions

Our purpose has been to uncover the determinants of relative changes in earnings associated with involuntary job separations. Our analysis has sought in particular to isolate the effects of previous job duration and the length of the intervening spell of unemployment.

An important result is that tenure on the first job is associated with higher earnings in the postdisplacement job. This outcome could be produced by a job-shopping strategy (Jovanovic 1979), by the location of the individual on the rising portion of his general-training investment profile, or, more narrowly, by a switching phenomenon as might occur when workers who anticipate their impending displacement substitute general- for specific-training investments. Each of the above possibilities points to a positive correlation between job duration and earnings growth and each escapes detection in the conventional Mincerian earnings function.

Our methodology and empirical results together provide a strong indictment of approaches that, in computing earnings losses, focus on the sacrifice of firm-specific investments in human capital obtained from the coefficient on tenure in the predisplacement wage equation. While there may indeed be significant losses associated with yet-to-be fully depreciated specific-training investments (some hint as to which may be provided by the strongly negative coefficients attaching to changes in industry and occupation), the extent of such losses cannot be gauged from the tenure coefficient since the latter compounds a number of different effects. We have sought, however, to provide some measure of the overstatement of losses on firm-specific investments that results from assuming that the coefficient on tenure in the predisplacement wage equation represents a pure return to such investments.

Spell length of unemployment emerges as a potent source of earnings loss. Identifying the role of duration is difficult because of simultaneity: not only does unemployment duration influence postdisplacement wages via productive search, a declining reservation wage, and human capital depreciation, but also the postdisplacement wage affects duration via the wage-offer distribution and the arrival rate of job offers. The former effects seem to be of significance, but we have no way of gauging the component magnitudes of each. Nevertheless, we have been able to identify the direct effect of unemployment duration on wages, and we report an even stronger depressing effect than obtains in the OLS regressions. Our best guess is that a 10% increase in unemployment duration lowers accepted wages by about 1%. We note parenthetically that the coefficient on duration progressively increases in absolute magnitude as the predisplacement wage enters the postdisplacement wage equation with coefficients of 0, .5, and 1. The reverse is true for the coefficients on almost all the other arguments that influence the postdisplacement wage only, which seems to point to the presence of unobserved individual heterogeneity.

Finally, some data limitations should be noted. To the extent that the returns to firm-specific investments take the form of fringes, we will have understated the losses to such investments. On the other hand, overstatement is implied by the existence of compensating differentials for job instability. In neither case do our data allow us to address these considerations. Moreover, the absence of *direct* information on job availability and the arrival rate of job offers is an additional lacuna of the present treatment. Nevertheless, the estimated effects of the postdisplacement wage on duration at least provide some hint as to the importance of these effects. For the future, each of these limitations should be tackled. One of the most intriguing questions that remains to be answered is whether chance, rather than choice, dominates in determining unemployment duration (Mortensen and Neumann 1984).

Appendix A

Table A1
Descriptive Statistics

Variable	Full Sample Mean	Full Sample SD	Reemployed Sample Mean	Reemployed Sample SD
AGE	35.615	10.431	35.131	9.975
SCHOOL	12.270	2.567	12.542	2.447
TENURE1	4.679	5.955	4.564	5.779
lnDURATION	...	...	2.128	1.752
lnWAGE1	5.885	.483	5.907	.469
lnWAGE2	...	...	5.693	.592
lnWAGE2-lnWAGE1	...	...	−.214	.544
WHITE	.889	...	.920	...
MARRIED	.694	...	.728	...
SOUTH	.120	...	.132	...
SMSA	.551	...	.550	...
PART-TIME	...	...	.029	...
UNSKILLED	.383	...	.358	...
MOVEOCC	...	...	.551	...
MOVEIND	...	...	.614	...
MOVEREG	...	...	.245	...
NOTIFIED	.553	...	.520	...
CLOSED	.399	...	.426	...
ABOLISH	.105	...	.113	...
YEAR:				
1979	.097	...	.117	...
1980	.141	...	.162	...
1981	.193	...	.217	...
1982	.275	...	.285	...
N	3,223		2,007	

Appendix B

Table B1
The Predisplacement Wage Equation

Variable	Coefficient
Constant	4.939*
	(70.65)
SCHOOL	.036*
	(8.71)
AGE:	
20–35	.019*
	(7.71)
36–50	−.017*
	(3.65)
51–65	−.023**
	(2.49)
TENURE1	.029*
	(6.91)
TENURE1SQ	−.00074*
	(4.57)
WHITE	.193*
	(5.39)
UNSKILLED	−.060*
	(2.84)
SMSA	.073*
	(3.77)
SOUTH	−.129*
	(4.52)
YEAR:	
1979	.092*
	(2.66)
1980	.076**
	(2.41)
1981	.083*
	(2.86)
1982	.065**
	(2.38)
R^2	.174

NOTE.—Absolute t-values are given in parentheses.
* $p \leq .01$.
** $0.1 < p \leq .05$.

Appendix C

Table C1
Probit Reemployment Estimates

Variable	Coefficient
Constant	−2.110*
	(5.93)
SCHOOL	.088*
	(7.67)
AGE:	
20–35	−.009
	(1.14)
36–50	.013
	(.98)
51–65	−.055*
	(2.98)
TENURE1	−.008
	(.68)
TENURE1SQ	.0001
	(.19)
MARRIED	.348*
	(6.03)
WHITE	.557*
	(6.97)
UNSKILLED	−.169*
	(3.02)
SMSA	.050
	(.87)
CLOSED	.193*
	(3.42)
ABOLISH	.177***
	(1.94)
NOTIFIED	.042
	(.82)
lnWAGE1	.118***
	(1.92)
Log-likelihood	−1,626.9

NOTE.—Asymptotic absolute t-values are given in parentheses. The equation also included eight regional dummies, four year dummies, and seven industry dummies.
* $p \leq .01$.
** $.01 < p \leq .05$.
*** $.05 < p \leq .10$.

Appendix D

Table D1
Maximum-Likelihood Estimates of the Reduced-Form Unemployment-Duration Equation

Variable	Coefficient
Constant	4.97*
	(.530)
SCHOOL	−.120*
	(.017)
AGE:	
20–35	.021***
	(.012)
36–50	−.024
	(.019)
51–65	.050***
	(.028)
TENURE1	.044**
	(.017)
TENURE1SQ	−.00011***
	(.00006)
MARRIED	−.544*
	(.088)
WHITE	−.877*
	(.133)
UNSKILLED	.362*
	(.086)
SMSA	.031
	(.084)
CLOSED	−.570*
	(.083)
ABOLISH	−.248***
	(.131)
NOTIFIED	−.216*
	(.075)
lnWAGE1	−.038
	(.092)
Sigma	1.982
	(.031)
Log-likelihood	−5305.92

NOTE.—Asymptotic SEs are in parentheses.
* $p \leq .01$.
** $.01 < p \leq .05$.
*** $.05 < p \leq .10$.

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