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# Layoffs and Wage Growth of Male Household Heads

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More than 30% of the males 14–63 years of age who were interviewed for the University of Michigan "Panel Study of Income Dynamics" experienced at least one dismissal over the years 1969–79.[1,2] Their experience suggests that the impact of layoffs on the financial well-being of workers deserves study. The aim of this study is to examine the impact of layoffs on the rate of wage growth, particularly through the effect of layoff-induced losses of firm specific human capital.

A layoff can influence the financial well being of an individual in at least two ways. First, a laid off worker who does not immediately start a new job will incur a loss in earnings for the period in which he did not work. The problem of unemployment (and of under-employment) while searching for a new full-time job is well known. Secondly, when re-employment is found it may be at a lower hourly wage than the individual previously earned. The estimation of this second effect, i.e., the effect of layoffs on the rate of growth of hourly wages rates, is the focus of this paper.[3] An important aspect of any negative effects of a layoff is their length. If layoffs decrease the rate of wage growth, it is important to know whether the damage is permanent, or, if not, how long it takes an individual to recover. Further, it is relevant to know how the impact of a layoff depends on a worker's personal characteristics, e.g., age, education, race and current job tenure. Both of these issues are addressed in this paper.

The job search literature has detailed two mechanisms by which dismissals might depress wage growth.[4] Since some dismissals are fires for cause, prospective employers may suspect dismissed workers of being less productive than average. If so, such workers will face a lower wage distribution then comparable job seekers who have not been laid off.

Dismissal also implies the loss of some specific on-the-job-training (OJT). The more firm specific OJT a worker has with the current firm, the larger his current wage relative to his alternative wage with other firms and the larger the potential decrease in log wage growth from a layoff.[5]

While recognizing the potential impact of both of these mechanisms, this paper focuses on the second one. A potential problem with measuring the effect of layoffs on the rate of wage growth is that job mobility may be correlated with unobserved differences among workers in firm specific OJT investments since the incidence of mobility is expected to be negatively related to a worker's specific OJT. Negative correlation between layoffs and unobserved differences in OJT would cause spurious negative correlation between wage growth and layoffs, because layoff prone individuals tend to have low wage growth regardless of their frequency of job change. One hypothesis to be tested in this paper is that individual differences in specific

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OJT accumulation can be proxied by the individual's personal characteristics and, therefore, that the impact of a layoff on wage growth depends on these same characteristics.

Several authors have written about measuring OJT, and its interaction with job mobility. Mincer and Jovanovic (1981)[6] hypothesize that individual differences in firm specific investment and quality of the job-worker match produce differences in mobility behavior and wages. They attempt to capture the differences in firm specific investment by including in their wage and job separation regressions, controls for variables hypothesized to be related to OJT, e.g., education and prior mobility. This paper undertakes to explicitly test the effects of education on investment behavior by disaggregating the wage growth regressions by education; it also undertakes to interrelate the mobility variables with prior tenure (which is inversely related with propensity to change jobs and presumably, directly related to the tendency to invest in specific OJT). The regressions are also disaggregated by race and age to allow for differences in OJT accumulation by age and race.

Borjas and Bartel (1981)[7] also tested for differences in investment in specific OJT between mobile and non-mobile individuals; their method was to examine the relationship between mobility and wage growth prior to the mobility, using wage growth prior to the period of a potential job change as a proxy for OJT investment. They found no solid evidence that wage growth prior to the turnover period was different for (most) mobile individuals than for non-mobile individuals. In particular, there was no evidence that workers who were laid off experienced lower pre-turnover wage growth (and therefore lower pre-turnover training investment) than non-mobile workers.

On the other hand, Borjas and Rosen[8] (1980) by correcting for selection bias in the division of workers into mobile and non-mobile individuals show that mobile and non-mobile individuals differ; non-mobile individuals gain by staying on their current job and mobile individuals gain from job changes. However, when the mobile individuals are divided into quitters and those laid off, it appears that those laid off do not gain from a job change.

This study will focus on the effects of layoffs on wage growth although quits and heterogeneity behavior of individuals are also controlled for. My method is to disaggregate the wage growth regressions by some variables I expect to be correlated with specific OJT and to interrelate mobility variables with tenure in the specification of one of the regressions. The rest of this paper is organized as follows. In Part I, it is shown that the effect of layoffs on a worker's wage growth depends on those personal characteristics which are expected to capture heterogeneity in investment behavior. This result follows from 1) a model of OJT accumulation in which it is hypothesized that the individual's OJT accumulation varies with certain personal characteristics and 2) a model of wage growth in which it is shown that the effect of layoffs on wage growth is directly affected by the amount of specific OJT accumulated by the worker on the job from which he was laid off.

Together these models imply that individuals with characteristics associated with large investments in specific OJT, may, cet. par., be expected to suffer large losses if they are laid off.

Part II, presents some empirical evidence about the effect of layoffs on wage growth, and on how the individual's age, education, race and tenure prior to the layoff influence this effect. It is established that layoffs that occur after four or more years on the job significantly decrease the rate of wage growth and, further, that young whites suffer a decrease in wage growth subsequent to a layoff. However, there is no evidence that young blacks suffer a decrease in wage growth subsequent to a layoff or that layoffs decrease wage growth for individuals with



fewer than four years of tenure. The effect of age on the loss from a layoff depends on the individual's educational level. Conclusions are presented in Part III.

## UNDERLYING MODELS

Economic theory suggests that the loss from a permanent layoff is directly related to firm specific OJT acquired on the "old job." Unfortunately, there is no large data set containing measures of OJT to test this hypothesis. Consequently, I develop proxies for OJT investment using the model developed below. A subsequent section will present a model of wage growth in which the effect of a layoff is shown to depend on prior OJT accumulation.

### *A Model of Firm Specific OJT*

Assume there exist n demographic groups whose members are identical in age, education, job tenure and race within groups. Assume also, following Becker (1973) that workers share the cost as well as the benefits of specific training with employers.[9] All members of any group have identical demand schedules that relate the quantity of OJT desired to the proportion of OJT costs paid for by workers in that group, who face identical supply schedules that relate the quantity of specific OJT employers are willing to supply to the proportion of OJT costs the firm pays.

The intersection of the aggregate supply and demand schedules determines the equilibrium quantity of OJT accumulated for any individual in time t. The equilibrium amount of OJT accumulated by an individual is modelled as a function of the characteristics of members of his group so that heterogeneity in OJT accumulation can be proxied by the (hopefully measurable) heterogeneity of individual characteristics.

More formally, it is hypothesized that

$$C_t = g(X_t)$$

where $C_t$ = the quantity of OJT accumulated by the individual on his current job as of time t and $X_t$ is a vector of personal and demographic characteristics of the individual in time t that affect either the individual's demand or supply of OJT. Components of $X_t$ that are positively related to investment in specific OJT are expected, cet. par., to be associated with a decline in wage growth following a layoff. Clearly, many variables could be included in $X_t$; discussion of those on which the empirical work will focus occurs in Section II below.

### *A Model of Wage Growth*

Wage growth between $t_1$ and $t_2$ is hypothesized to be

$$\Delta \ln w = f(\Delta Z, C_i, L01) \tag{1}$$

where

$\Delta \ln w$ = the change in the individuals's log wage between $t_1$ and $t_2$.[10]

$\Delta z$ = a vector whose components are changes in control variables for the individual between $t_1$ and $t_2$.

L01 = 1, if the individual experienced a permanent layoff between $t_1$ and $t_2$.

Further, let $C_i$ be defined as specific OJT acquired with the $t_1$ firm prior to $t_2$. $C_i$ impacts the effect of a layoff on OJT stocks in $t_2$, in the following way: If, in $t_2$ the individual is still

employed by the same firm as in $t_1$, $C_i$ is assumed to affect the individual's productivity, and therefore, his wage rate, in $t_2$. Otherwise, $C_i$ has no effect on the wage rate in $t_2$; clearly, part of the "cost" to the individual of being laid off is the loss of $C_i$. Using a semi-log specification for $\Delta \ln w$, $\Delta \ln w$ can be modelled more explicitly:

$$\Delta \ln w = a_0 + a_1 S \cdot C_i + a_2 L01 + a_3 \Delta Z + \epsilon \tag{2}$$

where $S = 1$ if the individual was with the same firm in $t_2$ as in $t_1$, otherwise $S = O$.
(Note that $S = O$ for quitters as well as for those laid off).
$a_1$ captures the return to specific OJT that accrues to the individual provided he did not experience a job separation between $t_1$ and $t_2$.
$a_2$ captures the effect of a layoff on wage growth aside from the loss of specific OJT (e.g., the negative signalling effect). Note that the effect of a layoff on the rate of wage growth due to loss in specific OJT is captured by setting $a_1 \cdot SC_1$ equal to O for individuals who are laid off.
$a_3$ captures the rate of return to change in the control variables.

Equation (2) can be transformed as follows:

$$\Delta \ln w = b_0 + b_1 \cdot LO1 + b_2 \cdot \Delta Z + u \tag{2.1}$$

where

$$b_0 = a_0 + a_1 \cdot C_i$$
$$b_1 = a_2 - a_1 C_i \text{ and } b_2 = a_3.$$

That is, $b_1$, the loss from a layoff from a job held in $t_1$, includes the loss of the return to specific OJT ($a_1 C_i$) plus any additional effects ($a_2$) attributable to a layoff. Note that $b_1$ depends on $C_i$, the amount of firm specific OJT the individual accumulated on the job as of $t_1$, prior to his employment in $t_2$. Although it is impossible to measure $C_i$ directly, some of the determinants of $C_i$ (i.e., of some of the $X_t$ discussed above) are measurable as discussed below in Section 2. It is hypothesized that $b_1$ is affected by the variables that affect $C_i$. This claim is tested in the empirical work.[11]

### *Data Specification and Empirical Results*

The data set made available in the *Panel Study of Income Dynamics*, follows each family head through the years 1968–1982. The sample consists of 830 male household heads from the PSID who:

1. were 25–53 years old in 1968
2. had no serious health problems in 1968 ($t_1$) or 1980 ($t_2$)
3. had a non-zero wage rate in both 1968 and 1980
4. worked at least 40 weeks in each year 1968–1980.

Data was gathered on layoffs (and quits), wage rates and other relevant variables for each head in each year, allowing calculations of both wage growth and changes in the control variables and tracing of mobility behavior over the period $t_1$ to $t_2$.

The regressions are based on equation (2.1) with $(\ln w_2 - \ln w_1)$, as the dependent variable, including as independent variables, the levels of the individual's $t_1$ characteristics where appropriate as well as changes in these variables and mobility variables. These level variables are hypothesized to be related to the individual's ability to accumulate OJT (general and specific) between $t_1$ and $t_2$, and therefore to serve as a proxy for growth in OJT. Recall that



$C_i$ is proxied by measurable personal variables, (the components of $X_t$). The variables used are length of job tenure, age, race and education.

Focus is directed to the age and tenure variables because human capital theory provides strong predictions about the effect of age on post-mobility human capital investment, and therefore a clue to how age affects recovery from a layoff and, theory predicts unambiguously that tenure should be correlated with firm specific OJT prior to a layoff, which suggests that long tenure workers should have the most difficulty recovering from a layoff. The effect of education is of particular interest because education is commonly presumed to provide a worker with versatility in the market and therefore with some financial security. The predicted effects of race on the loss from a layoff are ambiguous, since non-whites are likely to face discrimination in hiring and training both prior to a layoff and afterwards. However, determining empirically, whether this discrimination makes blacks particularly vulnerable to layoffs is of interest.

The relationship between each of these four variables and the loss caused by a layoff depends on

1) the relationship of the variable to pre-turnover specific OJT investment because the greater this specific OJT investment, the greater the loss from a layoff (cet. par.)

2) the relationship of the variable to post-turnover rate of wage growth. The greater the post-turnover rate of wage growth, the less the loss from the layoff.

Consider, first, the effect of current job tenure on OJT. The longer the worker's job tenure, the more time he has had to accumulate specific OJT on the current job. Also, the return to OJT accumulated in any year on a job is positively related to expected total job duration. Assuming that current job tenure and expected total job duration are positively related, it is expected that total accumulation of specific OJT on the job as of year t is positively related to current job tenure in year t. Further, since specific OJT investment discourages job mobility, I expect those who have accumulated the most specific OJT in past years to have the longest current job tenure. Consequently, I hypothesize that total OJT accumulated on the job as of time t is positively related to tenure in t.

On the other hand, the relationship between tenure and post-layoff log wage growth is ambiguous. If tenure is positively related to the motivation and ability level of the worker, tenure would be expected to be positively correlated with post-turnover log wage growth. However, to the extent that tenure is correlated with the proportion of OJT that is specific, one might expect those laid off after acquiring long tenure to be limited to secondary sector (low wage growth) jobs immediately after a layoff.

Age affects OJT investment through its influence on expected remaining time in the labor force. The older a worker is, the fewer years, cet. par., he has left in the labor force. This tends, on the one hand, to discourage older workers from (pre-layoff) OJT investment since profitability from such investment is directly related to the number of years of work life the individual has left on the current job, and time left in the labor force.[12] On the other hand, having fewer remaining years of work life discourages older workers from potential quits which tends to increase the number of years the individual will be on the current job (holding remaining time in the labor force constant) and this effect should increase the return to pre-layoff specific OJT for older workers. Since these two effects work in opposite directions, it is not obvious which one will predominate. Thus, the effect of age on new OJT investment is ambiguous.[13]

Stronger predictions can be made about the effect of age on the post-layoff rate of wage

growth. The older the worker is, the less incentive there is for the worker to engage in extensive post-turnover job search or most other types of human capital formation, and therefore, the smaller his expected post-layoff rate of wage growth.

The effect of education on OJT is ambiguous. On the one hand, education and OJT may be complements. The more knowledgeable an individual is, the greater the return to resources devoted to learning new skills, and thus the greater incentive he has to acquire specific OJT. However, education may also be a substitute for specific OJT (Blau and Kahn 1981).[14] Education is likely to be positively correlated with potential wage growth after re-employment, because education increases productivity in itself, increases the rate of absorption of new skills, and possibly increases an individual's efficiency at job search.

Lastly, it is hypothesized that blacks receive less training than whites in accordance with the findings of Robert Flanagan (1974).[15] This is consistent with the institutional theory of the labor market which suggests that some groups (e.g., blacks and women) tend to be crowded into unskilled jobs that require little training. Further, discrimination may increase the cost of OJT and reduce the return on investments in human capital for blacks even on skilled jobs, if promotion opportunities are limited for them. Blacks are also likely to face discrimination in finding re-employment after a layoff and therefore to have lower post-layoff wage growth than whites.

To summarize, the above discussion implies that whites acquire, cet. par., more OJT than blacks, that specific OJT is positively related to current job tenure and that the effects of age and education on OJT are uncertain.

On the other hand, the post-layoff rate of wage growth is expected to be lower for blacks and older workers than for whites and younger workers, and to be positively correlated with education. The relationship between pre-layoff tenure and post-layoff wage growth is ambiguous.

The logic of the empirical work is that worker characteristics associated with heavy OJT accumulation should also be associated with a larger loss in wage growth if the individual is laid off, since the more specific OJT a worker has, the greater the loss of specific OJT from a job change. However, worker characteristics that are correlated with high post-turnover wage growth will be associated with a small loss in wage growth.

None of these four factors (tenure, age, education or race) is unambiguously correlated with both higher pre-layoff specific OJT and lower post-layoff log wage growth (both of which contribute to vulnerability to loss in wage growth after a layoff). However, since the sign of the correlation of post-turnover wage growth to tenure is ambiguous, the positive effect of tenure on pre-turnover OJT is expected to dominate and so those with the most tenure are expected to lose the most from a layoff.

For a similar reason (the ambiguity of the effect of age on specific OJT accumulation) the post-turnover wage growth effect should predominate when workers are distinguished by age and older workers are expected to lose the most from a layoff. It is also expected that, more educated workers will lose less from a layoff, and that the effect of race will be ambiguous.

The validity of the model is tested by disaggregating the regressions sample by age, education and race. The layoff dummy introduced above is neither the only possible specification of mobility or necessarily the ideal one. First, it does not take account of the timing of the turnover. Persons laid off 10 years ago may have fresh OJT by $t_2$ and may have recovered from any negative signalling. Therefore, layoffs that happened 10 years ago should not have the same effect on wage growth as those which happened last month,[16] and this is taken into account in empirical work.



It is also necessary to test the hypothesis that the effect of turnover on log wage growth is dependent on tenure prior to turnover. Several studies use tenure to proxy OJT,[17,18] but its effect on wage loss from layoffs has never been studied.

The regressions reported below use the following specifications for the layoffs variables.[19]

### *Alternate Layoff Specifications*

One possible specification differentiates among layoffs according to their timing[20]:

LOE1 = 1 if the respondent had one or more *early* layoffs (1969–1973)
LOL1 = 1 if the respondent had one or more *late* layoffs (1974–1979)

The rationale for separating those layoffs that occurred in the first half of the survey period from those that occurred in the second half is that workers may "recover" from a past layoff as new OJT replaces firm specific OJT lost through a layoff and "negative signalling" is eroded by new work experience. Recent layoffs should have a more dramatic effect on $w_2$ (the wage rate in 1980) and therefore on ln ($w_2$). LOL1 is expected to have a negative coefficient. LOE1 should also have a negative coefficient, but one smaller in absolute value than LOL1.

A second layoff specification differentiates layoffs by 1968 job tenure[21]:

LO1S = 1 if the respondent had one or more layoffs (1969–1979) and less than 2 years of job tenure in 1968
LO1MED = 1 if the respondent had one or more layoffs (1969–1979) and 2–3 years of tenure in 1968
LO1LG = 1 if the respondent had one or more layoffs (1969–1979) and 4 or more years of job tenure in 1968

Recall that firm specific OJT is expected to be positively related to current job tenure. Further, institutional factors may cause senior workers to have more to lose from a layoff. The coefficients of LO1S, LO1MED and LO1LG are expected to be negative and in ascending order of absolute value.

Independent variables include layoff variables plus variables controlling for quits, union status, marital status, health status, changes in these variables 1968–1980, education, race and age (where appropriate).[22]

## FINDINGS

For each regression specification, the results are stratified by age cohort (by whether the individual was less than 38 years old in 1968 or at least 38). Part A of each table displays regression results for the younger age group in column one; for the younger group stratified by education in the next three columns, and for the younger group stratified by race in the last two columns.[23] Part B of each table displays analogous results for the older group.

Tables 1 and 2[24] present the regressions results for the layoffs variables only. Information on the coefficients of all other variables in each regression is available from the author upon request.

### *Early vs. Late Turnover*

Table 1 includes the results for specification A; that is for regressions in which turnover is distinguished by when it occurs. As expected, in general, recent layoffs appear to have a negative effect on the rate of wage growth.[25] The coefficient of the dummy variable denoting a

**TABLE 1**
Regressions Results for Layoffs Distinguished by Timing (t statistics are in parentheses)

| | (1) Young n = 421 | (2) Young With <H.S. n = 136 | (3) Young With H.S. n = 135 | (4) Young With Coll n = 81 | (5) Young Whites n = 307 | (6) Young Blacks n = 100 |
|---|---|---|---|---|---|---|
| | (Part A) Young Workers (23–37 years of age in 1968) | | | | | |
| LOE1 | .079 | .098 | .190* | −.248 | .038 | .196 |
| | (1.269) | (.956) | (1.870) | (1.010) | (.516) | (1.489) |
| LOL1 | −.162*** | −.120 | −.254** | −.167 | −.200*** | −.135 |
| | (2.365) | (1.185) | (2.007) | (.771) | (2.385) | (1.000) |
| | Old n = 409 | Old With <H.S. n = 186 | Old With H.S. n = 103 | Old With Coll n = 72 | Old Whites n = 307 | Old Blacks n = 93 |
| | (Part B) Older Workers (38–53 years of age in 1968) | | | | | |
| LOE1 | −.045 | −.090 | −.109 | −.029 | −.077 | −.047 |
| | (.587) | (.707) | (.740) | (.162) | (.859) | (.269) |
| LOL1 | −.131* | −.164* | −.095 | .230 | −.107 | −.236 |
| | (1.925)* | (1.671) | (.641) | (.742) | (1.468) | (1.222) |

*coefficient significant at the 10% level
**coefficient significant at the 5% level
***coefficient significant at the 1% level

recent layoff is negative for all but one regression in Table 1. A layoff within the past six years appears to reduce wage growth by 16% for the younger cohort (Part A, column 1) and 13% for the older group (Part B, column 1). The difference between the coefficients of the old and young groups is insignificant at the 10% level. When the younger workers are further disaggregated by educational level, recent layoffs appear to significantly (at the 10% level) depress the rate of wage growth for those with exactly 12 years of education (Part A, column 3). Among the older workers, it is the high school dropouts (Part B, column 2) who are significantly hurt by recent layoffs.

The insignificance of the recent layoff coefficients for young high school dropouts may be due to the tendency of employers to attach a stigma to young high school dropouts, and therefore to confine them to jobs that do not require specific training. Their relative lack of specific OJT renders them less vulnerable to a wage decrease from a layoff than members of their cohort with more specific training. Since fewer than half of the older workers are high school graduates, lack of a high school diploma does not automatically make a member of this cohort too poorly educated to benefit from specific training.

Both types of layoff decrease wages for every group of older workers although most of the coefficients are insignificant. Tentatively, it may be concluded that older workers have trouble recovering from past layoffs.[26] These results are consistent with the expectation that older workers cet. par., acquire less post-turnover human capital than younger workers (and thus recover more slowly from a layoff).



Below, in Table 2, layoffs are distinguished by tenure acquired as of 1968 on the 1968 firm.

Table 2 shows regression results for specification 2 in which layoffs are distinguished by the tenure of the respondent in $t_1$ (1968). From these regressions, it appears that, for both age groups, layoffs decrease wage income only if they occur after at least four years of tenure. Such layoffs significantly decrease wage growth for the younger group as a whole (by 16% according to the regression) and for the older workers (by 13%). When disaggregating the younger group of workers by education and then by race, it appears that layoffs after four or more years of tenure significantly decrease the rate of wage growth for the more privileged groups, ie, young white, young high school graduates and college graduates. Among older workers, it is high school graduates who appear to suffer the most from lay-offs after four or more years of tenure. Note that the interaction of tenure with layoffs controls, to some extent, for unobserved characteristics that are correlated both with presence or absence of mobility and with wage growth. That is, although it is possible that mobile individuals might have experienced low wage growth even without mobility, because their tendency to be mobile means that they tend to have small investments in human capital, these results suggest that it is the individuals who tend to be fairly non-mobile (i.e., those with the firm for at least four years prior to mobility) who suffer a decrease in wage growth from a layoff.

**TABLE 2**
Regressions Results for Layoffs Distinguished by 1968 Job Tenure

| | Education Level | | | | Race | |
|---|---|---|---|---|---|---|
| | (1) Young | (2) Young With <H.S. | (3) Young With H.S. | (4) Young With Coll | (5) Young Whites | (6) Young Blacks |
| | (Part A) Younger Workers | | | | | |
| LO1S | −.061 | −.112 | −.130 | .386 | −.111 | −.047 |
| | (.736) | (.829) | (.954) | (1.480) | (1.091) | (.278) |
| LO1MED | .048 | −.033 | .314** | −.218 | .024 | .213 |
| | (.523) | (.217) | (2.257) | (.791) | (.218) | (1.085) |
| LO1LG | −.164* | −.025 | −.268* | −.700** | −.183* | −.112 |
| | (1.800) | (.184) | (1.663) | (2.192) | (1.683) | (.567) |
| | (1) Old | (2) Old With <H.S. | (3) Old With H.S. | (4) Old With Coll | (5) Old Whites | (6) Old Blacks |
| | (Part B) Older Workers | | | | | |
| LO1S | .043 | −.035 | .549* | .227 | −.007 | .067 |
| | (.404) | (.227) | (1.927) | (.838) | (.057) | (.277) |
| LO1MED | −.058 | .126 | −.263 | .130 | −.069 | .038 |
| | (.448) | (.556) | (1.357) | (.249) | (.477) | (.103) |
| LO1LG | −.130* | −.183* | −.305 | −.028 | −.102 | −.347 |
| | (1.655) | (1.635) | (2.030) | (.114) | (1.228) | (1.593) |

## CONCLUSIONS

The model presented here predicts that certain personal characteristics (e.g. age, education, race and tenure) influence the effect of a layoff on wage growth (e.g., the coefficient $b_1$ from the wage growth model in Part I) because of their effects on

1) the specific OJT acquired on the old job prior to the layoff, (and therefore lost as a result of the layoff) and/or

2) wage growth after re-employment.

The empirical results are definitive for race and tenure, while the evidence is mixed for age and education. Our prior expectations were that, on average, blacks had worst post-layoff prospects than whites, but also less specific OJT to lose from a layoff. The regressions presented here suggest that whites, but not blacks, appear to suffer a significant decrease in the rate of wage growth as a consequence of a layoff.

The prior expectation for tenure was that individuals with long tenure tend to have large quantities of OJT as well as other seniority privileges and that, consequently, these individuals will be hard hit by a layoff. As expected, pre-layoff tenure is directly correlated with the loss from a layoff.

The expected relationship between education and OJT was ambiguous, but more educated individuals were expected to have better post-layoff re-employment prospects. The results showed that individuals with average education for their age group lost the most from a layoff, presumably because they lost more OJT than other groups.

Recent layoffs appear to cause more damage than those in the distant past. There is some weak evidence that layoffs prior to 1974 hurt older workers, but no significant results for these earlier layoffs.

Judging from these results, the possibility must at least be considered that some of those who lose the most from a layoff do so because they are relatively well off prior to the job separation. The results reported here should not be interpreted as suggesting abandonment of, for instance, young high school dropouts and blacks who appear to lose little from a layoff because their jobs provided little in terms of training and income in the first place.

However, help is indicated for individuals who are laid off after acquiring at least 4 years of tenure. These individuals were found to suffer a significant decrease in the rate of wage growth. Subsequent to a layoff, those from declining industries may be helped by retraining and or relocation assistance as well as income support. Older high school dropouts could possibly benefit from completion of high school, particularly if they are sufficiently short of retirement. Younger high school graduates could be encouraged and assisted in increasing both their training and education.

## NOTES

1. Calculation by the author.
2. A dismissal can either refer to a discharge for cause (firing) or be a result of an employer's decision to decrease the number of workers retained by the firm. Below, I shall refer to both types of dismissals as layoffs.
3. Growth in the log of hourly wage rates, rather than levels of hourly wage rates, is used as the dependent variable because first differencing cancels out unobserved, but constant individual effects that may be correlated with the independent variable.
4. Cooke, William, "Turnover and Earnings: Some Empirical Evidence," *Industrial Relations,* 18(2), (Spring, 1979) p. 220.



5. This assumes that the share of the return to training that accrues to the worker is held constant as the amount of specific OJT varies.
6. Mincer, Jacob and Jovanovic, Boyan, "Labor Mobility and Wages," in *Studies in Labor Markets,* ed. by Sherwin Rosen, (University of Chicago Press, NBR, 1981).
7. Borjas, George and Bartel, Ann in Rosen, ibid.
8. Borjas and Rosen, "Income Prospects and Job Mobility of Younger Man," in Research in Labor Economics, Vol. 3, ed. by Ronald G. Ehrenberg, (JAI Press, Inc., Greenwich, 1980).
9. According to the theory of specific OJT (see Becker's *Human Capital*), firms will, in general, be willing to pay part of the costs of specific OJT.
10. Wage rates may also be related to hours of work although the expected effect of hours on wages is ambiguous; individuals who work shorter hours have higher average fixed (administrative) costs associated with their employment but may have a higher marginal product due to diminishing marginal returns. Hours of work are not included as a control variable, because of the well known simultaneity between the determination of hours worked and wage rates. Hopefully, such variables as marital status and education, which are presumably correlated with hours of work, will pick up some of the effects of hours of work. Further, individuals are excluded from the sample who worked less than 40 weeks in any year of the survey, despite the resulting sample selection bias. Also, since the sample consists of male household heads (only) approximately 98% of the individuals included worked full time.
11. $b_1$ is also affected by any variable that influences post-layoff wage growth (e.g., through the return to post-layoff job search). Variables may influence the effect of a layoff on the rate of wage growth through their impact on post-layoff wage growth as well as through their impact on pre-layoff OJT.
12. The effect of years remaining until retirement on the incentive to invest in OJT is mitigated, to some extent, by the tendency of OJT to eventually depreciate; depreciation implies that the useful life to current training is limited regardless of years left until retirement.
13. Similar conclusions are reached by Bartel and Borjas in Bartel, Anne and Borjas, George, "Specific Training and Its Effect on the Human Capital Investment Profile," *Southern Economic Journal,* 44, (2), (October, 1977), p. 336.
14. Francine Blau and Leonard Kahn, "The Causes and Consequences of Layoffs," *Economic Inquiry,* 1981.
15. Flanagan, Robert, "Labor Force Experience, Turnover and Racial Wage Differentials," *Review of Economics and Statistics,* (56), (November, 1974), p. 529.
16. Jacobsen, in Earnings Loss of Workers Displaced From Manufacturing Industries (Public Research Institute, Arlington, 1976) recognizes this in his study on loss from changing industry and uses dummies for the year of leaving the industry. Data limitations precluded my breaking down layoffs by year.
17. Borjas, G., "Job Mobility and Earnings over the Life Cycle," *Industrial Labor Relations Review,* (1978), pp. 365–376.
18. Borjas and Bartel, "Wage Growth and Turnover, an Empirical Analysis," NBER Reprint, #321, September, 1982, p. 79.
19. The PSID does *not* allow a separation of dismissals into layoffs and fires. Both are coded as the same number for all years. However, all dismissals recorded are permanent. None are temporary layoffs.
20. Three other specifications (beside these reported here) were tried. They were: 1) total number of layoffs during the period 1969–1979, 2) dummies for at least one layoff during the period, 3) total number of early layoffs, total number of late layoffs. The results reported in the paper are the most interesting, but results using the other three specifications are available from the author on request.
21. An alternative (and interesting) specification would be to interact the layoff/quit dummies with job tenure at time of the layoff or quit. I used the specification reported here because job tenure in 1968 is exogenous to any job mobility 1969–1979.
22. I also tried regressions including weeks unemployed as an additional control variable, but the results were uninteresting.
23. The groups disaggregated by education are young workers with less than 12 years of education (column 2), young workers with exactly 12 years of education (column 3) and young workers with at least 16 years of education (column 4). There were no regressions run on young workers with 13–15 years of education due to a sample size of less than 50. (There were no regressions for older workers with 13–15 years of education for the same reason).

24. Chow test for both specifications justified the sample divisions by age but not by race. Technical problems precluded Chow tests by educational level.
25. It should be mentioned that wages may affect turnover as well as the other way around since theory suggests that workers who consistently make more than their marginal product will be dismissed. In particular, Blau and Kahn (1981) suggest that when OJT is imperfectly measured the wage is, in principle, a proxy for the portion of specific human OJT which is owned by the worker which has, in itself no value as a deterrent to layoffs. Alternative theories suggest that the layoff probability is function of total specific OJT (including that owned by the worker) because in any job separation the initiator will compensate the other party for the loss involved in the separation. In this case, as Blau and Kahn observe, layoffs may be negatively related to the wage rate, because, in the absence of perfect measures of total specific OJT, the wage rate is a proxy for total specific OJT. Due to space limitations, the issue of possible simultaneity between layoffs and wage rates is not pursued here. The best I can hope is that I have captured most of the relevant unobserved characteristics in the four variables that are used to proxy OJT.
26. If the true coefficient of LOE1 is O in every equation for the older group the probability of any regression coefficient being negative is .5. If the regressions were independent of each other, the number of negative coefficients out of 6, y, would be a binomial random variable with probability distribution

$$p(y) = (.5)^y (.5)^{6-y}$$

and

$$p(6) = (.5)^6 = .00156$$

So the probability of getting 6 negative values of LOE1 for the older group is less than 1% if the true coefficients are all O.

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