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UNION EFFECTS: WAGES, TURNOVER,
AND JOB TRAINING

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ABSTRACT

This study explores the existence of a net union premium and of the
extent of rationing by quality of the resulting excess supply. The net
union premium was estimated by relating changes in wages to changes in
union status of the same worker in longitudinal panels (NLS and MID), and
by two cross-section wage level regressions, a "prospective" and "retrospective"
which permit more direct observation of selectivity in hiring. Over a
half of the cross-section differential of over 20% for the "same" (stan-
dardized) worker is a net union rent and much of the rest reflects a quality
adjustment in hiring, as measured by wages. This conclusion was less reliable
for older workers.

Subsequent analysis explores the effects of successful union wage
pressure on: quit rates, fringe benefits, wage profiles, and training.

The reduction in quit of union joiners depends on the size of the
net wage premium. Quit rate differentials are also positively related to
the gross, cross-section wage differentials within groups of workers,
classified by location and occupation, less so by industry.

In Section 4, it is hypothesized that the imposition of larger fixed
labor costs (such as fringes) helps to deter employers from preferring
reductions in hours to reductions in men, and it helps to stabilize employ-
ment in the face of fluctuating demand, by a more frequent use of overtime
and of temporary layoffs in the union sector. This hypothesis links the
size of fringe benefits to the union wage gain. An analysis of firms in
70 industries confirms this link.

Union pressure is exerted on the whole tenure profile of wages. The
explicit linking of wage levels to seniority reduces incentives for worker
investment in general (transferable) training. The total volume of training
is indeed reported to be smaller in union jobs, and this is consistent with
the flatter profile.

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Introduction

Growing numbers of empirical studies confirm the prevalence of the following

features which distinguish unionized from non-unionized labor markets: (1) Higher

wages (2)A larger share of fringe benefits in total compensation (3) Lower quit

rates and lesser turnover, and (4) Flatter age-profiles of wages. These findings

appear to hold both in the aggregate, as when more unionized industries are

compared with less or non-unionized, and in micro-data where a variety of

variables are used to control for personal and, less frequently, firm characteris-

tics which might affect these differences.

Traditionally, economists have been concerned with the relative wage and

much less the employment impact of unionism, just as they have been concerned

with price-quantity impacts of product monopolies. On the other hand, students

of industrial relations tend to emphasize non-wage aspects of unionism in the

work setting. Some time ago (1958) H.G. Lewis dichotomized aspects of unionism

into "monopolistic" and "competitive". In the first category are the imposition

of wages above competitive levels,either by union restriction of supply or by

threats of strikes implicit in collective bargaining. In the second are all those

activities of unions and work rules espoused by unions which need not be in-

consistent with competitive wage-setting. However, these "two faces" of unionism

do not correspond to wage and non-wage aspects of unionism: union induced non-

wage conditions may also impose higher costs on employers, inconsistent with competi-

tion. Conversely,union-nonunion wage differentials need not reflect monopoly rents,

but may be merely compensatory, reflecting a faster and more regimented pace of work (Duncan and Stafford, 1980) or higher quality of workers in unionized jobs.

The wage-push hypothesis receives almost little or no attention in analyses of other features of unionized labor markets listed in the first paragraph. Thus with "voice", as a substitute for "exit", unionized workers quit less frequently than others. With longer expected stay, the probability of receiving initially non-vested pensions increases, hence larger pensions (a major part of fringe benefits) are demanded by union workers. Other explanations of larger union fringes run from union democracy which favors the older worker to union management of pension funds as an instrument of power. Finally, the flatter union wage profile is seen as a result of union egalitarian or bureaucratic compression of the wage structure.

Although some or all of these hypotheses may be valid, it is possible to view the union wage pressure as a source, if not necessarily an exclusive one of all of the features I have enumerated. It is the purpose of the present study to test this proposition empirically.

1. Union Wage Gains

Before I proceed to explore the effects of union wage pressure on turnover, fringe benefits, and wage profiles, it is necessary to establish that unions indeed succeed in pushing their wages above competitive levels. A large literature (see review by Parsley, 1980) answers this question positively, by observing a differential in favor of union members after controlling for a large number of worker personal characteristics in wage level regressions. However, unmeasured differences in labor quality may still be responsible for part or all of the wage gap. The question is whether, indeed, the same worker, and not merely his statistical surrogate, receives higher wages in union than in nonunion employment.

If the existence of a net union wage gain is confirmed in this fashion, we would expect its size to be smaller than the cross-sectionally observed union wage differential. This is because above equilibrium wages imply queues on the supply side which must be rationed.[1] Only in the case of probabilistic rationing ("first come, first served") and of rationing by price (union dues) or by discrimination and nepotism would the observed net wage gain equal the gross wage differential observed in the cross-section. But, employers have incentives to reduce the increase in labor costs imposed by the union by systematic rationing, that is by hiring more productive workers from the queue. The increase in labor costs cannot, however, be completely offset in this fashion partly because the marginal cost of screening for quality is positive and because of technological constraints in the production function. Moreover, where unions have a voice in hiring, as in the closed shop and in union hiring halls, there are no obvious incentives for upgrading of labor quality, and rationing is largely probabilistic. Rationing by price (in corrupt unions) or by discrimination and nepotism is outlawed, in principle, but its existence cannot be excluded.

The empirical analysis described below relies on observing wages received by the same worker before and after his change of union status. This approach was used most recently by Duncan and Stafford (1980)[2] and, thus far, most comprehensively by Wesley Mellow (1981). Duncan and Stafford observed wage changes of a small sample of union joiners between 1968 and 1971 in the Michigan Panel Survey of Income Dynamics (MID), and Mellow in two one-year intervals (1974-75 and 1977-78) in the much larger sample of the Current Population Survey (CPS). Large data sets, like the CPS sample, are needed to observe adequate numbers of workers who change union status, since their proportion is quite small. An alternative to using the CPS is pooling of the longitudinal samples. Although the annually surveyed samples in the MID and NLS (National Longitudinal Surveys) are much

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[1] The existence of union induced queues has been inferred in an econometric analysis of Farber and Abowd (1981).

[2] Other references are: Brown (1980), Duncan (1979), Chamberlain (1978), Kenny (1978), and Raisian (1981).

smaller than the CPS, their advantage lies in some of the information not available in the CPS. Especially useful in this context is information on job mobility and job tenure.

In the work to be described I pooled the MID sample in order to relate annual changes in wages of white men over the 10-year period 1968-1978 to changes in their union status. Since the NLS panels contain information on union status in the years 1969 and 1971,[1] I utilized this single interval for the same analysis on the two NLS panels of young white men (who were 17-27 years old in 1969) and older white men (who were 48-64 in 1969). The MID contains all ages, but as in the NLS, I limited the sample to non-students and to a maximum age of 64. For (partial) comparability with the NLS, I also stratified the MID panel into young (less than 30) and older (30-64) subsamples.

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[1] I did not use the less adequate telephone survey of 1970 which also contained questions on union status.

The statistical analysis relies on wage functions, where the dependent variable is the logarithm of hourly earnings, and the independent variables are education (Ed), experience (x) and its square (x^2), length of job tenure (T) and its square (T^2), marital status (Mar), health status (Hlth), local (Lun) and national unemployment rate (Nun). These are the "standardizing" variables. The main focus is on the additional (dummy) variables Uij, where i and j index the first and second period, and have values (0,1), 0 denoting non-union status and 1, union membership. Thus Uoo means non-union both years, U_{01} = union joiner (between the first and second survey), U_{10} = union leaver, and U_{11} = union stayer. These four union status categories were also cross-classified by mover-stayer status, and for movers by form of separation (quit, layoff) and whether moved within or between industries.

Wage level equations were used separately for the first and second year bracketing the changes.[1] Both equations contain the same union dummies. Thus in "year 1" equations, the dummies indicate prospective changes (or continuation) of union status, while in "year 2" they indicate recent change. Both are useful in exploring selectivity in hiring. The specification of wage change equations is derived by taking first-differences of the variables in the level equations. Thus the experience variable becomes $\Delta x = 1$ for all, and its coefficient enters the intercept, but Δx^2 differs with the level of experience. The tenure variable ΔT equals 1 for job stayers but becomes negative (-T, where T is length of job tenure on the preceding job). Correspondingly, ΔT^2 is positive for stayers, but is negative and equals $-T^2$ for movers. Standardization for tenure is important: since wages grow with tenure (experience held fixed), wage change estimates from regressions which omit tenure depend on how long recent movers stayed on the previous job. Put another way, estimates which omit tenure indicate the immediate wage change in moving from one job to another, while the present

[1]In the NLS the first year is 1969, the second 1971. In the MID the pairs of years are adjacent.

specification estimates the wage change from the prior to the current job, at comparable tenure levels. This is, clearly, a more appropriate measure of returns to job mobility, or to change in union status.

Although the pooling of ten periods provides an effectively much larger sample of the MID observations, the smaller NLS sample is in some respects superior: fewer observations are lost due to incomplete responses, and hourly wage refers to the current job rather than to the calendar year as in the MID. The NLS also contains an alternative union status definition in addition to union membership (UM), a question on whether the job is covered by collective bargaining (CB)[1].

In the MID, job and union status changed between surveys, which were taken in the second quarter of each year. Wage changes, however, refer to calendar years. Since we bracketed status changers by adjacent years, if the reported wages are indeed weighted annual figures, they could underestimate the true change by as much as 33%. We checked on the degree of bias in several periods (1976, 1977, 1978) when the question on wages was asked for the current job. The so estimated bias was on average 21%. The bias could be removed by leaving out two years between "year 1" and "year 2". But this would have eliminated a large fraction of job changers, whose tenure at moving is short.

Table 1 shows sample means (i.e. proportions) of the union status categories for the young and old men panels by alternative definition of union status in NLS and in MID. 28.3% of the young men and 36.3% of the old men were union members in 1969. A somewhat larger proportion (33.2% and 39% respectively) were covered by collective bargaining agreements. Union-nonunion turnover is quite large among the young and much smaller among older job movers. A surprising and somewhat puzzling statistic is the number of union status changers who do not change jobs (firms) -- it is as large as the number of union joiners who are movers among the young and even larger among the old. The preponderance of stayers who are union status changers is even stronger in MID, in which the proportion of all movers appears to be smaller than in the NLS.

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[1]However, (UM) is, in effect, a subset of (CB), since only those reporting collective bargaining coverage were asked whether they are union members. The (CB) definition is available in MID only for a few periods, so it could not be utilized in the pooled sample.

Table 1

Sample Proportions in Union Status Categories
NLS and MID

(A) <u>NLS</u>

	Young Men, n = 1160			Old Men, n = 1588	
Movers	UM %	CB %		UM %	CB %
U_{00}	24.3	21.9		6.5	5.9
U_{01}	4.6	4.9		.8	1.0
U_{10}	4.4	5.2		.5	.8
U_{11}	4.4	5.7		2.6	2.7
Stayers					
U_{00}	38.6	36.0		52.2	49.4
U_{01}	4.2	4.0		4.2	4.7
U_{10}	1.5	2.0		2.0	2.8
U_{11}	18.0	20.3		31.2	32.7

(B) <u>MID</u>

	All, n= 9,987	Age $<$ 30, n=3,069	Age $\geqq$ 30, n=6,905
Movers			
U_{00}	7.3	13.4	4.6
U_{01}	1.0	2.0	.5
U_{10}	.9	1.8	.5
U_{11}	1.3	2.0	1.0
Stayers			
U_{00}	56.8	53.6	58.3
U_{01}	2.6	2.7	2.5
U_{10}	2.3	2.4	2.3
U_{11}	27.8	22.1	30.3

Although some stayers may become union members, after a short period in the union shop, or by switching jobs within the firm, or by the firm becoming unionized, the figures for job stayers who change union status appear to be inflated by misreporting or misclassification. This is especially likely in the MID samples, where wages of union status changers who are job stayers are about the same (relative to the base group) before and after the change, (see Table 5).

Table 2 presents estimates of the 1969 to 1971 wage changes in the NLS by union change category, net of the other variables (these are shown in Appendix Table A1), classified by mobility status, age, and definition of union sector (UM: union membership, CB: coverage of wage by collective bargaining agreement). Numbers in Table 2 are regression coefficients, with t-statistics shown in parentheses.

Using point estimates and the union membership criterion, young men who joined unions by changing firms got a 17.6% increase in wages, while older men gained 7.4%. However, by the collective bargaining criterion, the gain was similar for young and old (13.5% vs. 11.6%). The figures for the older men are barely significant. The wage changes are adjusted for inflation and are net of the wage change experienced by the base group of non-union stayers. Firm stayers who report joining unions show smaller gains, with lesser statistical significance. Although it may be advisable to discount (or ignore?) the estimates for stayers, there is still a question whether the figures for job movers should be viewed as the net union premium: even if they did not join a union, young movers between non-union firms (U_{00}) gained 3.5 to 6.4%. If such gains measure the return on costs of mobility, the net profit of joining unions by moving is reduced to between 7% and 14% for the young movers, and is comparable for the older union joiners although the statistical reliability of the estimates for the older group is much weaker.

Table 2

Wage Growth 1969 to 1971 in the NLS

Movers	Young (UM)	(CB)	Old (UM)	(CB)
U_{01}	17.6 (4.0)	13.5 (3.2)	7.4 (1.5)	11.6 (1.2)
U_{10}	-26.0 (5.7)	-28.7 (6.8)	-8.2 (.6)	-20.4 (1.8)
U_{11}	7.7 (1.8)	7.9 (2.0)	6.9 (1.1)	6.6 (1.1)
U_{00}	3.5 (1.2)	6.4 (2.3)	-3.2 (.7)	-3.9 (.8)
Stayers				
U_{01}	10.6 (2.4)	6.3 (1.4)	4.7 (.4)	6.2 (1.3)
U_{10}	-6.6 (1.0)	-6.3 (1.0)	2.6 (.4)	1.6 (.3)
U_{11}	4.0 (1.7)	4.3 (1.9)	5.2 (2.4)	4.6 (2.1)
U_{00} (Base)				

When the movers' separation is distinguished by quit and layoff (see
Table A3, upper panel), the gain from joining a union is about the same
regardless of manner of separation, though for movers within the non-
union sector (U_{00}) the gain is a positive 5-6% for quitters and near zero
for those laid-off. These distinctions are not perceptible for the older
union joiners whose sample further divided into quit and layoff becomes
miniscule. However, old non-union movers (U_{00}) experience zero gains in
quitting and significant losses by layoff.

In turn, when movers are distinguished by moves between or within in-
dustry (Table A3, lower panel), (at the 2-digit level), gains to young union
joiners were observed only for inter-industry movers. They are not significant
for within-industry movers, but the latter comprise no more than 20% of
young movers who join unions. Again, no significant differences can be
observed in the small samples of older movers.

Movers who left unions $(U_{10}$ in Table 2) suffered losses which exceeded
the gains of movers who joined unions. The losses were even greater when the
separation from the union job was by layoff, and when the move was between
industries. Again less confidence should be attached to such findings for
the older men. If the loss of union leavers is to be viewed as another
measure of the union premium (with a negative sign), it is not clear why it is
so much larger than the positive measure. The discrepancy may reflect further
sorting by layoff from union firms, but the basis for such speculation is weak.

Table 3 presents estimates of annual wage change equations, pooled over the
period 1968 to 1978 in the MID. Only one definition of union status (union
membership) is available for all the years. The findings, coefficients of union
dummies, are shown for all ages and for the two age groups, below age 30, and
30-64, separately.

Table 3

Annual Wage Growth 1968 - 1978, MID, pooled

Movers	All	Age < 30	Age $\geqq$ 30
U_{01}	9.0 (3.2)	12.9 (3.4)	4.0 (1.0)
U_{10}	-1.8 (.6)	-.3 (.2)	-1.6 (.3)
U_{11}	6.4 (2.4)	5.3 (1.4)	8.2 (2.4)
U_{00}	4.3 (3.2)	7.0 (3.6)	.7 (.4)
Stayers			
U_{01}	1.7 (1.0)	4.5 (1.7)	-.4 (.2)
U_{10}	-2.7 (1.5)	-3.9 (1.0)	-2.1 (1.0)
U_{11}	-.6 (.9)	-.8 (.5)	-.6 (1.0)
U_{00} (Base)			

Young men who join the union by changing firms gain about 13% in wages; the older men's gain of 4% is not statistically significant. The average gain for all is 9%. This reduces to 6% for the young and less than 5% for all, if the union premium is viewed as net of the return to mobility into a non-union job. However, when the moves of union joiners are classified by quit and by layoff, (Table A4, upper panel), the gain by quitting was significantly larger: 19.8% for the younger and 11.8% for the older group. The non-union quitters had gains of 12.4% and 4.6% respectively. A net premium which would take into account the returns to mobility makes the net differential about 7% for quitters in both age groups. Again, as in the NLS, the gain for union joiners who are movers shows up in inter-industry mobility.

Compared to the NLS, MID estimates of the average union premium appear to be somewhat smaller.[1] A major difference is in the estimate of losses of union leavers: they are larger than the gains of joiners in the NLS, but insignificant in MID. However, the insignificance in the MID sample applies only to the young group. Among those over 30, union leavers who quit gain 12% while those laid off lose 8.6%. No significant wage changes are observed for persons who did not change jobs. Reported changes of union status in this group are more questionable than in the NLS sample.

To summarize the findings in both the NLS and MID: (1) Estimates of wage changes of union joiners are near 15% for young (< 30) white men and 4-12% for older men. If gains from (non-union) mobility are subtracted from these estimates, the net union premium is reduced to between 6-14% for the young joiners.

(2) The union premium appears to be larger for the young than for the older men. This is not true if collective bargaining rather than union membership is a criterion (in the NLS). Nor is it true in MID if union joiners quit from the preceding job.

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[1]Recall (p.6, above) that MID estimates of wage changes tend to be understated.

(3) The union premium is clear and significant mainly if union joiners quit the preceding job and moved between (2-digit) industries, as 80% of them did.

(4) Union leavers lose more than union joiners gain in the NLS. In the MID sample, only union leavers over 30 who were laid off lose as much or more than joiners gain. There is no evidence of losses for the younger union leavers.

2. Cross-Section Differentials and Selectivity in Hiring

Tables 4 (for NLS) and 5 (for MID) present coefficients on union status dummies in two cross-section wage regressions, year 1, prior to the change in union status, and year 2 after the change. Thus for job movers the coefficient on U_{01} in the "prospective" regression (year 1) estimates the wage of prospective movers from non-union to union jobs (e.g. a coefficient of 1.2 means 1.2% points larger than the wage of the base group = nonunion stayers). The coefficient of U_{01} in the "retrospective" regression (year 2) estimates the wage on the new union job of these recent union joiners, again relative to the base group in year 2.

If the existence of a union premium induces employers to select more productive labor, we would expect prospective union joiners to have higher wages in the prior, nonunion job than other nonunion workers. We have to be careful, however, to compare wages of new hires into union jobs with wages of new hires into non-union jobs: employers select among new hires. Thus the difference between the coefficients of U_{01} and U_{00} of job movers in the prospective regression (year 1) measures the upgrading in hiring into union jobs. Although union joiners had a prior non-union wage about the same as the base group of non-union stayers, their wages were significantly higher (about 10%) than the prior non-union wage of new hires into non-union jobs. This selectivity differential appears to be similar for the young and old NLS samples and is roughly comparable in size (a bit smaller) to the net union premium estimated from the wage changes.

Table 4

Wage level equations NLS

<u>1969 = year 1</u>

Movers	Young (UM)	(CB)	Old (UM)	(CB)
U_{01}	1.2 (.3)	-2.6 (.6)	-3.4 (.2)	-4.5 (.7)
U_{10}	20.8 (4.6)	20.7 (4.9)	3.1 (.2)	-2.9 (1.0)
U_{11}	33.3 (7.4)	27.2 (6.6)	42.9 (6.1)	43.7 (6.4)
U_{00}	-9.7 (4.0)	-10.2 (4.0)	-11.6 (2.6)	-11.0 (2.3)
Stayers				
U_{01}	6.1 (1.3)	7.2 (1.5)	-9.8 (1.8)	-9.4 (1.8)
U_{10}	11.7 (1.6)	6.9 (1.0)	-6.4 (.8)	-6.3 (1.0)
U_{11}	16.0 (6.2)	15.2 (6.0)	.2 (.1)	.3 (.1)
U_{00} (Base)				

<u>1971 = year 2</u>

Movers	Young (UM)	(CB)	Old (UM)	(CB)
U_{01}	22.0 (4.0)	14.4 (2.6)	7.6 (.5)	9.5 (.8)
U_{10}	1.8 (.3)	-4.5 (.1)	.2 (0)	-27.0 (2.0)
U_{11}	43.4 (7.9)	38.7 (7.6)	55.5 (7.6)	56.2 (7.1)
U_{00}	-3.2 (1.6)	-.5 (.2)	-9.2 (1.6)	-8.6 (1.5)
Stayers				
U_{01}	-1.5 (.2)	14.1 (2.6)	-1.5 (.3)	-3.5 (.6)
U_{10}	-2.8 (.3)	3.7 (.5)	-2.8 (.3)	-4.0 (.6)
U_{11}	20.0 (7.2)	19.6 (7.0)	4.5 (1.6)	3.4 (1.2)
U_{00} (Base)				

Table 5

Wage level estimates, MID - Pooled

Year 1	All	Age $\leq$ 30	Age $\geq$ 30
Movers			
U_{01}	-12.9 (3.9)	-6.2 (1.8)	-24.2 (3.8)
U_{10}	-8.6 (2.3)	2.2 (.5)	-17.2 (3.0)
U_{11}	17.6 (5.6)	21.5 (4.4)	13.7 (3.2)
U_{00}	-16.0 (11.3)	-14.1 (7.5)	-15.2 (7.2)
Stayers			
U_{01}	5.5 (2.3)	10.4 (2.6)	2.7 (.9)
U_{10}	3.5 (1.6)	4.4 (1.2)	3.1 (1.1)
U_{11}	17.8 (19.9)	22.8 (13.8)	15.4 (14.5)
U_{00} (Base)			
Year 2			
Movers			
U_{01}	-1.4 (.4)	13.2 (2.8)	-10.5 (1.8)
U_{10}	-4.7 (1.1)	1.0 (.2)	-2.6 (.4)
U_{11}	26.1 (7.4)	32.7 (6.2)	23.7 (5.7)
U_{00}	-14.2 (8.9)	-4.9 (2.0)	-15.2 (6.5)
Stayers			
U_{01}	5.3 (2.1)	10.7 (2.3)	3.7 (1.2)
U_{10}	1.9 (.7)	3.5 (.8)	1.8 (.5)
U_{11}	16.7 (17.3)	22.4 (11.2)	14.8 (13.1)
U_{00} (Base)			

The comparison of coefficients on U_{01} and U_{00} of MID movers (Table 5) in year 1 yields the conclusion that workers newly hired into union jobs had, on average, higher wages on the preceding job in the non-union sector than non-union workers hired into non-union jobs. The difference is smaller in the Michigan Panel than it was in the NLS. But this is a result of two opposite differentials by age. The young workers hired into union jobs had 8% higher wages than the prospective non-union hires, a figure comparable to the selectivity differential estimated for young NLS workers. However, the older group of new hires (30+) had lower wages than the comparison group of non-union hires.

Inspection of year 2 wages of new hires in their new jobs shows that the wage differential between non-union workers (in year 1) who moved to union and non-union jobs respectively (in year 2) has just about doubled (roughly from 10 to 20%) between year 1 and year 2 in both NLS and MID samples of young workers. Thus the union-nonunion differentials (among new hires in year 2) reflect selectivity in hiring and a net union wage premium in about equal measure. For the older worker the results are mixed: In the NLS the union-nonunion differential among new hires (year 2) more than doubled compared to year 1, while in the MID the differential changed from negative in year 1 to positive in year 2.

A sharper view of these comparisons is shown in Table 6. A little arithmetic is helpful in inspecting the table. First, it is clear that the <u>increment</u> in the wage differential between union and non-union new hires from year 1 (on the old jobs, both non-union) to year 2 (on the new jobs, one union, the other non-union) is in principle,[1] equivalent to the wage change regressions estimate of the union wage premium, net of selectivity, since:

(1) $\left[(\beta_{01} - \beta_{00})_2\right] - \left[(\beta_{01} - \beta_{00})_1\right] = \left[(\beta_{01})_2 - (\beta_{01})_1\right] - \left[(\beta_{00})_2 - (\beta_{00})_1\right]$

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[1] In practice, the estimates differ largely because of a differing structure of errors in levels and in changes.

Table 6

	Wage Differentials New hires $(\hat{\beta}_{01}-\hat{\beta}_{00})_2$	Selectivity in hiring $(\hat{\beta}_{01}-\hat{\beta}_{00})_1$	Implicit Wage Gain (1) - (2)	Estimated Net Gain of joiners (Tables 2&3)	Wage Differential of Stayers
	(1)	(2)	(3)	(4)	(5)
NLS					
Young (UM)	25.2	10.9	14.3	14.1	20.0
(CB)	14.9	7.6	7.3	7.1	19.6
Old (UM)	16.8	8.2	8.6	10.6	4.5
(CB)	18.1	6.5	11.6	15.5	3.4
MID (UM)					
Young	18.1	7.9	10.2	5.9	22.4
Older	4.7	-9.0	13.7	3.3	14.8
All	12.8	3.1	9.7	4.7	16.7

The β_{ij} are the regression coefficients on the respective union dummies U_{ij}.
The first term on the left hand side is the cross-sectional union-nonunion wage
differential for new hires, the second term measures upgrading in union hiring,
and the right hand term is the implicit union wage premium $(\Delta \beta_{01} - \Delta \beta_{00})_1$,
with returns on costs of mobility netted out.[1]

The interpretation expressed on the left hand side of (1) justifies the use
of the second term on the right hand side: Emphasis on new hires coming from the non-
union sector; where wages may be assumed to reflect the marginal value product,
justifies the netting out of returns to mobility of workers moving within the
non-union sector only.

Column (1) of Table 6 shows the cross-section union non-union differential
for new hires. Column (2) shows the selectivity (upgrading) differential, which
is a component of the union-nonunion wage gap (1). Column (3) is the difference
between (1) and (2), the implicit net union wage premium obtained by new hires.
For comparison, Column (4) shows the net union wage premium estimated from the
wage-change regressions, and Column (5), the cross-section union-nonunion wage
differential among stayers (coefficients of U_{11} of stayers in year 2), which
comes closest to the usual regression estimates of the union-nonunion differential
in the cross-section.

Judging by the first two columns of Table 6, selectivity in hiring, as measured
by prior wages, accounts for almost one half of the union-nonunion wage differential
among young new hires in the NLS, but less in the MID. The time-series estimates of
the net wage premium (col. 4) are comparable to the implicit cross-section estimates
(col. 3) in NLS, but they are smaller in the MID. For the older group in MID,
selectivity into the union (col. 2) appears to be perverse, that is negative.

The analysis summarized in Table 6 contains several innovations: (1) It relies on estimates of wage changes of workers whose union status changes as they move between firms. This is because reports of changes in union status while staying in the firm appear to be less reliable and show small or no effects. Previous studies which do not contain information on job mobility, therefore, probably attenuate the effects of unions on wages. (2) Information on mobility also permits more direct estimates of selectivity in hiring, as shown in col. 2 of Table 6. (3) Finally, the information on tenure makes possible estimates of gains from mobility as upward shifts of the whole tenure profile of wages from the old to the new job. Usually observed instantaneous wage changes may be negative, as is often found, but the longer run effects obtained in the present analysis are more likely to be positive.[1]

As to numerical results, it appears that estimates of net union gains in wages, as seen in new hires, range from 7 to 15% in the NLS and from 3 to 14% in the MID. Since the latter is biased downward, upward adjustments would put the central tendency of both NLS and MID a little over 10%. Selectivity accounts for almost as much in the NLS and the young workers in MID, but not so in the over 30 group in MID.

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[1] Even though superior as measures of gains from mobility, the shifts in tenure wage profiles are not fully commensurate with gains in present values. It can be shown that in order to secure the same gain in the present value of wages, moves after longer tenure require a larger shift in the tenure profile on the new job. This was confirmed empirically by positive and significant coefficients on a job change x prior tenure interaction variable in the wage change estimates (not included in the present study).

3. Turnover in Union Employment

If the wage received in union employment exceeds the worker's opportunity wage in non-union employment, he is less likely to quit a union job than a non-union job.

As Table 7 indicates, quit rates in the union sector are about half as large as in the non-union sector for young workers and are about one third as large for men over 30 (in MID), and one seventh for men over 48 (in NLS). The differences are smaller for separations, since layoffs are somewhat larger in the union, at least in the NLS.

Although one third of the white male workers are unionized (somewhat less among the young), less than 10% of non-unionized job seekers find employment in the union sector, while a half of the young and two-thirds of the older unionized job movers find employment in the union sector. Since these statistics apply to "unstandardized" workers, they may reflect differences in tastes and in geography, industry, and occupation. At face value, at any rate, the differences are consistent with our findings: non-union workers could get a bigger wage gain by moving to union jobs, but are evidently prevented by fewer vacancies and non-probabilistic rationing (stricter hiring standards, nepotism, etc.) resulting from the wage premium in the union sector. Union workers, however, can gain from mobility within the union sector, but face a wage loss when they leave the union sector. (Tables 2 and 3 above). Thus they tend not to leave the firm unless there is a good chance of landing another union job.[1]

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[1] If the existence of grievance procedures ("voice" instead of "exit") were the major force that ties groups of workers with similar preferences to a firm, we would expect lesser quit from firms with such labor relations. But, in the absence of a wage premium in unionized firms, those who separate from such firms would be largely indifferent whether they moved to a union or non-union firm. It is, incidentally, not clear why grievance procedures cannot exist need be or less effective in non-unionized firms, if reduction in turnover reduces labor costs.

Table 7

Turnover Rates (%), by Union Status and Age

	NLS			MID	
(a) Quits	Young	Old		Age < 30	Age ≥ 30
NU	26.7	6.5		14.6	4.7
U	14.3	.9		7.5	1.7
(b) Layoffs					
NU	13.5	5.0		7.0	3.0
U	17.7	7.0		6.6	2.7
(c) Separation					
NU	40.2	11.5		21.6	7.7
U	32.0	7.9		14.1	4.4

If unionization reduces job mobility, this reduction should be observed on the same individual by comparing his mobility before and after joining a union firm. And if the wage premium gained by moving to a union firm matters, the reduction in mobility should be greater the greater the wage gain.

According to Table 8, young NLS men who joined unions between 1969 and 1971 have prior quit frequency (in 1967-1969) which was not smaller than the quit of other non-unionized workers at that time.[1] This is based on regressions with quit as a dependent variable (0,1) and 1969 values of the same standardizing variables as used in the wage level regressions. However, the frequency of quit was lower by about 13% points in 1971-1973 after joining a union (coefficient of U_{01}) in the interval 1969-1971. This reduction is as large as the unstandardized cross-section difference in Table 7. In 1971-1973 union stayers (U_{11} in 1969-71) had about the same low quit rates as union joiners.[2]

No significant results were obtained for the older NLS men who joined unions, but union stayers had significantly lower quit rates than non-union workers both in 1967-1969 and in 1971-1973.

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[1] For a similar finding, based on other data, see Freeman (1980).

[2] Recall that union status was not reported in 1967, so many of the young union stayers in 1969-71 were likely to have been joiners in 1967-69.

Table 8

Quits and Layoffs, Before and After

Change in Union Status, NLS

	1967-69			1971-73	
Quits	Young	Old		Young	Old
U_{01}	-1.8 (.9)	+ .9 (.4)		-12.7 (2.8)	-1.1 (.5)
U_{10}	2.9 (.6)	- .8 (.2)		- 4.9 (.9)	-2.3 (.8)
U_{11}	-4.4 (1.5)	-2.2 (1.7)		-16.4 (5.2)	-2.7 (2.7)
Layoffs					
U_{01}	6.0 (1.2)	3.4 (1.1)		10.3 (2.7)	-1.0 (.4)
U_{10}	16.5 (3.2)	7.8 (1.8)		4.8 (1.1)	3.6 (.9)
U_{11}	-9.4 (3.3)	- .7 (.5)		4.1 (1.6)	2.6 (1.9)

A comparable analysis was performed on the full MID sample (including movers and stayers) using all ages and changes in quit (ΔQ) as the dependent variable which assumes values (-1, 0, +1). Again, the standardizing variables were the same as in the Δ ln wage equations. The results are shown in col. (1) of Table 9.

Union joiners experienced significant reductions in quit rates compared to all other groups. Larger wage gains (for all movers and stayers) also reduced quits in the next period. Here we do not distinguish wage gains of union joiners from those of everyone else.

To observe this distinction we restricted the sample to job movers (about two-thirds of the sample had at least one move during the 10-year panel period). We constructed a mobility index (M), which is a count of numbers of firms in which the person worked up to the current job (i.e. number of separations + 1) divided by the time interval over which mobility was recorded. The denominator of the index is the interval since entry into the labor force, or since 1958, if entry into the labor force was before this date. Wage changes associated with job moves were recorded separately for union joiners and for others between 1968 and 1973, and their effects estimated on the change in M (and in ln M) between the move and 1978, the last year of the panel.

The numerator of the mobility rate includes permanent layoffs in addition to quits, but is dominated by quits. We could not restrict the index to quits, since the pre-1968 mobility is reported as separations. However, as Table 8 showed, although (permanent) layoffs are somewhat larger in the union sector in the NLS, they are not larger in the MID sample. But, even in the NLS differences in separations are dominated by differences in quits.

Table 9

Changes in Quit Related to Changes in Union Status and in Wages
(MID)

Changes in Quits (Full Sample)		Changes in Mobility Rates (Sample restricted to movers prior to 1974)		
	ΔQ		$\Delta \ln M$	ΔM
U_{01}	-6.1 (3.5)	U_{01}	- .24 (4.4)	- .01 (.3)
U_{10}	.7 (.4)	$\Delta \ln W_{01}$	- .27 (2.0)	- .20 (2.4)
U_{11}	1.2 (1.3)	$\Delta \ln W_{00}$	- .05 (1.7)	- .07 (1.6)
$\Delta \ln W$	-4.9 (3.4)			

A preliminary cross-section analysis (Table A5, panel (a)) showed that the mobility rate (M and lnM) declines with experience in a decelerating fashion, is reduced by education, marital status, union status, and by economy wide unemployment.

The dependent variable (ΔM, or ΔlnM) in col. 2 and 3 of Table 9 is the difference between the index in 1978 and the first move observed between 1968 and 1973 (inclusive) for non-union people (for most of them this was not the first move, since they moved also before 1968) or the move at which they joined a union. The union dummy U_{01} distinguishes union joiners from non-union movers and wage gains of joiners (ΔlnW_{01}) are distinguished from wage gain of non-union movers. The standardizing variables are in the form of differences between 1978 and levels at the time of the move. Dummies were added to account for entry into the labor force before 1968, and 1958, and years observed in the panel.

The results show that union joiners reduce their mobility more the larger the wage gain from joining a union. A similar, but much smaller, and barely significant, effect is observed for wage gains of movers in the non-union sector. This is true whether the index or the changes in wages are logarithmic or arithmetical.

The weak effect of an episode of mobility among non-union jobs is not surprising: unless the rate of return on the cost of moving is unusually high, there are no disincentives to future job search. However, the larger the wage premium for union joiners the greater the potential loss from future mobility. This reasoning applies more clearly to quit behavior, but the effects on total separations, if weaker, are similar.

Whether mere unionization without a wage premium produces a reduction in mobility is not as clear in Table 9. The union dummy has a strong effect on the change in the logarithmic, but not arithmetical, mobility index. It is also strong in Col. (1) where the wage gain of union joiners was not separated from others. Undoubtedly it represents a response to other union benefits, in fringes and in working conditions, including the more certain advantages of seniority. Nevertheless, our findings support the inference that the union wage gain is not merely a compensatory differential for quality of labor, or costs of mobility, or for inferior conditions in union jobs. Union rents appear to be real, although relatively small (closer to 10% than to 20%).

Although cross-section union-nonunion wage differentials (gross) over-state the size of the (net) union premium, it may be useful to explore their effects on differentials in quit rates in the cross-section, on the assumption that the gross and net wage differentials are positively correlated. Again the purpose is to find out whether it is not merely union-status but the size of the differential that affects mobility.

The companion analysis whose reults are shown in Table 10 exploits the large size of the MID panel in another way: we explore whether union-nonunion wage differentials affect quit rates within , alternatively, industries, regions (states), and occupations. The analysis proceeds in two steps: first we added industry (or state, or occupation),union,and industry x union membership dummies to the cross-section wage function. The coefficients of the cross-product dummies (β_i) are estimates of within-group union-nonunion (relative) wage differentials (estimates shown in Table A6). In the second step we included industry (or other groups), and a union dummy as well as its cross-product with the estimated union-nonunion wage differentials (β_i from the wage regressions) in quit regressions.

Table 10

Quit Differentials within Groups

MID, 1968-1978

Group*	U	U x β
Industry	-.032 (3.5)	.002 (.1)
State	-.024 (3.0)	-.080 (1.8)
Occupation	-.002 (.2)	-.21 (4.1)

* Listed in Table A6.

We ruled out groups
for which we had too few (less than 20) observations in the union or nonunion
sector. Thus we utilized 22 industries, 31 states, and 27 occupational
categories (Table A6).

The intra-group union-nonunion wage differentials β_i amounted, on average,
to 15-20%. They ranged from 0 to over 40% and were most prominent (larger and
more significant) in the occupational category, less so within states, and
least within industries.

Table 10 which reports the results of step 2, shows respectively the effects
of union status (U), and of wage differentials (β)
on differences in quit rates between unionized and non-unionized workers within
the various groups of workers. Effects of wage differentials are not significant
within industries (at a level higher than 1-digit. for workers with the same
measured characteristics , but they are significant within regions and occupations.
Union status alone (the U dummy) is sufficient to affect the intra-industry differences,
but plays no role other than via wage differentials inside occupations. Both
variables are significant within states. Note also that U is significant in all three
groups, when β is dropped, and conversely.

4. Fringe Benefits and Hours of Work

If numbers employed (N) and hours per worker (H) are viewed as separate factors of production, cost minimizing employers will determine their demand for N and H at the point where the ratio of marginal factor costs is equal to the ratio of marginal productivities, that is to the solpe of the optimal production isoquant.

Following Rosen (1968), the equilibrium marginal factor cost ratio is.:[1]

$$(1) \quad \frac{MC_N}{MC_H} = \frac{H}{N} + \frac{F(r + q)}{NW}$$

where F is the fixed cost of employment per worker, amortized per period by r, the interest cost of capital and by q, the worker quit probability which depreciates the capital sum creating a capital loss Fq per period.

To the extent that F is positive an increase in the wage rate W reduces the factor cost ratio shifting the relative demand away from hours toward numbers.[2] In the minimum wage case it may be argued that F is significant at most in terms of training expenses, but that minimum wages tend to reduce or eliminate such expenses, so that the predicted effects on hours may be observed only in the short run before the adjustment is completed,or it may be indeterminate.

---

[1]Let total labor costs be $C = NHW + NF (r+q)$, Then $MC_n = \frac{dC}{dN} = HW + F (r+q)$

and $MC_H = \frac{dC}{dH} = NW$

[2]This analysis of substitution between H and N ignores scale effects. The qualitative conclusions, however, remain valid so long as the elasticity of demand for numbers (N) with respect to wages (including fringes) is less than unitary.

In contrast to the minimum wage, union pressure on wages extends to most components of the wage package.

Indeed, union push on components other than directly paid out wages appear to be even stronger. Union fringe benefits exceed non-union benefits not only in dollar value but also as a proportion of the wage package (about 30%). A number of possible explanations have been conjectured, running from union democrasy which favors the older worker to union management of pension funds as an instrument of power. One economic argument relies on reduced turnover, which is a result of union wage push and of other gains. In the presence of incomplete vesting of pensions in the worker, longer tenure of union members means that the probability of ultimately receiving the pension is higher in union than in non-union jobs. Hence the incentive to push for larger pensions[1] (Freeman, 1978). But why increase fringe benefits by a larger percentage than the increase in the paid out wage? One reason is the higher marginal income tax rate, if the income elasticity of worker demand for fringes is otherwise unitary (Rice, 1966). But this may explain only a small part of the proportional increase (Donsimoni and Shakotko, 1979).[2]

The analysis of effects of wage push on hours may provide one rationale for union pressure on fringes: an increase in union wages W, with F unchanged, would lower the ratio of marginal factor costs both by raising the denominator in the second right hand component of equation (1) and by reducing q in the numerator. If hours are reduced, weekly earnings may not increase much even if wage rates rise significantly. To blunt the adverse effect on hours, more specifically, to prevent their reduction which would limit union gains in earnings, quasi-fixed costs, such as fringes, must be increased by a larger percentage than the paid out wage (W), since quits (q) decline when both W and F are increased.

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[1] For the same reasons, more vesting is less costly to union than non-union employers.

[2] They found that fringe benefits were substantially larger for union workers whose total compensation (wages + fringes) was the same as the total compensation of non-union workers.

In contrast to hypotheses which rely on non-wage aspects of unionism to explain the larger ratio of fringe benefits to paid out wages in union employ- ment, this analysis predicts a positive link between the percent union wage premium and the relative increase in fringe benefits (F). Moreover, the per- cent increase in F is expected to exceed the percent increase in the wage, since the larger the latter, the bigger is the decrease in the quit rate. Thus unions which achieve the biggest gain in paid out wages would also want the largest <u>proportion</u> of their total compensation in fringe benefits.

This proposition is tested on a sample of over 4,000 firms in 70 (2-digit) industries. Average paid out wages in union firms were 21% higher than in non- union firms within the industry (a simple average of 70 wage ratios), while average fringe benefits were 60% higher. The first regression (Table 11, Panel A) relates the ratios of fringes $\left(\frac{Fu}{Fn}\right)_i$, where u is union, n is nonunion, and i is industry to ratios of money wages, standardizing for age of employees, size of firm, geographic region, and union coverage of the industry.

As predicted, the coefficient on the wage ratio in Table 11 is significantly larger than unity (1.62). This implies that a 10% increase in the union wage premium would create a 16% increase in the union-nonunion ratio of fringe benefits. This ratio was, on average, 32% in the unstandardized data.

A second test was performed, at the firm level on union firms alone. First a wage and fringe function was estimated on non-union firms in order to impute non-union levels of wages and fringes to union firms. Thus the denominators of the wage ratio and the fringe ratio are imputed values which workers in union firms would receive if they were not unionized. This is a more stringent test because it relates fringes to wages in union firms only, and because errors in imputation bias the coefficients against the hypotheses (downward). The results (Table 11, Panel B) nevertheless, are once again as predicted, although the co- efficient on the wage-ratio is, indeed, smaller than in Panel A.

Table 11

Fringe Benefits and Wages, Regressions on Industry (A) and Firm (B) Levels

(A) Industry Level—Panel A				(B) Firm Level—Panel B		
Variables		T-Stat		Variables		T-Stat
F-BRAT	--Dep. Variable --			HFBRAT_U	--Dep. Variable --	
MWRAT	1.6183	6.0		HTWRAT_U	1.5167	41.3
TWRAT	--	--		HMWRAT_U	--	--
SIZE_N	-0.0422	-0.8		COVER	-0.4483	-6.6
SIZE_U	0.0575	1.1		C4	0.2748	3.2
SOUTH_N	0.1750	0.9		EDUCATION	-0.0503	-2.0
SOUTH_U	0.0315	0.1		MALE	0.1217	1.7
URBAN_N	-0.0638	-0.3		URBAN	0.0485	2.2
URBAN_U	-0.0212	-0.1		SOUTH	-0.0642	-2.4
Y1_N	0.9546	3.6		SIZE	-0.0172	-2.2
Y1_U	-0.6160	-2.2		YOUNG	0.0001	0.1
Y2_N	0.0161	0.1		OLD	-0.0018	-0.7
OLD	0.0052	0.9		Y1	0.0150	0.6
COVER	0.3413	1.8		INTERCEPT	-0.4607	-1.4
INTERCEPT	-0.9640	-1.7				
DFE		56		DFE		2551
F-RATIO		6.7		F-RATIO		146.6
R-SQUARE		.61		R-SQUARE		0.46

Table 11

Fringe Benefit Regressions

Variables and Sources

Dependent Variable: F-BRAT, union/non-union fringe benefit ratio.

Independent Variables:

INDUSTRY-LEVEL REGRESSIONS: Panel (A)

MWRAT   --union/nonunion money wage ratio

SIZE   --size of establishments measured by the logarithm of the total
          number of employees, industry average. Source: a.

SOUTH   --proportion of establishments in the south. Source: a.

URBAN   --proportion of establishments in urban areas. Source: a.

Y1,Y2   --proportion of firms surveyed in 1967-1968 wave and 1969-1970
          wave respectively. Reference group was 1971-1972 wave. Source: a.

OLD   --industry average of the proportion of workers over age 50.
          Source: a.

COVER   --proportion of workers covered by collective bargaining agreements.
          Note: the industry definition for this variable is slightly
                different from the industry definition used in the main
                part of the study, but no industries are more broadly
                defined in the COVER variable. Source: c.

FIRM-LEVEL REGRESSIONS: Panel (B)

COVER   --see above.

C4   --Four-firm sellers' concentration. Weighted sum of 4-digit level
          concentration figures with value of shipment used as weight, and
          raw data corrected to account for regional and national markets.
          Source: b.

EDUCATION --average years of education of blue-collar workers. Source: b.

MALE   --proportion of male blue-collar workers. Source: b.

URBAN   --dummy, 1 if firm is in an urban area. Source: a.

SOUTH   --dummy, 1 if firm is in the south. Source: a.

SIZE   --see above, log of number of employees in firm. Source: a.

Firm-level regressions cont.

    YOUNG   --percentage of firm employees under age 35. Source: b.

    OLD   --percentage of firm employees over age 40. Source: b.

    Y1,Y2   --see above, dummy form. Source: a.

Source   a   Expenditures for Employee Compensation. Three surveys conducted by the Bureau of Labor Statistics in 1968, 1970, and 1972 have been pooled. They include 4,073 establishments, 2,580 unionized and 1,493 nonunionized.

              b   May Current Population Survey. Three surveys conduced 1973, 1974, and 1975 have been pooled. They include 50,000 households with 49 percent of union members.

              c   Richard Freeman and James Medoff. "New Estimates of the Industrial Unionism in the U.S." Industrial and Labor Relations Review, January 1979, 143-174.

In the data analyzed in Table 11 the value of union fringes is 80% higher than of nonunion fringes, while wages are 30% higher. Assuming that the net union effect is a half of this figure (i.e. 15%), and that quit rates are half as large in the nonunion sector (as observed in Table 7), we can calculate the required <u>ratio of fixed costs</u> $\frac{F_u}{F_n}$ which would remove employer incentives to cut hours schedules of union workers. Given N (numbers employed), the condition is, by eg. (1):

$$\frac{F_u(r+q_u)}{W_u} = \frac{F_n(r+q_n)}{W_n}$$

Assuming $r = 10\%$, and $q_n = 10\%$, $\frac{F_u}{F_n} = \frac{W_u(r+q_n)}{W_n(r+q_u)} = 1.15 \times \frac{4}{3} = 1.53$.

The actual <u>ratio of fringe benefits</u> is larger (1.80), presumably because fringes represent only a part, although a major one, of fixed labor costs.

An important consequence of higher fixed costs (in hiring and in fringes) imposed on union employers is greater stability of employment--reduced fluctuations in N, when labor demand fluctuates (see Rosen, 1968). As a result, the major means of adjustment to fulctuating demand in union employment are the use of overtime when labor is short and the use of temporary layoffs (recall unemployment) in slack times. Temporary layoff is favored by union workers, as it implies lesser income loss than corresponding reductions in weekly hours, because of unemployment compensation and other unemployment benefits (see b). For their part, employers can expect less attrition, since temporarily laid off union workers are less likely to look for other jobs than comparable non-union workers.

The evidence that average weekly hours are not less in the union than non-union sector, that overtime is more prominent and that temporary layoffs are more frequent and a larger porportion of total layoffs in the union sector is available[1], and con- sistent with our hypothesis. In this light union pressure on fringe benefits is not merely (or at all) a tradeoff for higher wages, but a policy which increases both earnings and job security of union workers.

[1]In MID, straight-time in weekly hours are about 4% shorter in union jobs, but total hours are no less in union than in nonunion jobs. See also Blau and Kahn (1981), and Raisian (1981).

5. Union Wage Profiles and Job Training

The cross-section union-nonunion differentials in wages diminish with age. They are reduced from over 22% to about 15% between ages less than and more than 30 for union stayers (U_{11}) in MID (Table 5), and reduced further to a little over zero for those over 48 in NLS (Table 4). The implication is that the typical union age (experience) wage profile, although higher in level is flatter than the typical non-union profile. This difference has been found in other studies and has been ascribed to union policy of compressing wage differentials across firms and workers, who may differ in sex, education, race, and age.

The policy of wage compression has been attributed to union pursuit of equity, to administrative convenience in collective bargaining, and to union efforts to reduce competition from lower wage firms. However, a more direct explanation of flatter wage profiles involves union emphasis on wage progression by seniority rules: Within the firm, the wage structure and other rules of collective bargaining agreements are specified in terms of seniority, or job tenure. Although wages grow as tenure lengthens in non-union jobs as well as in union jobs (Table 12), the explicit seniority rules in union firms are a great deal more rigid: Seniority is the necessary condition for promotion in most union firms. If the higher job level requires additional training, union clauses often provide that senior employees are to be trained in order to fill the higher level vacancies.[1] Such provisions limit the supply of trained workers from the outside. At the same time they severely reduce the benefits from transferable training. Consequently, incentives for general (transferable) training are reduced for union workers, both because such training is not adequately rewarded within the union firm, and because union workers are less likely to move in the first place.

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[1] See BLS Bulletin, "Union Status and Benefits of Retirees", July 1973.

Table 12

Union and Non-union Wage Functions

MID, Pooled, 1968-1978

(A) Wage Level (lnW)

	Union	Nonunion
Ed	5.0 (22.4)	9.6 (46.4)
X	1.7 (10.6)	3.7 (21.2)
X^2	-0.2 (7.4)	-0.6 (17.9)
T	2.4 (12.5)	2.1 (9.3) 1.6 (6.5)
T^2	-.06(8.1)	- .03(3.2) -.02(2.2)
Mar	.4 (.2)	13.1 (7.1)
Hlth	-7.3 (4.4)	-12.7 (6.1)
LocUn	1.2 (6.1)	2.7 (11.1)
M	-.8 (.9)	-5.5 (5.1)

(B) Wage Growth (ΔlnW)

	Union	Nonunion
	.1 (.5)	-.2 (1.1)
	2.0 (.5)	6.5 (1.9)
	-.04(1.7)	-.06(2.3)
	1.0 (1.5)	1.0 (2.2)
	-.03(.9)	-.02(.3)
	-.3 (.2)	-.4 (.3)
	-4.2 (2.5)	2.6 (1.3)
	.1 (.7)	.1 (.7)
	.2 (.3)	-1.3 (1.6)

This helps to explain the declining with age proportion of job movers who join union firms (Table 1), and the smaller coefficient of experience[1] at fixed levels, in the wage function of union compared to non-union workers (Table 12).

Whatever training exists in union firms, and certainly some initial (apprentice) training exists in crafts, and is provided for the purposes of filling more skilled vacancies from within, almost by definition, all of it is specific to the firm. If wage growth with tenure at fixed levels of experience reflects the growth of specific capital, tenure-wage profiles in union firms need not be flatter and may even be steeper than in non-union firms.

This is, indeed, found in Table 12, where the coefficient on tenure in the union wage equation is no smaller than in the non-union equation. Table 12 is restricted to MID data because estimates of experience and tenure coefficients in NLS are sensitive to truncation by age. The non-union tenure coefficient in the wage level equation (left-hand panel) is smaller in col. (3), where an attempt was made to adjust the estimate for heterogenity bias[2] by introducing the (prior) mobility rate variable. This variable (M) has a negative and significant effect in the non-union equation, where it also reduces the tenure coefficient, but has no effect in the union equation. The right-hand panel of the Table shows separate wage growth equations for union and nonunion workers. The comparative results are similar to the results in wage levels, but all coefficients are smaller and less significant.

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[1] And of education, if job training is complementary.

[2] For an introduction of this approach, see Mincer and Jovanovic (1978).

The experience coefficient is much smaller in the union equation. Thus the flatness of gross union wage profiles (by age, or experience) when tenure is ignored, or by tenure when experience (or age) is ignored is due solely to lesser worker investments in general training.

The flatter gross union profile also suggests that the total volume of training is smaller in union firms. In principle, this need not be the case, since the wage profile measures returns to workers, but not returns on costs borne by employers. The reluctance of union workers to quit should provide incentives for union employers to invest in specific training of their workers, since the risk of a capital loss by worker quit is smaller. Union workers have a corresponding insurance against capital losses only to the extent that seniority rules reduce layoffs.This is true at higher levels of seniority at the expense of low-tenured workers. Thus, in contrast to the usual (competitive) analysis (Becker, 1975) where worker turnover - both quits and layoffs - are a result of investments in specific human capital shared by workers and employers, such investments are the result, rather than cause, of turnover patterns induced by union pressure.

Because of lower quit, we might expect employer investments in specific training of union workers to exceed the corresponding investments of non-union employers. But the fact that _permanent_ layoff rates of union workers are no smaller than in the non-union sector raises doubts about such expectations. However, higher rates of union layoffs concentrated at low tenure levels may both reflect additional screening of new hires and may be a substitute form of adjustment for reduced quit.[1] If the volume of specific training is not clearly larger in union firms, total training,including the general component, is likely to be smaller. This is apparently confirmed by the more direct evidence on reported training, to which we proceed:

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[1] Blau and Kahn find that unionism increases permanent layoffs among the young, not among the old NLS. Our own research (not presented here) shows that while quits and layoffs decline with tenure at about the same rate for non-union workers, layoffs decline more steeply than quits for union workers.

In the NLS, training was reported in response to questions: Do you receive or use additional training (other than school training) on your job? In MID, different questions were used in different surveys; the most appropriate for our purposes is the question asked in 1976-1978: "Are you learning skills on the current job which could lead to a better job or to promotion?"

In a recent paper (Mincer and Leighton, 1981) these questions were used to explore the effects of minimum wage laws on training on the job. Training so reported was used as a dependent variable[1] in the NLS and MID equations (the same set of independent variables was used as in the wage equations). A union dummy included in those equations was significantly negative in all periods in the MID and in the older NLS sample.

In the current study, we further classified the union status variable as joiners (U_{01}), leavers (U_{10}), and stayers (U_{11}), and explored the incidence of training before and after the change of union status.

In the NLS samples (Table 13, (A)) the old union stayers and union joiners show significantly less training than non-union stayers. The signs on coefficients of the young union joiners are negative, not significant before joining and almost significant after. Old union leavers show positive, but not significant coefficients.

The MID results (Table 13(B)) similarly show significantly less training among union stayers (U_{11}). They also show that less training is required on union jobs, though not as a precondition for hiring (coefficient U_{01} is not significant in year 1, Panel C). The coefficients of union joiners are negative and not significant before but significant after joining the union (year 1 and year 2 regressions). Inferentially, they received more training than union stayers before the start of the union job. Conversely, union leavers appear to receive more training than before on the non-union job to which they moved.

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[1]As an independent variable, the training dummies are positve and significant in both wage level and wage growth equations (not reported here).

Table 13

(A) Union Effects on Training, NLS 1969-1971

Training in 1969 Job				Training in 1971 Job		
	Young	Old			Young	Old
U_{01}	-3.1 (.7)	-6.5 (1.8)			-5.5 (1.3)	-4.9 (1.4)
U_{10}	-1.3 (.3)	4.6 (1.1)			- .3 (.1)	2.7 (.6)
U_{11}	- .4 (.1)	-5.1 (2.9)			-3.6 (1.3)	-7.4 (4.1)

(B) Union Effects on Training, MID 1976-1978

(Ages $\leq$ 45)

Year 1 (1976)		Year 2 (1978)	
U_{01}	-6.3 (1.2)	-11.5 (3.0)	
U_{10}	-14.4 (2.4)	-3.8 (.9)	
U_{11}	-15.0 (5.3)	-12.9 (5.8)	

(C) Is Non-School Training Required for Current Job?

(MID, Education $\leq$ 12)

Year 1 (1976)		Year 2 (1878)	
U_{01}	-3.9 (.6)	-6.8 (1.7)	
U_{10}	-7.5 (1.1)	3.6 (.8)	
U_{11}	-6.7 (2.1)	-9.7 (4.2)	

The findings on workers who <u>changed</u> union status are not very secure, but
they suggest that selectivity in hiring may involve prior training (in addition
to some concentration of it at the outset of some of the union jobs, such as in
crafts). This may explain the results for the very young NLS workers where
union effects on training were not significant.

The finding that union workers receive less training on the job than non-
union workers is also confirmed in the Michigan Time Study, as reported by
Duncan and Stafford (1980). They report that while non-union workers spend,
on average 6.1 hours per week on job training, comparable union workers spend
4.2 weekly hours on such training.

Summary

Amoung a number of features that distinquish unionized from non-unionized
labor markets, at least four have been repeatedly observed in empirical studies:
(1) Higher wages, (2) Larger fringes, more than in proportion to wages, (3) Lesser
turnover, reflecting lesser quit, though not less layoffs, and (4) Flatter age-
wage profiles. Although the labor monopoly hypothesis has been used (and disputed)
as an explanation of higher union wages, it does not receive much, if any,
attention in analyses of the other features.

This study explores the existence of a net union premium and of the extent
of rationing by quality of the resulting excess supply. The net union premium
was estimated in Section 1 by relating changes in wages to changes in union status
of the same worker in longitudinal panels (NLS and MID). In Section 2, two cross-
section wage level regressions, a "prospective" and "retrospective" permit more
direct observation of selectivity in hiring. Over a half of the cross-section
differential of over 20% for the "same" (standardized) worker is a net union rent and
much of the rest reflects a quality adjustment in hiring, as measured by wages.
This conclusion was less reliable for older workers.

The next step (Section 3)was to ascertain whether the net union wage premium is responsible, together with other advantages secured by the union such as fringes and seniority rules, for the lesser turnover, especially quit of union workers. The answer is positive: the reduction in quit depends on the size of the net wage premium in an individual analysis, and is also positively related to the gross, cross-section wage differentials within groups of workers,classified by location and occupation, though not by industry. Fringes and seniority rules, or other union advantages did not explicitly enter this analysis but they are likely to be embodied in the union membership variable.

In Section 4, it is hypothesized that the imposition of larger fixed labor costs (such as fringes) helps to deter employers from preferring reductions in hours to reductions in men, and it helps to stabilize employment in the face of fluctuating demand,by a more frequent use of overtime and of temporary layoffs in the union sector. This hypothesis links the size of fringe benefits to the union wage gain. An analysis of firms in 70 industries confirms this link.

Section 5 explored the consequences of union pressure on the tenure profile of wages by the rather rigid linking of wage levels to seniority in the job. The consequently reduced incentives for worker investment in general (transferable) training, even if no such reduction needs to apply to specific training, is consistent with observed flatter experience (or age) profile of wages of union workers. In sum, total training is likely to be less frequent in union firms and this is confirmed by direct responses in survey reports.

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Table A1
NLS Wage Equations

(a) Wage Growth

$\Delta \ln W$ (1969 to 1971)

	Young Men				Old Men		
	Mean	β	t		Mean	β	t
$\Delta \ln W$	8.2				2.6		
Intercept		.10	(5.5)			.038	(.9)
ΔX^2	22.8	-.002	(2.7)		152.0	.0005	(1.3)
ΔT	1.19	.018	(2.4)		1.22	-.0004	(.6)
ΔT^2	8.11	-.001	(.8)		46.22	.0002	(1.0)
Mar	.63	-.007	(.4)		.92	-.017	(.7)
Hlth	.09	-.034	(1.2)		.21	-.0006	(.9)
Lun	3.7	-.0003	(.5)		3.8	-.0002	(.5)

$R^2 = .35$ (Young Men) $R^2 = .10$ (Old Men)

(b) Wage Level (69)

	Young Men				Old Men		
	Mean	β	t		Mean	β	t
$\ln W$	1.12				1.34		
Intercept		-.19	(2.4)			1.21	(2.5)
Educ	12.2	.078	(15.6)		10.3	.053	(11.5)
X	6.7	.045	(4.5)		39.0	.033	(1.4)
X^2	57.8	-.001	(1.3)		1551.4	-.0003	(1.0)
T	1.65	.024	(2.2)		14.8	.003	(1.1)
T^2	7.12	-.002	(1.5)		342.2	-.0001	(1.4)
Mar	.63	.096	(4.5)		.92	.238	(5.9)
Hlth	.09	-.024	(.8)		.21	.043	(1.5)
Lun	3.7	-.002	(.4)		3.8	.001	(1.5)

$R^2 = .36$ (Young Men) $R^2 = .28$ (Old Men)

Table A2

MID Wage Equations
1968-1978, annual, pooled

(a) Wage Growth

	All			Age $<$ 30			Age $\geqq$ 30		
	Mean	β	t	Mean	β	t	Mean	β	t
Δ lnW	2.25			4.91			1.06		
Educ	12.5	-.001	(1.1)	13.0	.001	(.4)	12.2	-.001	(1.1)
Δ X	1.0	.048	(2.3)	1.0	.021	(.4)	1.0	.042	(1.6)
Δ x^2	40.2	-.0005	(3.5)	13.1	-.002	(1.4)	52.2	-.0002	(1.2)
Δ T	.64	.010	(2.7)	.57	.021	(2.1)	.68	.007	(1.9)
Δ T^2	12.8	-.0002	(1.2)	3.6	-.001	(1.5)	16.8	-.0007	(.4)
Mar	.92	-.005	(.5)	.88	.017	(.9)	.94	-.02	(1.5)
Hlth	.08	.004	(.4)	.06	-.013	(.5)	.09	.009	(.8)
Lun	5.6	.001	(.8)	5.6	-.002	(1.0)	5.6	.0015	(1.2)
Nun	.3	-.004	(1.8)	.3	-.013	(2.8)	.3	-.0003	(.1)
	$R^2 = .007$			$R^2 = .012$			$R^2 = .003$		

(b) Wage Level (year 1)

	Mean	β	t	Mean	β	t	Mean	β	t
lnW	1.68			1.53			1.76		
Intercept		.522	(18.5)		.563	(9.3)		.541	(12.0)
Educ	12.5	.073	(51.0)	13.1	.062	(18.3)	12.2	.075	(44.2)
X	19.5	.028	(23.9)	6.1	.058	(7.5)	25.8	.025	(11.4)
X^2	542.6	-.0005	(19.7)	46.2	-.003	(5.4)	771.4	-.0004	(11.0)
T	7.4	.022	(14.6)	2.2	.056	(10.9)	9.8	.018	(10.9)
T^2	118.5	-.0004	(7.0)	11.1	-.003	(5.9)	168.0	-.0002	(4.2)
Mar	.91	.083	(6.2)	.86	.065	(3.4)	.94	.086	(4.5)
Hlth	.08	-.108	(7.9)	.06	-.076	(2.8)	.09	-.119	(7.5)
Lun	5.6	.021	(13.2)	5.6	.023	(8.5)	5.6	.020	(10.2)
Nun	4.8	-.022	(4.6)	4.8	-.024	(4.5)	4.8	-.019	(4.1)
	$R^2 = .43$			$R^2 = .37$			$R^2 = .43$		

(a) Wage Gains (%) of Movers, by Union Status, and by Quit (Q) and Layoff (L)
NLS, 1967-1969

	Young Men (UM)		Young Men (CB)		Old Men (UM)		Old Men (CB)	
$U_{01}Q$	17.3	(2.8)	13.1	(2.6)	5.5	(.4)	16.7	(1.1)
$U_{01}L$	15.1	(2.5)	10.1	(1.6)	-8.1	(.4)	4.2	(.3)
$U_{10}Q$	-24.4	(4.3)	-23.6	(4.4)	-9.4	(.6)	-17.6	(1.0)
$U_{10}L$	-31.4	(5.0)	-37.0	(6.7)	-16.7	(.6)	-27.9	(1.8)
$U_{11}Q$	10.6	(1.6)	11.4	(2.0)	10.6	(.5)	10.2	(.5)
$U_{11}L$	3.4	(.6)	3.0	(.6)	3.8	(.6)	3.4	(.5)
$U_{00}Q$	5.1	(1.8)	6.0	(2.0)	-1.9	(.4)	-3.9	(.7)
$U_{00}L$	-4.8	(1.4)	2.3	(.6)	-12.0	(1.7)	-10.9	(1.5)

(b) Wage Gain of Movers, by Union Status and by Industry[*] Change
(B = moved between industries; W = moved within industries)

	Young Men (UM)		Young Men (CB)		Old Men (UM)		Old Men (CB)	
$U_{01}B$	18.8	(4.0)	13.4	(2.9)	2.4	(.1)	8.1	(.7)
$U_{01}W$	4.3	(1.5)	5.6	(1.7)	2.3	(0)	7.7	(.5)
$U_{10}B$	-33.5	(6.4)	-35.0	(7.7)	-9.1	(1.4)	-35.7	(2.4)
$U_{10}W$	-12.9	(1.7)	-13.0	(1.7)	-17.9	(.9)	-10.0	(.6)
$U_{11}B$	5.7	(.8)	6.9	(1.2)	3.7	(.3)	3.4	(.3)
$U_{11}W$	6.1	(1.2)	6.0	(1.1)	3.2	(.4)	2.9	(.4)
$U_{00}B$	-.8	(.3)	3.2	(1.3)	-28.9	(4.8)	-30.7	(4.8)
$U_{00}W$	7.2	(2.0)	8.0	(2.2)	15.1	(2.6)	13.9	(2.4)

Table A4

Wage Gains in MID

(a) Movers, by Union Status and by Quit and Layoff
MID, pooled

	All		Age $<$ 30		Age $>$ 30	
$U_{01}Q$	17.2	(4.4)	19.8	(3.6)	11.8	(1.8)
$U_{01}L$	-2.8	(.6)	-1.9	(.3)	-3.3	(.5)
$U_{10}Q$	.4	(.1)	-9.9	(1.6)	12.1	(3.0)
$U_{10}L$	-2.8	(.4)	-7.4	(1.0)	5.2	(1.0)
$U_{11}Q$	10.6	(2.7)	10.9	(1.8)	11.6	(2.0)
$U_{11}L$	5.6	(1.6)	-2.3	(.3)	10.7	(2.6)
$U_{00}Q$	9.0	(5.6)	12.4	(4.6)	4.6	(2.0)
$U_{00}L$	.7	(.3)	2.4	(.7)	-.8	(.3)

(b) Movers, by Union Status, and Industry Change

	All		Age $<$ 30		Age $>$ 30	
$U_{01}B$	10.1	(2.1)	11.2	(3.4)	4.2	(.6)
$U_{01}W$	-.4	(.7)	5.6	(.7)	-6.9	(2.0)
$U_{10}B$	-7.2	(1.7)	-10.1	(1.5)	-5.6	(.9)
$U_{10}W$	-.2	(.4)	-.6	(.1)	4.5	(.9)
$U_{11}B$	12.1	(2.7)	15.2	(2.0)	12.4	(2.1)
$U_{11}W$	3.0	(.8)	1.3	(.2)	5.5	(1.1)
$U_{00}B$	6.9	(3.8)	8.7	(3.0)	7.3	(2.8)
$U_{00}W$	3.2	(1.5)	8.5	(2.4)	-1.4	(.4)

(a) Mobility Rate Regression, MID, Cross-section pooled

	lnM		M	
Educ	-.050	(7.3)	-.015	(4.8)
X	-.048	(7.0)	-.031	(9.9)
x^2	.0004	(3.0)	.0002	(1.2)
Mar	-.092	(12.8)	-.050	(15.1)
Hlth	.001	(.3)	-.001	(.2)
Nun	-.454	(9.2)	-.421	(18.5)
Union	-.093	(2.5)	-.042	(2.6)

$$R^2 = .62 \qquad\qquad R^2 = .55$$

(b) Quit Rate Regressions, MID, Cross-section

	Union			Non-union		
	Mean	β	t	Mean	β	t
Q	.039			.083		
Intercept		.157	(6.2)		.275	(13.1)
Educ	11.4	-.0016	(1.1)	13.0	-.0053	(4.5)
X	22.4	-.0040	(3.5)	18.6	-.0068	(6.8)
x^2	666.2	.00004	(2.0)	500.0	.00007	(3.6)
T	9.0	-.0036	(2.5)	6.6	-.0080	(5.4)
T^2	151.8	.0001	(2.2)	99.4	.00027	(5.1)
Mar	.94	-.027	(2.0)	.91	-.012	(1.1)
Hlth	.09	.018	(1.6)	.08	.023	(1.9)
lnM*	-.0048	.010	(1.7)	.0038	.043	(6.8)

$$R^2 = .03 \qquad\qquad R^2 = .06$$

*Prior mobility rate.

Table A6

Estimates of Group Wage Differentials, MID, Pooled

I. A category was excluded if the number of union observations in it was less
than 20 in the full sample.

The coefflicients reported below represent the percent union differential
for each category. The dependent variable is log defalted average hourly
earnings. No intercept was used so that levels rather than contrasts could
be read for the group dummies (not reported here).

(1) The first set of dummies refers to detailed industries:

VARIABLE	PARAMETER ESTIMATE	T RATIO	
IU01	0.192876	2.6447	mining and extraction
IU02	0.030812	o,8054	metal industries (manuf.)
IU03	-0.027250	-0.9070	machinery, incl. electric (manuf.)
IU04	0.098406	3.2413	motor vehicles, other transportation equipment (manuf.)
IU05	0.120014	3.4501	other durables (manüf.)
IU06	0.117478	2.5718	food and kindred products (manuf.)
IU07	0.281642	4.1110	textiles, apparel, shoes (manuf.)
IU08	0.028735	0.3106	paper and allied products (manuf.)
IU09	-0.110420	-2.5741	chemical, allied products (manuf.)
IU10	0.349179	12.4312	construction
IU11	0.130612	3.9211	transportation
IU12	0.084606	1.4231	communication
IU13	0.168903	3.5534	other public utilities
IU14	0.273232	7.1365	retail trade
IU15	-0.043016	-0.8688	wholesale trade
IU16	0.110787	1.2200	repair service
IU17	0.142907	1.4794	personal services
IU18	0.184840	2.7868	printing, publishing, allied services
IU19	-0.073909	-0.9493	medical and dental and health services, public or private
IU20	0.136501	3.9973	educational services, public or private
IU21	0.219599	2.7867	professional and related services other than medical or dental
IU22	0.138157	4.3291	government, other than medical or educational services; NA whether other

F RATIO	4701.32
DFE	8463
R-SQUARE	0.9682

(2) The second set of dummies refers to states of the continental U.S.A. These states were dropped:

Alabama	New Hampshire
Colorado	New Mexico
Delaware	North Dakota
District of Columbia	Rhode Island
Idaho	South Carolina
Kansas	Vermont
Louisiana	West Virginia
Montana	Wyoming
Nebraska	
Nevada	

VARIABLE	PARAMETER ESTIMATE	T RATIO	
SU01	0.132901	1.7058	Arizona
SU02	0.268430	4.3792	Arkansas
SU03	0.177617	6.6824	California
SU04	-0.105243	-1.1892	Connecticut
SU05	0.106447	1.4246	Florida
SU06	0.005364501	0.0696	Georgia
SU07	0.004690769	0.1242	Illinois
SU08	0.029036	0.7859	Indiana
SU09	0.367906	8.0557	Iowa
SU10	0.062387	1.4132	Kentucky
SU11	0.283081	3.1941	Maine
SU12	0.160142	1.7876	Maryland
SU13	0.089620	1.8404	Massachusetts
SU14	0.026547	0.8876	Michigan
SU15	0.081187	1.8014	Minnesota
SU16	0.101594	1.3480	Mississippi
SU17	0.275612	7.9395	Missouri
SU18	-0.073130	-2.0036	New Jersey
SU19	0.083717	3.0567	New York
SU20	0.419438	5.6320	North Carolina
SU21	0.185828	6.9361	Ohio
SU22	-0.114174	-1.3375	Oklahoma
SU23	0.134037	3.1064	Oregon
SU24	0.105055	3.6566	Pennsylvania
SU25	0.454598	6.2741	South Dakota
SU26	0.407114	8.3154	Tennessee
SU27	0.149538	2.5615	Texas
SU28	0.165572	1.8644	Utah
SU29	0.184553	4.7105	Virginia
SU30	0.218756	4.3970	Washington
SU31	0.107793	2.1447	Wisconsin

F RATIO	4633.36	
DFE	11758	
R-SQUARE	0.9655	NOTE: NO INTERCEPT TERM IS USED.

(3) The third set of dummies refers to aggregate occupations cross-classified by aggregate industries:

VARIABLE	PARAMETER ESTIMATE	T RATIO	
Durables manufacturing			
IOU01	-0.161345	-2.7482	professional or managerial
IOU02	0.017291	0.2213	clerical and sales
IOU03	0.106036	3.9576	craftsmen, foremen and kindred
IOU04	0.223627	8.1542	operatives and kindred
IOU05	0.153847	1.9201	laborers and service workers
Non-durables manufacturing			
IOU06	-0.222012	-2.4278	professional or managerial
IOU07	0.020362	0.2382	clerical and sales
IOU08	0.102038	1.7847	craftsmen, foremen and kindred
IOU09	0.258602	6.3823	operatives and kindred
IOU10	0.136277	1.2682	laborers and service workers
Construction			
IOU11	0.380947	4.0230	professional or managerial
IOU12	0.394693	11.8224	craftsmen, foremen and kindred
IOU13	0.341520	3.8748	laborers and service workers
Transportation, communication, public utilities			
IOU14	0.035351	0.5346	professional or managerial
IOU15	-0.00629608	-0.0954	clerical and sales
IOU16	0.184263	4.2493	craftsmen, foremen and kindred
IOU17	0.201402	4.0272	operatives and kindred
IOU18	0.318112	3.1273	laborers and service workers
Wholesale and retail trade			
IOU19	0.073211	1.0787	professional or managerial
IOU20	0.123114	2.2794	clerical and sales
IOU21	0.359201	5.9317	operatives and kindred
IOU22	0.412222	5.1120	laborers and service workers
Educational services, private or public			
IOU23	0.171116	4.3852	professional or managerial
Those and all other services			
IOU24	0.068333	2.3846	all occupational categories, except IO23
Government, excl medical or educational services; NA whether other			
IOU25	0.148524	2.3907	professional or managerial
IOU26	0.145901	1.9485	clerical and sales
IOU27	0.247951	3.7062	craftsmen, foremen and kindred

F RATIO	4083.84
DFE	8284
R-SQUARE	0.9688

NOTE: NO INTERCEPT TERM IS USED.

Table A7

Determinants of In-Firm Training on Current Job

Variable	NLS Young Men, 1969-1971 White, Educ ≦ 12				MID, All Men, 1973-1975 White, Educ≦12	
	β	t			β	t
Educ	.013	(7.2)	.039	(6.0)	-.0007	(.01)
X	.009	(1.4)	.015	(1.1)	-.001	(0)
X^2	-.001	(.9)	-.0005	(.5)	-.001	(0)
T	.055	(3.4)			-.005	(.7)
T^2	-.005	(2.1) ?			.00	(.3)
Mar	.018	(1.1)			.014	(.1)
Hlth	-.044	(1.3)			.005	(.1)
Union	.014	(.7)			-.164	(26.8)
Min. Wage*	-.196	(2.7)			-.220	(2.4)
	R^2 = .06				R^2 = .08	

* Inverse of one plus state wage differential, multiplied by coverage
  (see Leighton and Mincer, 1980).