# THE EXIT-VOICE TRADEOFF IN THE LABOR MARKET: UNIONISM, JOB TENURE, QUITS, AND SEPARATIONS*

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This paper examines the effect of trade unionism on the exit behavior of workers in the context of Hirschman's exit-voice dichotomy. Unionism is expected to reduce quits and permanent separations and raise job tenure by providing a "voice" alternative to exit when workers are dissatisfied with conditions. Empirical evidence supports this contention, showing significantly lower exit for unionists in several large data tapes. It is argued that the grievance system plays a major role in the reduction in exit and that the reduction lowers cost and raises productivity.

In the exit-voice model of the social system [Hirschman, 1970, 1976] individuals react to discrepancies between desired and actual social phenomena in one of two ways: by the traditional free market mechanism of "exiting" from undesirable situations; or by directly expressing their discomfort to decision-makers through "voice." While little attention is paid to the labor market in Hirschman's book [1970], the exit-voice dichotomy provides a potentially fruitful framework for analyzing the major employee institution of capitalist economies—the trade union. From the perspective of the dichotomy, voice is embodied in unionism and the collective bargaining system by which workers elect union leaders to represent them in negotiations with management, while exit consists primarily of quits. A major feature of the model is a predicted tradeoff between the two adjustment mechanisms: when workers have a voice institution for expressing discontent, they should use the exit option less frequently and thus exhibit lower quit rates and longer spells of job tenure with firms.

Is unionism associated with lower quit rates and higher job tenure of workers, as predicted by the model? To what extent can any reduction in quits due to unionism be attributed to union "voice" as opposed to other routes of union effects, notably wage gains?

Despite a sizeable literature on labor turnover and on the economic effects of unions, extant empirical evidence provides no clear answer to these questions. The turnover literature has focused on quit rates for aggregated manufacturing industries, which provides only

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weak evidence on the behavior of individuals; has not treated job tenure or permanent separations as dependent variables; and has only rarely sought to estimate the effect of unionism (see Parsons [1972] for a useful summary). As a result, the impact of unionism on turnover has been at best estimated imprecisely, differing with sample and control variables (Burton and Parker [1969]; Stoikov and Raimon [1958]; Pencavel [1970]; Parsons [1977]; and Kahn [1977]). The union literature has dealt almost exclusively with union wage effects. Summarizing the state of knowledge in his textbook, Reynolds concluded that "it is questionable whether collective bargaining has produced a major change in the pattern of labor turnover" [1974, p. 568].

To provide a better test of the relation between unionism and exit behavior, this study analyzes data on *individuals* from three surveys—the National Longitudinal Survey (NLS), the Michigan Panel Survey of Income Dynamics (PSID), and the Current Population Survey (CPS)—which contain detailed information on the personal attributes of workers and characteristics of jobs that is better suited for analysis of individual behavior than industry aggregates. Longitudinal data in the NLS and PSID and retrospective data on the CPS allow for the analysis of the effects of union status and other variables on *actual quits or separations over time,* while survey questions relating to job tenure provide information on past exit behavior. By examining several bodies of information, each of which has certain weaknesses and strengths, I hope to obtain a better fix on the hypothesized behavioral relation from that given in previous studies.

The principal finding of the paper is that, *with wages and other measures of pecuniary rewards held fixed,* trade unionism is associated with significant, large reductions in exit behavior. Diverse calculations designed to adjust the union effect for potential omitted variable biases relating to union monopoly power or selectivity do not eliminate the union effect. While interpretation of the impact of unionism in terms of "voice" is open to some question, the empirical analysis provides support for the hypothesis that trade unions alter work place relations and worker behavior in ways *not* captured by standard monopoly wage models of the institution. Some evidence is presented that the observed reduction in exit is at least in part attributable to the operation of unions as a voice institution in the job market, though the fact that all collective bargaining involves voice in negotiations and in day-to-day work activities makes any definitive separation from the other components of unionism exceedingly difficult.

The paper is divided into five sections. The first sets out the reasons for expecting unionism to reduce exit. The second develops the methodology for the empirical analysis. Sections III and IV present the empirical results, with the former focusing on the effect of unionism on exit, and the latter probing the voice interpretation. The paper concludes with a brief evaluation of the economic consequences of the union-induced increase in the attachment of workers to firms.

## I. UNIONISM AND EXIT BEHAVIOR

Trade unionism can be expected to reduce exit behavior through "monopoly routes" of impact and through "voice routes" of impact.

In the context of the standard monopoly model of unions, exit is likely to be lowered by union-induced improvements in wages, fringes, and work conditions. Since the union wage effect is nonnegligible and high wages are likely to reduce quits significantly, the "monopoly wage" route of impact may be quite potent, and must be controlled in empirical analyses seeking to isolate the voice channels of concern. The major empirical problem in this study is to hold fixed monopoly wage effects of unionism, some of which may be unobserved, so as not to produce biased estimates of the union voice-exit tradeoff.

There are several ways in which the operation of unions as an institution of worker "voice" is likely to reduce exit behavior, producing the exit-voice tradeoff that is central to the model.[1]

First, unionism creates distinctive mechanisms for treating industrial relations problems that offer a substitute for classical exit behavior. Perhaps the most important such institution is the *grievance and arbitration system*, which offers dissatisfied workers who are considering quitting an alternative means of expressing discontent and possibly changing work conditions. Ninety-nine percent of major U. S. collective bargaining contracts provide for grievance procedures and 95 percent for arbitration [U. S. department of Labor, 1977, p. 94], making grievance and arbitration virtually synonymous with trade unionism. By contrast, only 30 percent of nonunion firms in the Bureau of National Affairs Personnel Policies Forum have formal grievance procedures, and only 11 percent allow outside arbitration to settle grievances not resolvable at lower levels [Bureau of National

1. For a detailed discussion of unions as a voice institution, see Freeman [1976] and Freeman and Medoff, forthcoming.

Affairs, 1968, p. 2]. The potential impact of a grievance system is clear: workers who feel themselves unfairly treated or who believe their supervisors erred in interpreting work rules will seek a solution through the grievance procedure before invoking the more drastic exit remedy. If the grievance is successful, the incentive to quit will be removed. Even if it is not and the aggrieved ultimately leaves, the overall rate of exit will be reduced as a result of the *delay* in the quit decision during the grievance procedure.[2]

The regular process of collective negotiation of labor contracts can also be expected to reduce exit behavior. Workers wanting new conditions who, in the absence of a bargaining alternative, might have quit will instead seek first to obtain the particular changes through bargaining. If some of the worker demands are met, quits are likely to be lower than would otherwise be the case. For work conditions and rules that are "public" to the enterprise, where standard public goods arguments suggest that enterprises would have great difficulty in eliciting true worker preferences, considerable mobility might be needed in the absence of unionism for these conditions and rules to be provided. Unions might obtain and aggregate preferences in such a manner as to produce the desired arrangements more efficiently, and with lower mobility in the market.

Union "voice" may also reduce exit by creating particular work rules and conditions of employment (which may or may not be costly to employers, once unionism is "in place") that are desired by workers, particularly what industrial relations experts call the *industrial jurisprudence system*. Under this system managerial authority is diluted by requiring that many work place decisions be made on the basis of *negotiated* rules, for instance seniority, as opposed to supervisory judgment (or whim). By straightforward application of compensating differential arguments, if workers desire these conditions and if they are provided largely by unions, then with *pay and other pecuniary benefits held fixed*, separation rates should be lower for union workers.

### *Voice in the Absence of Unions*

If "voice" institutions such as grievance/arbitration and individual jurisprudence are desirable work conditions that reduce

2. There are two conditions for the delay effect to operate. First, there must be some *nonzero* probability of redressing the grievance, so that the worker is willing to try the option. Second, the length of employment must be finite, for otherwise delays will not affect the steady-state solution. If, on average, the length of employment were initially, say ten years, then a delay in quitting for, say one half year, would reduce the quit rate from 0.10 to about 0.095, a nonnegligible though by no means large effect.

urnover, the question naturally arises as to why nonunion enterprises o not generally adopt them as part of a profit-maximizing trategy.

One reason for the general absence of voice or industrial juris-rudence practices in the nonunion sector is that the essence of voice s to reduce managerial power and create a dual authority channel vithin the firm. Such a change in power relations would be difficult o attain in the absence of a genuine independent union or union-like rganization. During the 1920s many firms experimented with so-alled "employee representation" plans designed to provide a nonu-ion voice mechanism for workers (see National Industrial Conference Board [1933]). Many of these plans ended in failure, despite the best ntentions, as workers are unwilling to express their desires for fear f retaliation by management and because of their own lack of power o affect decisions. Other plans led to the formation of company nions, which, in several industries, became the building blocks of ndependent unionism in the 1930s [Galenson, 1963]. Under current aw, of course, company unions are illegal. The dilemma is that if nanagement gives up power, it creates seeds of genuine unions; if it loes not, employee representation plans face severe difficulties. This s not to say that *no* nonunion firm will have a grievance/arbitration system, for some have such systems, in part to reduce worker desire or unions.[3] The point is that it is more difficult (costly) to institute n effective system in the absence of unions or union-type rganizations.

A second more subtle reason for the concentration of voice in-stitutions in the organized sector relates to the nature of price signals in unorganized and organized markets. In the unorganized market, the desire of workers for a given condition of work is conveyed by the marginal evaluation of the condition by the marginal worker, as re-flected in the reduction in wage he would take to obtain the condition. In the union market, the desire for the condition is conveyed by some average of preferences of workers: in a median voter model, by the (marginal) desire of the median worker; in a "consumer (worker) surplus"-maximizing model, by the average intensity of preferences for the condition. Assuming that "inframarginal" workers have greater desire for voice and industrial jurisprudence or that such systems generate worker surplus or both, there will be a more intense demand for the condition under trade unionism. Given fixed costs to setting

3. In discussion with ten large nonunion firms with grievance systems, all but one reported that the system was instituted in part to reduce worker desires for unionism (private interviews, 1978).

up voice institutions, the profit calculus might reject their development, while a benefit-cost calculation using the benefits to the median worker or taking account of consumer surplus would favor their development.[4]

## II. Empirical Analysis

Exit behavior is measured in this study by three variables: by job tenure, defined as the number of years a worker remains with a firm; by quits in a specified period; and by total separations in a period. Each of these variables has certain strengths and weaknesses for analysis of the exit-voice tradeoff. Tenure has the advantage of reflecting longer run and more permanent behavior than quits or separations because it relates to attachment between workers and firms over an extended period of time. The main disadvantage is the absence of data on characteristics of the job years earlier. Quits are useful because they measure worker behavior, which is at the core of the union voice model, but face the problem that the distinction between worker and employer-initiated changes is at least partially arbitrary (an employer may harass a worker to quit; a worker may quit because of potential plant closings or may perform poorly until fired). Separations do not have this problem but include such forms of mobility as those due to plant closings, which are not directly relevant to the model. By examining each measure, we are able to obtain a firmer set of conclusions than would otherwise be the case.

The decision to exit is analyzed in the framework of a probability model in which each person has a specific propensity to exit in a given year $Q$, dependent on a set of explanatory factors $X_i$, including unionism. Because exit is a dichotomous variable and probabilities are bounded by 0 and 1, the logistic provides an appropriate functional form for the relation:

$$(1) \qquad Q = \left(\left[1 + \exp - \sum_i (B_i X_i)\right]\right)^{-1} \qquad \text{with} \qquad \frac{dQ}{dX_i} = B_i Q(1-Q).$$

Tenure is treated as a backward waiting time variable dependent on $Q$. When $Q$ is fixed, the probability that tenure in year $n$, $T_n$, is a

4. This argument can be put more formally. Let $L(W,C)$ be the supply curve facing the firm, where $C$ = desired condition of work, and $L_W$, $L_C > 0$. Then the "supply price" for the condition, defined as the wage that is needed to maintain a given work force $L$ at various levels of $C$ can be written as $W(L,C)$, where $W_C < 0$. The marginal evaluation of the condition is $W_C(L,C)$; the median worker evaluation is $W_C(L_m,C)$, where $L_m$ is the median. The average of worker marginal evaluations is $[\int_0^L W_C(X,C)dX/L]/L$. If, as assumed in the text, $W_{CL} < 0$, then $W_C(L_m,C) < W_C(L,C)$, and $[\int_0^L W_C(X,C)]/L < W_C(L,C)$ so that whether the union represents the median worker or the average desire of workers, it will place a greater weight on the condition than the competitive signal. See Viscusi [1978] for elaboration of models of this type.

specified value $t$ can be written as

$$PR(T_n = t) = (1 - Q)^t Q, \tag{2}$$

which is a geometric distribution. Since $T_n$ reflects behavior over $t + 1$ periods of time, while changes in one time interval reflect behavior over only one period, $T_n$ conveys greater information about exit propensities than dichotomous quit or separation measures.

The mean of the *completed* tenure ($T$) distribution has a well-known relation to $Q$, which can be fruitfully used in analysis:

$$E(T) = \frac{(1 - Q)}{Q} = \exp\left(\sum_i B_i X_i\right). \tag{3}$$

If $Q$ depended on the $X$'s as in (1) and was independent of past tenure, the appropriate function form for $T$ would be the exponential, $T = \exp(\Sigma_i B_i X_i) + U_i$, where $U_i$ is a random error. With a fixed $Q$, renewal theory guarantees that the mean of the distribution of incompleted tenure equals the mean of the distribution of completed tenure, justifying use of the completed spell functional form. When, as appears to be the case, $Q$ is not constant but depends on the length of tenure (separations fall as tenure increases), the exponential is no longer appropriate. The functional form of the tenure equation will depend on the slope of the hazard function (the relation between cumulated tenure and the probability of separation) and can be quite complex. The most useful way of analyzing tenure in this case is to use the linear form, which can be viewed as a first-order Taylor series approximation to more complex functions:

$$T = \sum_i B_i X_i + U_i. \tag{4}$$

Calculations show that the linear function is much superior statistically to the exponential, presumably because of the dependence of $Q$ on tenure.

*Controlling for Monopoly Compensation Effects*

To isolate the nonmonopoly wage impact of unionism on exit, it is necessary to control carefully for other determinants of exit (themselves correlated with unionism), such as pecuniary and nonpecuniary compensation at the current job and at alternative jobs and personal characteristics like age or sex, which affect the transactions cost of mobility.

There are three problems in controlling for compensation at the current job. First, the surveys of individuals to be analyzed lack adequate information on fringe benefits, which are increased by unionism [Freeman, 1978a] and can be expected to reduce exit. This problem is dealt with by adjusting estimated coefficients on unionism for the omitted fringe variable using standard omitted variable bias formulae and outside information on the effect of unionism on fringes. Second, measures of nonpecuniary work conditions (above and beyond those represented by the voice or industrial jurisprudence conditions) are notoriously poor. Detailed industry and occupation dummies are used to narrow some of the possible range of variation among workers. In addition, measures that might be taken to reflect worker evaluation of nonpecuniary conditions, such as indicators of job satisfaction, are entered when available. Some effort is also made to control for omitted work conditions (and other factors) in the context of an unobservables model to be described shortly. Third, when tenure is the dependent variable in the analysis, there is a clear dual causal relation with tenure raising wages at the same time that high wages reduce $Q$ and raise tenure. Because of the likely magnitude of the coefficient on tenure in the wage equation, simultaneity can be expected to bias upward the estimated coefficient of wages on tenure.[5] This in turn is likely to bias downward the estimated re-

5. More precisely, let $W$ = wages, $T$ = tenure, $U_w$ = error in the wage equation and $U_t$ = error for tenure equation. Then, ignoring other factors for simplicity, we have the equation with wages as the dependent variable:

$$W = aT + U_w, \tag{1}$$

or rewritten with $T$ as the dependent variable,

$$T = (1/a)W - (1/a)U_w = a'W + V_t. \tag{2}$$

We also have the equation with tenure as the dependent variable:

$$T = \beta W + U_t. \tag{3}$$

Now the OLS estimate is

$$\hat{\beta} = \frac{\sum WT}{\sum W^2} = \beta + \frac{1/N \sum WU_t}{1/N \sum W^2}, \tag{4}$$

where $N$ = number of observations. So

$$\text{plim}\,\hat{\beta} = \beta + \frac{\text{plim}\ 1/N \sum WU_t}{\text{plim}\ 1/N \sum W^2}. \tag{5}$$

But since $W = (U_t - V_t)/(a' - \beta)$,

$$\text{plim}(1/N)\ WU_t = \text{cov}(WU_t) = (1/(a' - \beta))\sigma_U^2. \tag{6}$$

When the covariance between $V_t$ and $U_t$ is 0, the denominator of the right-hand expression in (5) becomes

$$\text{plim}(1/N)W^2 = \text{var}(W) = (1/(a' - \beta)^2)(\sigma_V^2 + \sigma_U^2). \tag{7}$$

But then

$$\text{plim}\,\hat{\beta} = \beta - (\beta - a')[\sigma_U^2/(\sigma_U^2 + \sigma_V^2)]. \tag{8}$$

We assume that $a \approx 0.02$ so that $a' \approx 50$ and that $\beta$ is much smaller. Thus plim $\hat{\beta} = \beta$ + positive term $> \beta$. The bias is upward.

gression coefficient on unionism. As inclusion of wages in the tenure calculations tends to work *against* the exit-voice hypothesis, I shall operate as if the causality were uni-directional and ignore the simultaneous bias.

Since the set of options facing a worker cannot be measured directly but must be inferred from his or her general characteristics, it is more difficult to obtain adequate measures of alternative compensation. The major indicators of alternatives are education, which should (wages fixed) raise exit propensities due to the better opportunities of the more educated or, in the context of a model of specific human capital, as a result of the inverse link between general and specific human capital at any fixed wage level; years of work experience, which should raise outside earnings and thus exit; and the state of the local labor market. Occupation and industry dummy variables can also be interpreted as reflecting outside opportunities. Because standard earnings regressions that include such variables as education, experience, and occupation rarely explain more than one-third of the variance in log earnings, however, it is unlikely that these variables will adequately index alternative opportunities. If, as seems reasonable, the unobserved components of alternative opportunities are correlated with current wages, as both current and alternative possibilities depend on omitted human capital or personal characteristics, statistical analyses will understate the negative effect of current compensation on exit and bias the estimated coefficient on unionism even if the unobserved alternative opportunities are uncorrelated with unionism.

The effect of the omitted components of alternative compensation on the estimated impact of wages and unionism on exit can be analyzed with regression formulae that do or do not control for the omitted factor. Let $W$ = compensation on the present job; $W_A$ = compensation in other jobs; $U$ = unionism; and $Q$ = the propensity to exit. Then, using subscripts to specify partial regression coefficients with the first subscript reflecting the dependent variable, the second the independent variable, and additional subscripts reflecting controls, the least squares coefficient relating $Q$ to $W$ and with $W_A$ omitted $[b_{QWU}]$ and the coefficient relating $Q$ to $U$ with $W_A$ omitted $[b_{QUW}]$ are linked to the "true" coefficients with $W_A$ included ($b_{QWUW_A}$ and $b_{QUWW_A}$) as follows:

$$b_{QWU} = b_{QWUW_A} + (b_{QW_AUW})(b_{W_AWU}) \tag{5}$$

$$b_{QUW} = b_{QUWW_A} + (b_{QW_AUW})(b_{W_AUW}), \tag{6}$$

where all of the coefficients are conditional on the other variables in the equation. The difference between the estimated and "true" coefficients, $(b_{QWU} - b_{QWUW_A})$ and $(b_{QUW} - b_{QUWW_A})$, depends on the signs and size of $b_{QW_AUW}$, $b_{W_AWU}$, and $b_{W_AUW}$. Increases in $W_A$ should increase exit, making $b_{QW_AUW}$ positive. The term $b_{W_AWU}$ is positive, by the assumption that the omitted factors are positively correlated with current pay. With $b_{QW_AUW} > 0$ and $b_{W_AWU} > 0$, there will be a downward bias in the estimated coefficient on wages. To obtain some notion of the magnitude of the bias, assume that $W$ and $W_A$ have similarly sized but oppositely signed effects on exit ($b_{QWUW_A} = -b_{QW_AUW}$). Then (5) can be rewritten to obtain the "correct" coefficient:

$$b_{QWUW_A} = b_{QWU}/(1 - b_{W_AWU}). \tag{7}$$

If the coefficient from the regression of $W_A$ on $W$, conditional on $U$ and all other variables is sizeable, say 0.5 to 0.7, then the true coefficient will be significantly above the estimated coefficient, suggesting that the coefficient on wages be raised considerably to estimate better the true wage effect.

The bias in estimating the effect of unionism on exit in (6) depends on $b_{W_AUW}$, whose sign is unclear. If union workers are *more* able than others, in ways not captured by $W$, $b_{W_AUW}$ will be positive, producing a downward bias in the absolute value of the estimated coefficient. Conversely, if union workers are, for whatever reason, less able than others, the absolute value of $b_{QUW}$ will overstate $b_{QUW_A}$. Assuming, as before, that $b_{QWUW_A} = -b_{QW_AUW}$, we obtain for the relation between the estimated and true coefficients on unionism

$$b_{QUWW_A} = b_{QUW} + (b_{W_AUW})(b_{QWUW_A}). \tag{8}$$

If, other factors fixed, $b_{W_AUW}$ is about equal to the union wage effect, say 0.10 to 0.20, the bias would be relatively modest, unless $b_{QWUW_A}$ were extremely large.

## III. Statistical Estimates of the Tradeoff

This section presents estimates of the impact of unionism on tenure, quits, and total *permanent* separations[6] for four individual data sets:[7] the NLS older male sample; the Michigan PSID sample;

6. These are to be distinguished from temporary separations due to temporary layoffs.

7. For discussions of these data sets see U. S. Department of Labor Research Monographs 15 and 16; Institute for Social Research (University of Michigan) and U. S. Bureau of the Census.

TABLE I

ESTIMATES OF THE EFFECT OF TRADE UNIONISM AND OTHER VARIABLES ON EXIT BEHAVIOR: OLDER MALE NLS SAMPLE, 1969–1971[a]

| Dependent variable | Mean and standard deviation | | Estimated coefficient and standard error | | | | | | | | $R^2$ (−ln likelihood) |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | Union | Nonunion | Union[b] | Log earnings | Years of schooling | Age | Retirement plan (1 = yes) | Job dissatisfaction[c] | Tenure 1969 | Other controls[d] | |
| 1. Tenure, 1969 | 17.4 | 13.2 | 3.64 | 3.49 | 0.11 | 0.41 | | | | 1–7 | 0.20 |
| (linear) | (10.6) | (11.0) | (0.58) | (0.59) | (0.08) | (0.06) | | | | | |
| 2. Tenure, 1969 | | | 2.96 | 2.41 | 0.07 | 0.47 | 4.65 | 0.02 | | 1–7 | 0.23 |
| (linear) | | | (0.58) | (0.60) | (0.09) | (0.06) | (0.57) | (0.29) | | | |
| 3. Tenure, 1969 | | | 0.56 | 1.59 | −0.00 | 0.15 | | | | 1–7 | 0.03 |
| (exponential form) | | | (0.35) | (0.29) | (0.05) | (0.04) | | | | | |
| 4. Quit, 1969–71 | 0.01 | 0.07 | −2.17 | −0.26 | −0.03 | −0.10 | | | | 1–7 | (267) |
| | (0.09) | (0.26) | (0.43) | (0.24) | (0.04) | (0.03) | | | | | |
| 5. Quit, 1969–71 | | | −2.07 | −0.01 | −0.02 | −0.10 | −0.85 | 0.36 | | 1–7 | (259) |
| | | | (0.43) | (0.27) | (0.04) | (0.03) | (0.30) | (0.14) | | | |
| 6. Quit, 1969–71 | −1.85 | −0.85 | | 0.19 | −0.02 | −0.08 | −0.24 | 0.35 | −0.16 | 1–7 | (228) |
| | | | (0.44) | (0.30) | (0.05) | (0.03) | (0.31) | (0.15) | (0.03) | | |
| 7. Separations, | 0.07 | 0.14 | −0.87 | −0.09 | −0.03 | −0.07 | | | | 1–7 | (521) |
| 1969–71 | (0.26) | (0.35) | (0.21) | (0.18) | (0.03) | (0.02) | | | | | |
| 8. Separations, | | | −0.81 | −0.08 | −0.02 | −0.06 | −0.51 | 0.27 | | 1–7 | (514) |
| 1969–71 | | | (0.21) | (0.19) | (0.03) | (0.02) | (0.20) | (0.10) | | | |
| 9. Separations, | | | −0.63 | 0.16 | −0.01 | −0.05 | −0.14 | 0.26 | −11 | 1–7 | (463) |
| 1969–71 | | | (0.22) | (0.21) | (0.03) | (0.02) | (0.20) | (0.10) | (0.01) | | |

*Source.* Based on National Longitudinal Survey data tapes for men aged 45–59 in 1966, 1973 tape version.

a. Number of observations = 1,735. Sample consists of private wage and salary workers who reported earnings and other variables in 1969 and who were employed in 1969 and 1971. Coefficients in lines 1–2 based on linear equation $T = \Sigma B_i X_i$. Coefficients in line 3 are based on exponential equation $T = \exp \Sigma B_i X_i$. Coefficients in lines 4–9 are based on logistic equation $Pr(Q) = 1/(1 + \exp -\Sigma B_i X_i)$.

b. Unionism measured by 0–1 dummy variable for whether or not wages are set by collective bargaining. This variable is virtually identical with a union member variable.

c. Job satisfaction measured as standard normal deviation, with positive values reflecting greater than average satisfaction and negative values the converse.

d. Other controls are defined as 1 = 9 dummy variables for industry; 2 = 6 dummy variables for occupation; 3 = race; 4 = number of dependents; 5 = 3 dummy variables for region; 6 = size of local labor market; 7 = rate of unemployment in local market.

the CPS sample; and the NLS younger male sample. Maximum likelihood calculations are used to estimate the determinants of quits and separations; nonlinear and linear least squares are used to estimate the determinants of tenure. The analysis finds a sizeable effect of unionism on all of the indicators of exit propensity, which is maintained after the various adjustments and corrections suggested in Section II.

### *Older Male NLS Results*

The estimated effect of unionism and other important explanatory variables on exit behavior in the older male NLS sample is given in Table I. The table records the mean and standard deviation of the measures of exit for union and nonunion workers[8] and the coefficients on unionism (measured by an 0–1 dummy variable for workers whose wages are set by collective bargaining), log earnings, dummy variables for presence of a retirement plan (=1 when a firm has a plan), an index of job dissatisfaction, and on lagged tenure, entered as an explanatory variable in some quit or separations equations. As specified in the notes to the table, the sample is limited to workers who remained in the labor force in the period and thus excludes persons who retired.

What stands out in the table are the differences in the exit propensities of union and of nonunion workers. The means and standard deviations show that union workers have considerably more tenure than nonunion workers (17.4 years versus 13.2 years) and have much lower quit rates (a miniscule 1.0 percent compared to 7.2 percent for nonunion workers) and lower separation rates (7.0 percent for union workers versus 14.0 percent for nonunion workers). Since the absolute difference in quit rates between union and nonunion workers (6.0) is almost equal to the absolute differences in the total separation rates, which also include employer-initiated changes, there are essentially no differences in other separations between union and nonunion workers.

In the multivariate statistical analysis, where a wide variety of other factors is held fixed, the different exit propensities of union and nonunion workers are evinced in the sizeable significant coefficients accorded unionism.

In the linear tenure calculations, unionism obtains a coefficient

8. Note that we can compare the exponential and linear forms directly because they have the same dependent variable $T$. Thus, the $R^2$ is the correct measure of fit.

that ranges from 1.8 to 3.6 years, indicating that upward of half of the mean difference in years of tenure is, in fact, attributable to trade unionism, other factors held fixed. Introduction of the retirement variable in line 2 reduces the union coefficient in the linear form but still leaves a sizeable significant effect. By controlling for retirement plans, which are more prevalent among union workers, the analysis assumes that all of the effect of unions on pensions represents monopoly wage gains and that none represents the effective transmission of worker preferences via voice, possibly understating the voice component of the union impact. The job satisfaction index, also entered in line 2 as a crude indicator of unmeasured nonwage aspects of the work place or of alternative opportunities (which, if especially good, should decrease satisfaction, all else the same), obtains an insignificant effect. In both lines 1 and 2, the coefficient on log earnings is large, of a similar magnitude to that on unionism. The estimated exponential tenure equation (equation (5)) in line 3 yields a smaller coefficient on unionism than on earnings but fits the data so poorly as to be readily rejected in favor of the linear form.[9]

The estimated effect of unionism on the probability of quits or total separations in lines 4–8 tells a similar story. In line 4, unionism obtains a sizeable highly significant coefficient of −2.17, which implies that unionism causes a sizeable reduction in the probability of quitting. Addition of the index of job dissatisfaction, which significantly lowers quits, and of the retirement plan variable in line 5 reduces the coefficient on unionism modestly. Comparable results are obtained in the logistic separation equations in lines 7 and 8, where the estimated effect of unionism varies from −0.81 to −0.87 with a standard error of 0.21. Given the higher level of separations in the sample (a mean of 0.11), these parameters translate into differences in the probability of separation of about 0.08. In sharp contrast to the sizeable significant impact of unionism, both the quit and separation calculations accord a small insignificant effect to wages, which makes the union exit tradeoff look quite powerful by comparison.

In lines 6 and 9, the log of tenure in 1969 has been added as an additional control in the quit and separation calculations. Inclusion of tenure can be interpreted in two possible ways. In the simplest interpretation, tenure is just another control variable, reflecting the dependence of exit propensities on cumulated tenure. Alternatively, however, tenure can be taken to represent unobserved factors that

9. A potential problem with this procedure is that it ignores possible selectivity bias in the group that separate from their job.

affect both future quits and separations and past employment stability. Formally, let $T$ and $Q$ depend on some omitted person or job-specific factor ($F_i$) as well as on unionism:

$$T = \alpha^T UN_i + F_i + \epsilon_{iT} \tag{9}$$

$$Q = 1/[1 + \exp(-\alpha UN_i - \lambda F_i + \epsilon_{iQ})], \tag{10}$$

where $T$ superscripts are used to denote coefficients in the tenure equation. In (9) and (10) the residuals have two parts: $F_i$ which reflects the omitted factor (given a unit coefficient in (9) and a scaling of $\lambda$ in (10)) and equation-specific components $\epsilon_{iQ}$, $\epsilon_{iT}$, which are themselves uncorrelated and are uncorrelated with the independent variables. The econometric problem is that $F_i$ is correlated with unionism (and possibly other explanatory variables also).

Solving (9) for $F_i$ and substituting into (10) yields

$$Q_i = 1/(1 + \exp[-(\alpha - \alpha^T\lambda)UN_i - \lambda T + \epsilon_{iQ} + \epsilon_{iT}]), \tag{11}$$

which removes the correlation between $UN_i$ and the residual but introduces a correlation between $T$ and the residual since $E(T\epsilon_{iT}) > 0$. If an instrument for $T$ could be found, such as an independent variable that does not enter $Q$, consistent estimates of the coefficient on tenure could be made. In the absence of instruments, $\lambda$ will be underestimated in (11), and given a positive correlation between $UN_i$ and $T_i$, the coefficient on unionism overstated. As can be seen in (11), however, the coefficient on unionism with the inclusion of tenure is $\alpha - \alpha^T\lambda$, rather than $\alpha$ itself. Given $\alpha^T > 0$, this implies an underestimate of the effect of unionism on quits. Inclusion of tenure in the regressions does not totally resolve econometric problems but does set up a difficult test of the union impact. From this perspective, the continued sizeable effect of unionism in lines 6 and 9, where tenure is entered as an explanatory factor, can be viewed as highly supportive of the postulated union-exit tradeoff. Unionized workers are much less likely to quit or separate from employers than nonunionized workers who have the same job tenure and thus having the same "stability" history.

To analyze the magnitude of biases due to inadequate controls on alternative opportunities, it is necessary to obtain estimates of $b_{W_AWU}$ and $b_{W_AUW}$. With such estimates, (7) and (8) can be used to evaluate the effect of the omitted part of alternative opportunities. The link between alternative wages and current wages and between alternative wages and unionism can be estimated, albeit crudely,[10] by examining workers who changed employers from 1969 to 1971, with $W_A$ measured by wages on the new 1971 job. Regressions of log wages in 1971 on log wages in 1969 and union status in 1969 and the other 1969 control variables for the 11 percent of the older male NLS sample who changed employers yield the following coefficients (standard errors in parentheses):[11] $\hat{b}_{W_A WU} = 0.32$ (0.09); $\hat{b}_{W_A UW} = 0.12$ (0.11). Using equations (7) and (8), these values of $b_{W_A WU}$ and $b_{W_A UW}$ imply a sizeable increase in the estimated effect of wages on exit and a modest increase in the effect of unionism on exit. Adjusting the figures in line 2, where wages were accorded a sizeable effect on tenure, for example, yields $b_{QWUW_A} = 1.47\ (2.47) = 3.54$ and $b_{QUWW_A} = 2.96 + 3.54\ (0.12) = 3.38$. Even with the adjustment, the estimated impact of unionism on tenure is large relative to the estimated impact of wages on tenure. Adjusting the quit and separation equations has even less effect on the results.

In sum, the data on the effect of unionism on job tenure, quits, and separations in the older male NLS sample show union workers to be much more attached to their firms than comparable nonunion workers.

*Michigan PSID Data Set*

Table II presents estimates of the effect of unionism on tenure and quits using data from the Michigan PSID sample. This data set covers the entire population and thus gives a more inclusive picture of the trade union impact than that in the older male NLS sample. In the PSID, information on unionism and mobility is available on an annual basis for the years 1968 to 1974, which provides data on exit in each of five separate years (1968–1969, 1969–1970, and so forth). To obtain a single large sample covering all of the years, individual year observations were pooled into one data set, with observations

10. This equation included the same controls as in Table I. The $R^2$ was 0.462 and the *SEE* was 0.552. Tenure was included as an explanatory variable. There were 207 observations.

11. These are based on 525 separations from 1969 to 1970. Comparable estimates for other years give smaller coefficients on unionism and, except in one case, on ln wages as well:

| | $b_{W_A UW}$ | $b_{W_A WU}$ |
|---|---|---|
| 1970–71 ($n$ = 473) | 0.04 (0.06) | 0.41 (0.04) |
| 1971–72 ($n$ = 493) | 0.10 (0.06) | 0.32 (0.04) |
| 1972–73 ($n$ = 590) | 0.11 (0.06) | 0.36 (0.04). |

In each regression, all of the control variables listed in Table II were included in the calculations. Tenure was excluded.

TABLE II

ESTIMATES OF THE EFFECT OF UNIONISM ON EXIT, MICHIGAN PANEL SURVEY OF INCOME DYNAMICS, 1968–1974[a]

| Dependent variable and # of observations | Mean and standard deviations | | Coefficients and standard errors | | | | | Additional controls[b] | (−ln $R^2$ likelihood) |
|---|---|---|---|---|---|---|---|---|---|
| | Nonunion | Union | Union | Log wage | Schooling | Age | Lagged tenure | | |
| Tenure, 1968 | 5.8 | 10.2 | 2.72 | 2.24 | −0.13 | 0.18 | | 1–9 | 0.243 |
| All family units, | (7.4) | (8.6) | (0.27) | (0.20) | (0.03) | (0.01) | | | |
| $n$ = 2,597 (linear) | | | | | | | | | |
| Tenure, 1972 | 6.6 | 8.8 | 1.06 | 2.60 | 0.04 | 0.22 | | 1–9 | 0.320 |
| All family units, | (8.0) | (8.2) | (0.27) | (0.19) | (0.04) | (0.01) | | | |
| $n$ = 2,322 (linear) | | | | | | | | | |
| Tenure, 1972 | | | 0.77 | 0.76 | −0.01 | 0.09 | | 1–9 | 0.160 |
| All family units, | | | (0.17) | (0.14) | (0.02) | (0.01) | | | |
| $n$ = 1,175 (exponential)[c] | | | | | | | | | |
| Annual quits, 1968–1974 | 0.090 | 0.058 | −0.469 | −0.238 | 0.001 | −0.045 | | 1–9 | (5,603) |
| $n$ = 21,173 | (0.286) | (0.235) | (0.070) | (0.051) | (0.009) | (0.002) | | | |
| Annual quits, 1968–1974 | | | −0.438 | −0.197 | 0.002 | −0.040 | −0.038 | 1–9 | (5,570) |
| $n$ = 21,173 | | | (0.070) | (0.051) | (0.009) | (0.002) | (0.005) | | |
| Annual separations | 0.130 | 0.092 | −0.412 | −0.253 | −0.009 | −0.044 | | 1–9 | (7,272) |
| 1968–1974 | (0.336) | (0.289) | (0.058) | (0.042) | (0.008) | (0.002) | | | |
| $n$ = 21,173 | | | | | | | | | |
| Annual separations | | | −0.379 | −0.205 | −0.008 | −0.038 | −0.043 | 1–9 | (7,212) |
| 1968–1974 | | | (0.058) | (0.043) | (0.007) | (0.002) | (0.004) | | |
| $n$ = 21,173 | | | | | | | | | |

*Source*. Based on Michigan Panel Survey of Income Dynamics, 1974 tape version.

a. Sample consists of private wage and salary workers who reported relevant variables in the given years. Coefficients in lines 1–2 are based on linear equation $T = \Sigma B_i X_i$. Coefficients in line 3 are based on exponential equation $T = \exp \Sigma B_i X_i$. Coefficients in lines 4–7 are based on logistic equation $PR(Q) = 1/(1 + \exp - \Sigma B_i X_i)$.

b. Additional controls are defined as 1 = 5 industry dummies; 2 = 8 occupation dummies; 3 = dummy for shortage of workers in local areas (1 = yes); 4 = unemployment in local areas; 5 = dummy for whether wages are high (= 1) or low; 6 = number of dependents; 7 = sex; 8 = race; 9 = dummy for whether person was in special low-income sample.

c. Sample size reduced due to computational problems in maximum likelihood search procedure.

consisting of dichotomous exit variables from year $t$ to $t + 1$ linked to the characteristics of the worker and job in year $t$. The pooled sample contains 21,173 observations, with certain individuals deleted in particular years due to changes in the survey sample. Since quits can be treated as independent from year to year while tenure cannot, tenure is examined in the initial year 1968 and in the intermediate year 1972.

The calculations in Table II confirm the basic finding of a significant union-exit relation. The mean values of tenure and quits for union workers and nonunion workers show substantially less exit among unionists, uncorrected for differences in other factors. Union members had 4.4 years more tenure in 1968 and 2.2 years more in 1972 (when the sample was somewhat smaller due to deletions). Over the period 1968–1973 union workers had an average annual quit rate of 5.8 percent compared with 9.0 percent for other workers, a quantitatively large and statistically significant difference. The separation rates also differed noticeably, with a rate for union workers of 9.2 percent compared with 13.0 percent for nonunion workers. In the regressions, which include controls for wages, years of schooling, occupation, local labor market conditions, age, sex, race, and industry, as specified in the table, unionism is always accorded a significant impact. In the OLS tenure regressions, the union coefficient varies noticeably between 1968 and 1972, but is large absolutely and relative to its standard error in both cases. While the linear form fits better than the exponential, the fit of the exponential in line 3 (limited to a smaller sample due to computational problems) is much superior to that in Table I and yields a union coefficient somewhat larger at the mean value of variables (1.32) then from the better-fitting linear form (1.06). In the pooled quit equations in lines 4 and 5, which first exclude and then include tenure in the precedent period, the estimated impact of unionism on the probability of quitting ranges from 0.036 to 0.038 points at the mean level of quits. Comparable results are obtained in the pooled separation equations in lines 6 and 7, with the union coefficient estimated to lower separation at the mean level by 0.044 with tenure excluded from the equation and by 0.040 with tenure included as an explanatory variable.

The magnitudes of the union and wage coefficients in the quit and separation equations in Table II differ noticeably from those in Table I. The impact of unionism is smaller in the Michigan PSID than in the NLS Older Male sample, possibly because of the inclusion of younger workers and women whose exit behavior is less likely to be affected by unionism than that of older men; the impact of wages on exit is significantly negative in lines 5–8 of Table II, in contrast to its weak effect in the estimates of Table I. With a smaller union and larger wage coefficient, the union-exit tradeoff becomes relatively more moderate compared with the wage-exit relation.

The wage and union coefficients in Table II can be "corrected" for potential bias due to lack of adequate data on alternative opportunities in the same fashion as done previously. Regressing log wages of persons who changed jobs in the Michigan sample on their previous wage and union status and other aspects of the job for each year yielded the following estimates for 1969–1970: $b_{W_A UW} = 0.13$ (0.05); and $b_{W_A WU} = 0.38$ (0.04) and comparable (somewhat lower) estimates for other years.[12] These figures suggest that in the tenure equation of line 2 the wage coefficient be raised to 4.19 and the union coefficient raised to 1.60, while in the quit and separation equations in lines 5 and 7 the wage coefficients be changed to −0.39 and −0.33 and the union coefficients be changed to −0.52 and −0.42, respectively. Even with these adjustments, the union impact is sizeable relative to the wage impact: an increase in wages of over 100 percent is needed to reduce quits (separations) by as much as unionism, while an increase in wages of over one-third is needed to raise tenure by as much as unionism.

Because the PSID sample lacks any information on fringe benefits, it is also important to adjust the coefficients for the potential bias due to absence of fringe data. To obtain some notion of the magnitude of the bias, consider the least squares equation linking partial regression coefficients to the coefficients that are not partial for the omitted variable:

$$b_{QUF} = (b_{QU} - b_{QF}b_{FU})/(1 - r_{FU}^2), \tag{12}$$

where

$b_{QUF}$ = the coefficient corrected for the omitted fringe variable
$b_{QU}$ = estimated coefficient on exit
$b_{FU}$ = the regression coefficient linking fringes to unions
$b_{QF}$ = the regression linking exit to fringes
$r_{FU}$ = the correlation of fringes and unionism,

12. Adjusting for $W_A$ we have

$$b_{QWUW_A} = 1.61(-0.57) = -0.92$$

and

$$b_{QUWW_A} = 0.52 + 0.13(0.92) = -0.74.$$

Adjusting for the effect of fringes on the union coefficient yields

$$b_{QUF} = [-0.74 - 0.13(0.92)]/0.91 = -0.68.$$

TABLE III

UNIONIZATION AND QUITS IN THE CURRENT POPULATION SURVEY TAPES, MAY 1973–1975[a]

| Dependent variable | Mean and standard deviation | Coefficient and standard error | Approximate logistic coefficient[b] |
|---|---|---|---|
| Quit and unemployed | 0.004 (0.007) | | |
| Coefficients and standard errors on explanatory variables | | | |
| Union | 0.23 (0.42) | −0.0026 (0.0006) | −0.62 |
| Log hourly earnings | 1.24 (0.60) | −0.0024 (0.0005) | −0.57 |
| Schooling | 12.3 (3.0) | −0.00001 (0.0001) | 0.00 |
| Age | 35.9 (14.4) | −0.0014 (0.0002) | 0.33 |

| Other controls | Numbers of controls |
|---|---|
| Industry dummies | 46 |
| State dummies | 26 |
| Sex dummy | 1 |
| Race dummy | 1 |
| Occupation dummies | 11 |
| Number of dependents | 1 |
| Marital status dummies | 3 |
| Year dummies | 2 |

*Source.* U.S. Bureau of the Census, Current Population Reports, May 1973, 1974, 1975 tapes.

a. Number of observations = 98,593. Sample consists of private wage and salary workers reporting relevant variables.

b. Estimated by dividing linear coefficient by $SSR/n$, where $SSR$ = sum of squared residuals, and $n$ = number of observations. See Nerlove and Press [1973] for discussion of this approximation.

and where all the coefficients are partial with respect to the other variables in the model.

Available information and the likely magnitudes of the coefficients in (12) suggest only a moderate upward bias in the estimated effect of unions, of at most 25 percent. For $b_{FU}$ and $r_{FU}$, estimates in Freeman [1978a, table 3] indicate that, conditional on straight-time pay, industry dummies, and other control variables, $r_{FU} = 0.30$ while $b_{FU} = 0.11$. An extremely high estimate of $b_{QF}$ would be the coefficient obtained on log wages in the regressions: since fringes constitute no more than one-third of the wage bill, this implies that a dollar of fringes is three times as effective in reducing exit as a dollar of wages. With these estimates, the union coefficient in line 2 of Table II (where wages have their greatest impact relative to unionism) is reduced to 0.79, while there is virtually no impact on the coefficients in other lines. Unless the magnitudes of the estimated coefficients linking unionism to fringes and fringes to exit are markedly off, correction for the omission of fringes still leaves a sizeable union coefficient.

*CPS Data Set*

The largest data set with information on union status and quit behavior is the Current Population Surveys for the month of May, which contain information on persons who quit *and* are unemployed at the time of the survey, an extreme form of exit, since most quitters have a job in hand, and information on unionism on the past or current job. These surveys have the advantage of offering an especially large sample size, which permits many industry, occupation, and area control variables but the disadvantage of relating to a distinct and relatively small group of quitters. To obtain as large a sample of unemployed quitters as possible, I amalgamated the May 1973, 1974, and 1975 CPS surveys into a single sample with over 98,500 persons. With this size of the sample and with numerous controls for occupation and detailed industry, maximum likelihood estimation of the logistic function became computationally infeasible. Instead, a linear probability model was fit. The linear coefficients can be transformed to obtain a first-order inverse Taylor series approximation of the logistic parameters by multiplication by $n/SSR$, where $n$ = sample size, $SSR$ = estimated sum of squared residuals (see Nerlove and Press [1973] for a discussion of this transformation).

Estimates of the effect of unionism, earnings, schooling, and age on quits in the CPS sample are given in Table III. The linear regression coefficient on union status gives the, by now familiar, result of a negative significant coefficient that is of approximately the same magnitude as the coefficient on log wages. With diverse characteristics fixed, unionism is associated with a 0.26 percent lower quit (into unemployment) rate, which is of similar magnitude to the effect of earnings in the sample. Transformed into logistic curve parameters, the coefficient on unionism is about 50 percent larger than that obtained in the quit calculation of Table II while the coefficient on earnings is roughly twice as great. Still, it would take almost a 100 percent increase in the log of average hourly earnings to reduce the probability of quitting by as much as the switch from nonunion to union status.

If the partial correlations between $W$, $W_A$, and $UN$ from the Michigan sample are taken to apply to the CPS sample as well (since both samples cover the entire population) and the CPS estimates adjusted for omission of $W_A$, the wage and union coefficients are raised noticeably. In the logit form the coefficient on wages rises to $-0.74$, while the union coefficient rises to $-0.71$. If, in addition, the union (but not the wage) coefficient is reduced for the omission of fringe benefits using the same procedure and data as before, the union impact drops to $-0.68$.[13] Even with these adjustments, the finding of a steep union-exit tradeoff remains; the switch from nonunion to union status reduces quits by as much as 0.92 log wage increase.

*Young Male NLS Sample*

Estimates of the effect of unionism on exit behavior were also made with the young male NLS survey, which contains information on the union status and tenure of men aged 17–27 in 1969 and in 1971. Because of a lack of direct information on the causes of job changes between 1969 and 1971, the analysis is limited to tenure and to separations, defined as having a different job in 1971 than in 1969. The analysis eliminates students and focuses on regular workers.

The discussion of the difference between the supply price in union and in nonunion markets of Section I suggested that unionism represents older "average" workers to a greater extent than younger mobile workers. This in turn suggests that the impact of unionism on exit would be weaker among younger, more mobile workers than among others, and might even possibly be positive rather than negative. The calculations summarized in Table IV show the expected weaker effects. The differences in mean tenure and separation rates between young union and young nonunion workers are proportionately smaller than in the previous samples. The estimated effects of unionism in the linear tenure regressions in columns 1 and 2 and in the logistic separation calculations in columns 3 and 4 also yield relatively small and moderate union effects, both absolutely and relative to the estimated effects of wages. While the coefficient on unionism in the 1969 tenure regression is larger than the coefficient on wages, in all of the other calculations, the coefficient on wages is larger, which contrasts to the general pattern in previous analysis. Adjustments in the coefficients for the omission of major components of alternative compensation enhance this greater effect of wages than of unionism on the exit behavior of the young. In the young male NLS sample, regressions of the log of wages of persons who changed employers from 1969 to 1971 on unionism, log wages in 1969, and the control variables in Table IV yield estimates of $b_{W_A W U}$ of 0.43 (0.04) and of $b_{W_A U W}$ of 0.005 (0.03), where numbers in parentheses are standard errors.[14] Using the formula in (7) and (8) to adjust the coefficients in the table

13. The equation included the same controls as Table IV. Tenure was excluded as an explanatory variable.

14. When a person never quits, the best estimate of his individual propensity in the fixed logit is that he has a constant term that is $-\infty$. When a person quits always, the model fits best with an individual constant of $\infty$. Hence, these persons drop from the sample.

TABLE IV

ESTIMATES OF THE EFFECT OF UNIONISM ON TENURE AND SEPARATIONS: YOUNGER MALE NLS SAMPLE 1969–1971[a]

| Dependent variable | Tenure 1969 | Tenure 1971 | Separations[b] | 1969–1971 |
|---|---|---|---|---|
| Mean and standard deviation | | | | |
| Union | 1.88 (1.85) | 3.12 (2.57) | 0.47 (0.50) | |
| Nonunion | 1.66 (1.93) | 2.74 (2.65) | 0.58 (0.49) | |
| Coefficients and standard errors on | | | | |
| Explanatory variables | | | | |
| Unionism | 0.28 (0.11) | 0.36 (0.14) | −0.20 (0.13) | −0.20 (0.13) |
| ln wage | 0.23 (0.14) | 0.62 (0.16) | −0.73 (0.18) | −0.67 (0.17) |
| Years of schooling | −0.01 (0.02) | 0.00 (0.02) | 0.01 (0.03) | −0.00 (0.03) |
| Age | 0.08 (0.02) | 0.16 (0.03) | −0.13 (0.02) | −0.08 (0.03) |
| Years of work experience | 0.16 (0.02) | 0.13 (0.02) | −0.001 (0.002) | 0.02 (0.02) |
| Tenure | | | | −0.11 (0.03) |
| Other controls | | | | |
| Industry dummies | 9 | 9 | 9 | 9 |
| Region dummies | 3 | 3 | 3 | 3 |
| Occupation dummies | 6 | 6 | 6 | 6 |
| Race dummy | 1 | 1 | 1 | 1 |
| SMSA dummy | 1 | 1 | 1 | 1 |
| Number of dependents | 1 | 1 | 1 | 1 |
| $R^2$ (−ln likelihood) | 0.252 | 0.243 | 1,050 | 1,076 |

*Source.* Based on National Longitudinal Surveys data tapes for men aged 14–24 in 1966, 1973 tape version.

a. Number of observations = 1,742. Sample limited to private wage and salary workers out of school in 1969 and 1971 who report relevant variables. Coefficients in columns 1 and 2 based on linear tenure equation $T = \Sigma\beta_i X_i$. Coefficients in columns 3 and 4 based on logistic equation $Q = 1/(1 + \exp - \Sigma\beta_i X_i)$.

b. Separations 1969–1971 calculated from 1971 tenure, with separation = 1 if person had less than two years of tenure.

raises the estimated effect of wages by 75 percent but leaves the union coefficients essentially unchanged. Because deferred fringes are likely to be received too far in the future to affect the young, further adjustments are not warranted. The principal conclusion to be drawn from Table IV is that unionism has a smaller impact on the exit behavior of young workers than on the exit behavior of older workers.

*Selectivity versus Behavior*

The analysis thus far has shown that with monopoly compensation gains fixed, unionism raises tenure and reduces quits. Is the estimated reduction in exit, with wages held fixed, due to unionization of relatively more stable persons, or is it due to actual changes in behavior caused by the specific work relations associated with the union institution?

The longitudinal data files provide a possible means of differentiating between these two effects and isolating the behavioral aspects of unionism of concern. With longitudinal data on the *same* person over time, that person's exit behavior when he or she is unionized and when he or she is not unionized can be compared, thereby eliminating the personal propensity to be a stable worker. The most direct way of controlling for individual effects is to add individual constants to the logistic probability function and estimate a "fixed effect logit model" (see Chamberlain [1978] and Freeman [1978b] for detailed discussions of this model). The fixed effect procedure has the advantage over possible random effect models of not requiring knowledge of a particular distribution for the individual propensities. The fixed effect model yields consistent estimates regardless of the distribution of individual propensities, whereas random effects models based on specific distributions yield inconsistent estimates when that distribution is incorrectly specified.

The Michigan PSID data set provides sufficient number of observations on individuals over time for the fixed effect logit to be estimated and has, accordingly, been used to test the selectivity interpretation of the union effect. Since, with individual constants in the equations, the behavior of persons who remain in their job over the whole period or who quit in each period is explained entirely by the constant, the sample drops from that used in Table II to 1,232 cases, consisting of 877 cases of a single quit, 276 cases of two quits, 67 cases of three quits, and 12 cases of four quits.[15]

15. The differential effect of the individual constants on the union and wage coefficients may reflect the fact that wages are more person-related than unionism, which is much more of a social phenomenon.

TABLE V

FIXED EFFECT LOGISTIC MODEL ESTIMATES OF EFFECT OF UNIONISM ON QUITS

| Explanatory variable | Coefficient and standard error in fixed effect logistic model of quits | Numbers of variables |
|---|---|---|
| Unionism | −0.462 (0.151) | |
| Log wages | 0.128 (0.104) | |
| Individual constants | | 1,232 |

*Source*. Same as Tables I and III.

The results of the calculations, given in Table V, yield coefficients on unionism of similar magnitude to those obtained in Table II but show a marked change in the coefficient on log wages, which changes from significant negative to insignificant positive.[16] Since controlling for individual propensities to quit has essentially no effect on the coefficient on unionism, we conclude that the union impact appears to operate by changing the behavior of the same person rather than by unionization of innately more stable persons. In an organized work place a given individual is less likely to quit than in a nonorganized work place, wages held fixed.

## IV. THE UNION VOICE INTERPRETATION

The analysis thus far has documented the existence of a significant inverse relation between trade unionism and various measures of exit behavior, which exists separate from the effect of unions on wages and the selectivity of innately more stable workers by unions. Because of the lack of a direct measure of "voice" components of unionism, however, the voice interpretation of the relation rests on the inability of other factors to explain the union effect, rather than

16. The highly satisfied group in the PSID consist of persons who responded "very or mostly enjoyable" ($n$ = 2,566); the moderately satisfied group were those who responded "somewhat enjoyable" ($n$ = 868); the rest were labeled dissatisfied ($n$ = 293). In the NLS older men, highly satisfied persons answered "like it very much" ($n$ = 899); moderately satisfied said "like it fairly well" ($n$ = 734); while the rest were dissatisfied ($n$ = 102). Numbers in parentheses are the number of respondents in those categories.

TABLE VI

EFFECT OF UNIONISM ON QUITS OF WORKERS BY LEVEL OF SATISFACTION

| State | Frequency of quits | | | | Logit of quit rates | | | |
|---|---|---|---|---|---|---|---|---|
| | PSID | | NLS older male | | PSID | | NLS older male | |
| | Union | Nonunion | Union | Nonunion | Union | Nonunion | Union | Nonunion |
| Highly satisfied | 4.8% | 9.1% | 0.8% | 5.5% | −2.79 | −2.30 | −2.82 | −2.84 |
| Moderately satisfied | 6.1 | 12.3 | 0.9 | 7.4 | −2.73 | −1.96 | −4.70 | −2.53 |
| Dissatisfied | 10.4 | 18.5 | 2.2 | 24.3 | −2.15 | −1.48 | −3.79 | −1.14 |
| Change from highly satisfied to dissatisfied | 5.6 | 9.4 | 1.4 | 18.8 | 0.64 | 0.82 | 1.03 | 1.70 |

*Source*. Michigan Panel Survey of Income Dynamics.

on positive support for the hypothesis. This section considers some direct evidence on the link between one of the major components of union voice, the grievance and arbitration system, and exit behavior, and finds some support for the hypothesis.

One direct test of the voice interpretation is to compare the effect of unionism on the exit behavior of workers more or less likely to use the grievance system. Persons with grievances should evince a sharper drop in the propensity to exit under unionism than persons relatively pleased with their job. Empirically, in terms of the information on the longitudinal files under study, workers who report themselves dissatisfied with their jobs are most likely to raise grievances and thus be affected by the grievance procedure. Does unionism reduce the rate of exit of dissatisfied workers more sharply than that of satisfied workers?

To answer this question, the quit rates of workers with varying *expressed* levels of job satisfaction were tabulated for union and nonunion workers separately from the Michigan and NLS older male surveys. In the 1972 Michigan survey, workers were asked, "in general, would you say that your job is very enjoyable, mostly enjoyable, . . . not enjoyable at all?" with five possible answers. In the NLS a similar question was asked ("How do you feel about your job?") with four categories of responses. For ease of presentation and to obtain a reasonably sized sample of dissatisfied workers (few express extreme dissatisfaction), the categories have been grouped into three classes: highly satisfied workers, moderately satisfied workers, dissatisfied workers; with the result shown in Table VI.

Measured by *absolute* differences in quit rates, unionism clearly has a greater effect on dissatisfied than on other workers. In the PSID, quit rates are 4.8 points lower for "highly satisfied" union workers than for highly satisfied nonunion workers and 6.2 points lower among the "moderately satisfied," and 8.1 points lower among the dissatisfied. In the NLS older male sample, the greater impact of trade unionism is even more marked, with a 4.7 point difference in quit rates among the satisfied rising to differences of 6.5 and 22.1 points with increased dissatisfaction. Measured as logits of the rates, the picture is less clear. The logit of the quit rate of union workers increases more slowly than the logit of the quit rate of nonunion workers as dissatisfaction increases from highly satisfied persons to dissatisfied persons and from highly satisfied persons to moderately satisfied persons. However, the logits increase more rapidly for union workers in comparisons of moderately satisfied and dissatisfied persons. Interpre-

TABLE VII

EFFECT OF GRIEVANCE CLAUSES ON TENURE OF UNION WORKERS

| Survey (sample size) | Regression coefficient and standard error for effect of percentage of contracts with nonrestrictive grievance clauses on tenure of union workers[a] |
|---|---|
| NLS Older Male ($n$ = 728) | 8.4 (2.2) |
| PSID ($n$ = 801) | 2.2 (1.6) |

a. In the NLS other variables included six occupation dummies, three region dummies, education, age, a retirement benefit dummy, and ln wage, dummies for SMSA and race, number of dependents and the percentage of contracts with layoff and seniority clauses. The $R^2$ was 0.146. In the PSID controls were the same except that they include seven occupation dummies, and a sex dummy and excluded the retirements benefit variable. The $R^2$ was 0.400.

tation of the evidence depends on the metric and group used to evaluate the differences.

A second way of evaluating the effect of grievance and arbitration on exit behavior is to compare the exit behavior of unionized workers in sectors of the economy having different types of grievance systems. The more inclusive or stronger the system, the greater should be the reduction in exit. While detailed knowledge of the operation of grievance machinery is needed for a definitive evaluation, readily available data on collective bargining clauses from the Bureau of Labor Statistics can be used to obtain a rough measure of the scope of grievance systems. The B.L.S. divides grievance clauses into those with "unrestricted" coverage (43 percent of the total), defined as expressing or implying "that any dispute or complaint could be processed as a grievance" and those with "restrictive" coverage (57 percent) because "they limit the grievance process to disputes arising under or relating to specific terms of the contract" [B.L.S. Bulletin 1425-1, p. 6]. The percentage of workers covered by unrestricted clauses varies sufficiently across industries to provide some indication of industrial differences in the scope of grievance procedures. Accordingly, the percentage of contracts with nonrestrictive grievance clauses reported by the B.L.S. [Bulletin 1425, table 1, p. 2] for thirty-one separate industries was added to the NLS and PSID data tapes and the effect of the new variable on the tenure of union workers estimated by linear regression analysis. To eliminate the possibility that "favorable" results would result simply from the distinctive work arrangements in construction (low tenure, limited grievance system), construction workers were deleted from the sample. To reduce the danger that the effect of other union work conditions correlated with nonrestrictive grievance systems might underlie any statistical results, two other measures of contracts were also entered in the calculations: the percentage of contracts with explicit layoff provisions; and the percentage with seniority provisions.[17] All of the various control variables used earlier, except for industry dummies, were also entered into the equation. The resultant coefficients on the grievance clause variable are given in Table VII. While the crude measure of differences in grievance procedures makes the results no more than suggestive, the positive impact of the scope of the grievance system on tenure is consistent with the model.

More direct evidence on the impact of a grievance system on turnover is given, for a limited number of hospitals, in a study by Sargent and Clawson [1974]. In their data set, hospitals with a written grievance procedure had a yearly separation rate of 0.50 while hospitals without a grievance system had a separation rate of 0.81; the simple correlation of separation rates with the presence of written grievance procedures is −0.72; the partial correlation between separation and grievance procedures is −0.74, conditional on bed size of hospital and distance from the center of the nearest city.[18] While absence of other control variables makes these relations suggestive rather than definitive, the significant relation between turnover and grievance procedure supports our interpretation of the evidence on the data tapes for individuals.

Finally, the voice interpretation of the apparently sizeable nonwage effect of unionism on exit, while novel in some respects, is consistent with the corpus of industrial relations analyses and findings. Industrial relations studies of unionism have traditionally stressed the importance of grievance, arbitration, and related factors in the impact of the institution [Slichter, Livernash, and Healey, 1975; Dunlop, 1970; and Kerr, 1978]. Analyses of why workers join unions almost invariably find that worker grievances, usually with specific managerial policies rather than possible higher wages, motivate organization [Seidman, Landon, and Karsh, 1957]. Companies

17. The percentage with layoff clauses was obtained from U. S. B.L.S. Bulletin 1425–14, table 1, p. 32. The percentage with seniority clauses was taken from U. S. B.L.S. Bulletin 1425–13, table 8, p. 53.

18. The figures reported are based on simple correlation coefficients reported in the Sargent-Clawson [1974] article.

that face organizing drives are usually advised by labor-management legal consultants to concentrate on personnel policy issues relating to what could easily be labeled voice issues. At the other end of the spectrum, AFL-CIO propaganda places great stress on the role of unions as the "voice" of workers. These pieces of information suggest that the interpretation of the depressant effect of unionism on exit given in this study is broadly consistent with the industrial relations picture of the operation of trade unions in the United States.

## V. Consequences and Implications

The evidence that trade unionism appears to reduce the exit propensity of workers, with at least some of the effect due to the operation of unions as an institution of collective voice, has implications for the functioning of the job market and for future research on trade unions. It suggests that unionism is a major force in the creation of a relatively permanent enterprise work force and thus of the types of market arrangements and adjustments to which permanent attachment gives rise (see Feldstein [1976] and the literature cited therein). For example, with quits playing a lesser role in labor force adjustments in the union sector, one would expect organized firms to make greater use of layoffs and recalls, as appears to be the case [Medoff, 1979]. At the same time, permanent attachment of workers can be expected to reduce wage adjustments over the business cycle and enhance employment fluctuations [Feldstein, 1976]. The composition of the compensation package is also likely to be affected, with the greater likelihood of remaining with a firm raising the deferred compensation share of the wage bill [Freeman, 1978a]. Investments in firm-specific human capital may also be increased by union-induced reductions in exit. In terms of productivity, the increase in job tenure and reduction quits can be expected to raise the efficiency of organized establishments by lowering the costs of turnover in the form of hiring and training expenses. Brown and Medoff [1978] find that a sizeable proportion of the relation between unions and productivity among two-digit manufacturing industries across states is, indeed, due to the lower quit rates of unionized workers.

Finally, with respect to research, this study has taken only a first empirical step in analyzing the nonmonopoly wage effects of unionism. The model and findings suggest that greater attention be given to the economic effects of unionism beyond the monopoly wage gains that are at the center of most modern empirical work.

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