JOB SATISFACTION, WAGES, AND UNIONS*

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ABSTRACT

This paper provides a systematic empirical analysis of the effect of union membership on job satisfaction and wages, and shows how the interaction between these effects leads to empirically observable relations between unionization and individual quit probabilities. Using the National Longitudinal Survey of Mature Men, several empirical results were obtained. First, union members, on average, report lower levels of job satisfaction. Interestingly, unionization causes greater dissatisfaction at higher tenure levels. These findings are attributed to both the politicization of the unionized labor force and the fact that union members face flatter earnings profiles. The importance of the latter effect is reflected by the empirical fact that unions have a strong negative effect on quit probabilities at low levels of tenure, but the effect diminishes (absolutely) as tenure increases.

I. INTRODUCTION

The analysis of the determinants of individual earnings has received intensive study in the past decade.[1] Much of this interest can be traced directly to the development of a theoretical framework that relates the individual's earnings profile to his human capital stock, the skills and abilities that are rented in the labor market.[2] It has been seen that this

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1 The most notable empirical analysis is Mincer's [24] study of white male earnings. Other studies were recently surveyed by Blaug [5].

2 The most important presentations of the human capital model can be found in Becker [3] and Ben-Porath [4].

approach provides a unified method of analysis that can explain the determinants of the market wage structure. These studies, however, have largely concentrated on the analysis of money earnings. Little systematic interest has been paid by economists to the determinants of full wages, where the full wage is defined as the sum of money earnings and any nonpecuniary job components. Recent studies by Duncan [8] and Hamermesh [15] have shown that working conditions and "job satisfaction" are systematically related to basic individual characteristics, and that these relationships lend themselves to intuitive economic explanations.

The purpose of this paper is to provide a systematic empirical analysis of the effect of union membership on job satisfaction and wages, and to show how the interaction between these effects leads to empirically observable relations between unionism and individual quit probabilities. The estimation of the union effect on money wages has long been a major topic of importance in labor economics.[3] This paper will document the fact that unionism has a strong *negative* effect on job satisfaction. Many hypotheses can be set forth to explain this finding. For example, a union "voice" effect may make workers more aware of what is wrong with the firm (Freeman [11]); or an argument that unionized jobs are inherently unpleasant (which may be why a union was created in the first place) and hence the union wage effect can be viewed as a compensating differential (Duncan and Stafford [9]). This paper reports empirical tests of these hypotheses and also presents evidence of an additional factor that helps to explain certain similarities in the effect of unions on satisfaction and quit rates, namely, that unions tend to flatten the earnings profile within the job.[4]

Section II of the paper contains a brief discussion of the framework used in this study. The basic empirical results, based on the National Longitudinal Survey of Mature Men, are presented in Section III and show the relationship between job satisfaction and unionism. Section IV relates the union effect on wages to the union effect on satisfaction and investigates the outcome of these processes on turnover behavior of union members. Section V briefly replicates the analysis using monetary measures of job satisfaction. Finally, Section VI summarizes the major empirical findings of the study.

II. A FRAMEWORK FOR THE ANALYSIS

Define the full wage, Y, as the sum of the money wage rate, W, and any nonpecuniary job-consumption components, V, measured in monetary terms:

3 See, for example, Lewis [22], Ashenfelter [1], and Boskin [7].

4 Perhaps it is important to point out that these alternative hypotheses are not mutually exclusive. They may *all* be working jointly in the labor market.

$$(1) \qquad Y = W + V$$

It should be obvious that V (and hence the full wage) depends on the nature of the individual's utility function. Thus, there exists a distribution of V and Y across individuals for the same job. In order to simplify the discussion that follows, I assume that tastes do not differ across individuals so that any particular job is characterized by a unique value of the full wage.[5]

Job satisfaction will be defined as a monotonic transformation of the full wage for individuals with a vector of characteristics Z. That is, $S = S(Y,Z)$, where $\partial S/\partial Y > 0$. The variables in the vector Z may be composed of variables that only affect measured job satisfaction without any effect on the full wage, or they may be composed of variables that both affect the full wage and have a *direct* effect on measured job satisfaction. Clearly, in a perfectly competitive market characterized by perfect information and costless mobility, the full wage would be constant across jobs, and under the assumption of homogeneous tastes and individuals (i.e., for a given Z), the distribution of S would be degenerate: the level of job satisfaction would be constant across individuals. Hence, in order to explain any differences in job satisfaction, we must resort to factors such as imperfect information and costly mobility, and the existence of noncompeting forces such as unionism.[6]

For example, uncertainty about the job exists since the nonpecuinary characteristics of the job are not immediately known. That is, after an individual and firm begin the job contract, both parties undergo a process of learning about each other in order to determine whether the match is a proper one.[7] If the match is found to be imperfect, both parties have an incentive to initiate a job separation; the firm finds that the individual's marginal product is not as high as expected, and the individual learns that the firm's full wage is lower than expected. In a life-cycle context, the individual "tries out" several firms until the proper match is found. Thus, the analysis implies that older individuals are more likely to be satisfied with their jobs since they have been sampling the job market for a longer period of time and are more likely to have drawn a successful match. The relationship between satisfaction and job tenure is less clear. It is not obvious why there should be any relationship between the two variables for a given individual once the matching period has elapsed. On the other hand,

5 The assumption of homogeneous tastes has been fruitfully applied to such problems as addiction, fads, and advertising by Stigler and Becker [25].

6 The reader should note that the assumptions made in this paragraph about the meaning of job satisfaction are far from innocuous. Measured job satisfaction is usually derived from answers to questions asking individuals how they feel about their jobs. The nature of the answers given has been debated heatedly in the psychological literature. An excellent statement of the advantages and drawbacks of using satisfaction variables in economics is given by Freeman [13].

7 See Jovanovic [20] for a theoretical model of the job-matching mechanism.

in a cross-section we observe individuals in both the pre- and postmatching periods. If job tenure is correlated with some measure of matching, then we may observe a positive relationship between satisfaction and tenure. Finally, individual traits that increase efficiency in job search (e.g., education) increase job satisfaction since they tend to increase the probability of a successful job match,

In analyzing the effects of unions on job satisfaction, it is important to note that since job satisfaction is a positive function of W and V, the union wage effect has a positive effect on job satisfaction, holding V constant. Thus, I confine the discussion initially to the effect of unions on nonpecuniary components, V. The simplest way to isolate this relationship (given the fact that nonpecuniary job rewards are unobserved) is by estimating the effect of unionism on S, holding the money wage constant. Thus, the model can be written as:

$$S = \alpha_0 + \alpha_1 \mathrm{U} + \alpha_2 \mathrm{W} + \alpha_3 \mathrm{Z} \tag{2}$$

where U is a dummy variable equaling unity if the individual is a union member and zero otherwise.[8]

Several hypotheses exist that predict that α_1 in equation (2) will be nonzero. The simplest way is to postulate the existence of a competitive market in which the firm has different methods of offering payments to workers, W and V. Suppose the union succeeds in increasing the firm's money wage rate. If the firm is faced with such an increase, and if the union did *not* have any monopoly power, then clearly the firm would simply compensate the workers' high wage rate with a corresponding decrease in nonpecuniary job rewards; thus $\alpha_1 < 0$. The higher the monopoly power of the union, the less able the firm would be in taking away V from its labor force and the less likely the negative "compensation" would occur.[9]

An alternative hypothesis is that workers in very unpleasant jobs have the greatest incentive to organize. Thus, bad jobs "cause" both dissatisfaction and union membership. Hence, equation (2) can be viewed as part of a

8 The simplest way to motivate equation (2) is the following: Suppose p gives the probability of being satisfied with the job and follows a logistic cumulative distribution. Thus $p = [1 + \exp\{-(\gamma_1 Y + \gamma_2 Z)\}]^{-1}$. Suppose further that unionization is the only variable to affect V (this is an innocuous assumption made only to simplify the notation since the multivariate extension is straightforward) so that $V = \beta_0 + \beta_1 U$. It can be shown that:

$$\ln [p/(1-p)] = \gamma_1 \beta \ + \gamma_1 W + \gamma_1 \beta_1 U + \gamma_2 Z$$

which is equation (2) in the text with $S = \ln [p/(1-p)]$. Thus a simple logit regression of the probability of being satisfied on wages and unions (and Z) yields an estimate of the union effect on V as long as unions do *not* affect satisfaction directly (do not enter the vector Z). However, as will be seen below, this is a very real possibility due to the "exit-voice" effect of unions.

9 In fact, if unions have a very strong monopoly power so as to demand not only higher W, but also higher V, then the union coefficient would turn positive.

simultaneous equation system, and since the union variable and the disturbance in (2) are correlated, the use of simple least squares yields biased estimates of α_1. This may lead to the researcher's concluding that unions affect V when in fact no such causal relationship exists.[10]

A third hypothesis that also leads to a negative relationship between job satisfaction and unionization was proposed by Freeman [11].[11] Using Hirshman's [18] theory of the "exit-voice" tradeoff, Freeman argues that one important effect of unionization is to provide an efficient mechanism that gives "voice" to the workers while reducing exit (i.e., quits). In other words, the exit-voice hypothesis argues that in order for the workers' voice to be heard effectively, it is important for the union to make them aware of what is wrong with their jobs. Thus, a by-product of unionization is the politicization of the firm's work force, and union members can be expected to express less job satisfaction than nonunion workers. That is, the exit-voice model states that in order for firms to hear the workers effectively, the firm's work force must express itself "loudly." Note, however, that this dissatisfaction is not genuine in the sense that it leads to quits, but is instead a device through which the union can tell the firm that its workers are unhappy and are demanding more. In terms of equation (2), the estimated coefficient of U may not only reflect a union effect on V, but will also include a *direct* effect of union on satisfaction, since union membership enters the vector Z.

The empirical work in the next sections tries to develop tests that will distinguish the validity of these hypotheses. It will be seen that the key identification device is the fact that unions do not decrease satisfaction for the entire work force in the firm. This will allow us to distinguish between the hypotheses as well as illustrate the relationship between the union effect on wages and job satisfaction.

III. THE DETERMINANTS OF JOB SATISFACTION

The empirical analysis is carried out on the National Longitudinal Survey of Mature Men.[12] The data analyzed in this paper are taken from the 1971 survey at which point the men are 50–64 years of age. To simplify the analysis, I restricted the sample to white working men who reported the key variables needed for the study. These restrictions reduced the working sample to 1873 observations.

Clearly, the success of the empirical work will depend to a large extent on the measure of job satisfaction used. In this section of the paper, I deal

10 The validity of this hypothesis will be tested empirically in Section III.

11 A more detailed discussion of the exit-voice model is in Freeman and Medoff [14].

12 See [27] for a description of the data base.

only with "quantal" measures, such as indexes and dummy variables.[13] These variables are created from the answer to the following question: "How do you feel about the job you have now? Do you like it very much, like it fairly well, dislike it somewhat, or dislike it very much?" From the response, it is possible to create many alternative measures of job satisfaction.[14] I use three different measures of satisfaction:

1. *INDEX*. This variable is created by assigning it a value of 1 if the individual dislikes his job very much; 2 if he dislikes it somewhat; 3 if he likes it fairly well, and 4 is he likes his job very much. This breakdown, of course, utilizes all the information on the job-satisfaction question.

2. *LOT*. This dummy variable is constructed by assigning it a value of unity if the individual likes his job very much, and zero otherwise.

3. *SATIS*. This dummy variable takes on the value of unity if the individual likes his job very much or fairly well, and zero otherwise.

The means of the dependent variables indicate that in this age range, men are highly satisfied with their jobs. For instance, 91.4 percent of the sample reports liking their jobs very much or fairly well, and 44 percent of the sample reports liking their jobs very much. Thus, the widespread popularity of job "alienation" as a prevalent characteristic of work is not substantiated by the data.[15]

Table 1 presents the regressions estimating equation (2) for the alternative dependent variables, using least squares for *INDEX* and maximum likelihood logit for *LOT* and *SATIS*. The coefficients reported for the logit regressions are marginal coefficients showing changes in probability due to unit changes in the independent variables, and are thus comparable (in units) to OLS coefficients.[16] Table 1 regresses job satisfaction on the wage rate, union membership, and a vector of personal and job characteristics. The job characteristics include job tenure and occupation dummies at the one-digit level. These are included since satisfaction (or nonpecuniary components of the job) may vary systematically with the type of work performed. The personal characteristics vector includes education, labor force experience, health, marital status, wife's education, number of dependent children, and family income. These family variables are introduced since they may affect the individual's "demand" for money wages,

13 McFadden [23] has a lengthy discussion of the economic problems and behavioral assumptions required when estimating equations with quantal dependent variables.

14 Actually, there exists more satisfaction data in the NLS since the men were also asked to mention three things they liked and disliked about the job. For an example of an index using this information, see Kalachek and Raines [21].

15 See [26] for an exposition of this view.

16 More precisely, if $p = 1/(1 + e^{-bx})$, then the tables report $dp/dx = bp(1 - p)$, where the mean probability is used in the calculation. The t-ratios refer to the logit coefficient b.

TABLE 1
JOB-SATISFACTION REGRESSIONS[a]

Variable	*INDEX* Coeff.	*INDEX* t-ratio	*LOT* Coeff.	*LOT* t-ratio	*SATIS* Coeff.	*SATIS* t-ratio
U	−.1046	(−2.97)	−.0865	(−3.10)	−.0297	(−2.02)
W	.0244	(3.34)	.0269	(4.03)	.0052	(1.29)
MARRIED	−.0397	(−.56)	−.0423	(−.81)	.0037	(.14)
WIFED	.0088	(1.48)	.0050	(1.20)	.0025	(1.16)
EXPER	.0084	(2.10)	.0123	(1.79)	.0039	(2.23)
EDUC	.0068	(.89)	.0067	(1.12)	.0013	(.42)
CHILD	−.0402	(−2.63)	−.0285	(−2.25)	−.0133	(−2.46)
HEALTH	−.0900	(−2.37)	−.0421	(−1.40)	−.0354	(−2.38)
INCOME	.0022	(1.05)	.0136	(.83)	.0007	(.71)
TENURE	−.0014	(−1.01)	−.0023	(−2.11)	.0006	(.98)
R^2	.076					
Log likelihood			−1200.0		−526.7	

a Key to variables: *MARRIED* = 1 if individual is married, spouse present; *WIFED* = wife's education; *EXPER* = years of labor force experience; *EDUC* = years of education; *CHILD* = number of children living in the household; *HEALTH* = 1 if health limits work; *INCOME* = family income; *TENURE* = current job tenure. Also included in the regressions were occupation dummies at the one-digit level.

thus changing the optimal mix of money wages and nonpecuniary components.[17]

From Table 1, it can be seen that, as expected, unionization has a strong negative effect on job satisfaction.[18] The coefficient is more significant when using *INDEX* and *LOT* since these variables are better able to distinguish between satisfied and dissatisfied workers. Note that not only is the

17 In fact, there was considerable experimentation concerning the independent variables in Table 1. The most obvious set of omitted variables are industry dummies. In fact, the introduction of 11 one-digit industry dummies results in an insignificant increase in R^2 (the F statistic in the *INDEX* regression was .72), and in little change in the union coefficient. Moreover, it seems reasonable to suppose, as a first-order approximation, that it is the *type* of work (i.e., the occupation) which is likely to affect job satisfaction, and in fact the occupation dummies are often sizable and significant.

18 It is worthwhile to note that despite the expense in getting logit estimates, the coefficients were almost indistinguishable from the OLS coefficients. For example, the OLS coefficient in column 2 was −.0782 ($t = -3.05$), and in column 3 it was −.0287 ($t = -1.92$). Due to this similarity, the linear probability model is used in all other regressions except where it is explicitly noted that an alternative statistical technique was employed.

statistical effect of U strong, but its numerical effect is substantial.[19] For example, in the *LOT* regression the effect of a one dollar increase in the wage rate is to increase the probability by .027 units, while being a member of a union decreases the probability by .087 units. The union effect is, therefore, equivalent to a wage cut of \$3.22. This result alone indicates that the union effect on satisfaction is much more than just the union effect on nonpecuniary job components: since the average wage in the sample is \$4.76, the union effect on satisfaction is equivalent to a 68 percent wage cut! That is, if this sizable effect were entirely due to a *real* reduction in nonpecuniary job components, one would have difficulty explaining the existence of unions. Therefore, there must be other factors at work.[20]

As was discussed earlier, two alternative hypotheses were that unpleasantness created unions and that union membership reduced exit while increasing the voice of the members. I now discuss each of these hypotheses in turn. First, it is easy to think of an indirect test for reverse causation. In particular, if unions have monopoly power over the full wage, and if it is bad working conditions that led to the creation of the union, then over time the factors underlying the dissatisfaction should disappear. Hence, we would expect that the older, more established unions (e.g., craft unions) would be less likely to have a negative correlation with satisfaction than more recent unions (e.g., industrial unions or, better yet, government unions). Table 2 presents the union coefficients from regressions similar to those presented in Table 1, by type of union. Since, as was noted in footnote 18, logit and OLS coefficients were very similar, the linear probability model was used in calculating the type of union effects. As can be seen, under any definition of job satisfaction, the statistical and numerical differences between the effects of craft and industrial unions are nil. For example, when using *INDEX*, being a member of a craft union lowers satisfaction by .133 units, while being a member of an industrial union lowers it by .137 units. Moreover, the effect of being in a government union is always weakest (statistically), contradicting the expectation that due to their recent formation, individuals in these unions are still likely to "suffer" from the unpleasant job conditions that led to union formation.

Although these results provide some insight into the underlying process, they are not conclusive. In particular, it may be that sectors that were organized long ago were considerably "worse" and that progress has

19 Although the main focus of this paper is the union coefficient, it is instructive to note the effects of the other variables. The wage rate is, of course, positive and significant. Both experience and education are generally positive as the discussion in Section II suggested, while the number of children and bad health have negative effects on satisfaction. The effect of job tenure is quite sensitive to the dependent variable used.

20 In fact, a simpler way of showing this is to note that the simple correlation between unionization and *INDEX, LOT,* and *SATIS* is −.142, −.155, and −.061, respectively.

TABLE 2
UNION EFFECT BY TYPE OF UNION

Type of Union	*INDEX*	*LOT*	*SATIS*
Industrial	−.1366	−.1119	−.0270
	(−2.95)	(−3.33)	(−1.37)
Craft	−.1328	−.1044	−.0271
	(−2.56)	(−2.69)	(−1.23)
Government	−.1141	−.0696	−.0384
	(−1.31)	(−1.10)	(−1.04)
Other	−.0093	−.0027	−.0191
	(−.15)	(−.06)	(−.70)

TABLE 3
UNION COEFFICIENTS IN SIMULTANEOUS SYSTEM

Dependent Variable	2SLS	OLS
INDEX	−.3306	−.1046
	(−2.78)	(−2.97)
LOT	−.1746	−.0782
	(−2.04)	(−3.05)
SATIS	−.1344	−.0287
	(−2.66)	(−1.92)

been slow. A more direct way of tackling the problem is to note that the hypothesis generates a simultaneous equation system in which union membership and satisfaction are jointly determined. By obtaining an instrument for union membership uncorrelated with the disturbance in the satisfaction equation, we can estimate the unbiased effect of unionization on satisfaction. A structural union membership equation of the form $U = U$ (satisfaction, education, experience, geographical region, urbanization, industrial dummies) is hypothesized to exist. Clearly, geographical region and industry enter this equation since unions have tended to organize in certain areas of the country and in particular industries. The resulting 2SLS coefficients of union membership in the job-satisfaction equation are presented alongside the biased OLS coefficients in Table 3. It is seen that the use of a predicted union membership variable does *not* take away the statistical significance of the union coefficient in the satisfaction equation, and if anything increases (in absolute terms) its numerical magnitude.[21] Thus we can conclude that the finding of a negative relationship between job satisfaction and unionization is not due solely to the fact that bad jobs create unions, but is partly due to the direct effects of union membership on nonpecuniary job rewards and/or job satisfaction itself.[22]

An alternative explanation for the negative effect of unionization on job satisfaction is the exit-voice tradeoff discussed in Section II. That is, to form an effective organization, unions focus attention on the working conditions of the firm, making workers more aware of what is wrong with their jobs and leading to unionized workers "expressing" more dissatisfaction. At this level, the results discussed are entirely consistent with this hypothesis. Thus, in order to test whether it is indeed the politicization of workers that leads to lower expressed levels of satisfaction, some additional structure is needed. In particular, the voice mechanism reflects the demands of some average of workers' desires. In the exit mechanism, on the other hand, it is likely to be the low-tenure workers who have a larger incentive to quit, and hence it is this group which conveys information to the firm. In other words, the voice mechanism gives greater weight to the voice of older, high-tenure individuals. Thus, in order to be effective, the voice mechanism (i.e., the union) must be successful at increasing discontent among high-tenure men.

This hypothesis, in effect, postulates an equation of the form:

$$S = \alpha_0 + \alpha_1 U + \alpha_2 t + \alpha_3 U \cdot t + \alpha_4 W + \alpha_5 Z \tag{2'}$$

where t is job tenure. The interaction term $(U \cdot t)$ is predicted to be negative. The estimated coefficients for (2′) are presented in the top panel of Table 4. As can be seen, once the interaction term is introduced into the equation, the coefficient of U becomes insignificant, and in the regression on *INDEX*

21 It is interesting to note that the effect of job satisfaction on unions is strongly negative, which indicates that indeed bad working conditions serve to create unions. The complete set of regressions is available from the author on request. Note also that the linear probability model was used in the estimation of the union equation and of the dichotomous satisfaction equations. This procedure leads to consistent estimates of all the parameters. To increase efficiency, maximum likelihood methods described by Heckman [17] can be applied.

22 It is important to point out that the estimation of the job-satisfaction–union-membership structural equation system made rather strong assumptions concerning the set of variables included in each of the equations. In particular, the main identification device was to include industry only in the union-membership equation and occupation only in the job-satisfaction equation. As I argued earlier, this may be a valid first-order approximation since unions have organized along industrial lines and since it is the type of work that mainly affects job satisfaction. Moreover, the reader should also note that a complete analysis of this simultaneous system should incorporate the role of wages (both expected and actual) in determining the incentives for union membership as well as the extent of job satisfaction.

TABLE 4
UNION INTERACTION EFFECTS

Dependent Variable	U	t	$U \cdot t$
	A. Holding the Wage Constant		
INDEX	−.0437	.0022	−.0040
	(−.80)	(1.09)	(−1.66)
LOT	−.0093	−.0001	−.0045
	(−.23)	(−.11)	(−2.28)
SATIS	−.0303	.0005	.0001
	(−1.31)	(.64)	(.09)
	B. Not Holding the Wage Constant		
INDEX	−.0071	.0011	−.0050
	(−.13)	(.66)	(−1.84)
LOT	.0145	.0006	−.0054
	(.37)	(.50)	(−2.71)
SATIS	−.0195	.0006	−.0001
	(−.84)	(.83)	(−.06)

and *LOT,* the interaction term is negative and significant at the 10 and 5 percent levels of significance, respectively. Therefore, as predicted by the exit-voice hypothesis, the increase in dissatisfaction due to unionism occurs at higher levels of tenure.[23]

It is also important to note that since the union effect on satisfaction is not constant throughout all tenure levels, the results are not consistent with a hypothesis that has the employer in unionized firms handing out smaller compensations in V, nor is it consistent with a view that the results are due solely to the unpleasantness inherent in union jobs. In order to make the results in Table 4 consistent with these views, one would have to argue that somehow the unpleasantness in the job varies systematically with tenure or that only the high-tenure individuals gain from unionization in terms of the money wage. Fortunately, it is easy to obtain some insight into the relationship between the union effects on job satisfaction and on money wages. Panel B of Table 4 reestimates equation (2′) omitting the wage from

23 An alternative way of obtaining this result is by estimating equation (2′) within tenure groups. If this estimation is conducted, I find that (using *INDEX* and *LOT* as dependent variables) unions have no significant effects on satisfaction for tenure levels below 14 years, but have a strong and significant effect beyond 25 years of tenure. For instance, using *LOT,* I find that the effect of U on job satisfaction is $-.017$ ($t = -.33$) for men with less than four years of tenure, but it is $-.158$ ($t = -3.08$) for men who have been at the job longer than 25 years.

the equation. The effect of this exercise on the union and union-tenure interaction coefficients is quite interesting. In particular, note that the omission of the wage reduces the negativity of the variable U on job satisfaction, which is what one would expect as long as unions have a positive effect on money wages. More interestingly, omitting the wage from the satisfaction equation makes the point estimate of the coefficient of the union-tenure interaction *more* negative. For example, using *INDEX* and *LOT*, the coefficient of $U \cdot t$ is increased by about 25 percent absolutely and the statistical significance of the interaction also rises. The fact that the interaction term is more negative when wages are omitted from the regression suggests that, if anything, the high-tenure individuals gain least in terms of money wages from unionization.[24] This result clearly conflicts with the popular opinion that it is the high-tenure men who gain most from unions. The result is studied further in the next section.

IV. SATISFACTION, WAGES, AND TURNOVER

The confirmation of the predictions of the exit-voice hypothesis in the previous section raises interesting new questions. If the effect of unions on satisfaction is so tenure-dependent, and since this result is sensitive to whether or not money wages are held constant, what is exactly the effect of unions on the wage structure within the firm? It has been noted that the structure determining union wages differs from the nonunion structure with respect to such factors as race [1], education and age [6, 19]), etc. This section provides empirical evidence on the difference in structures with respect to job tenure. Moreover, these findings will be useful in obtaining more insight into the union-tenure interaction effects in the job-satisfaction equations presented in Table 4. In particular, the wage-generating equation may be written as:

$$\ln W = \beta_0 + \beta_1 U + \beta_2 t + \beta_3 U \cdot t + \beta_4 X \tag{3}$$

where X is a vector of variables that affect wages.[25]

Table 5 presents estimates of (3) with and without union-tenure

24 It is very important to note that the statistical significance of these effects is not very strong. In particular, the coefficient of U is insignificantly different from zero whether or not the wage rate is included in the regression and the standard errors of the interaction terms are always relatively large. The direction of the effects, however, is suggestive of the underlying union wage structure.

25 The variables that enter the vector X have been the focus of much attention in recent years. In this study, I included education, labor force experience, marital status, health, urbanization, region of residence, and industrial dummies. I do not enter occupation dummies since in the human capital approach, occupation, like earnings, is the outcome of the investment process.

TABLE 5
EARNINGS FUNCTIONS[a]

Variable	Coeff.	*t*-ratio	Coeff.	*t*-ratio
U	.0485	(2.02)	.2307	(6.17)
t	.0092	(9.53)	.0136	(11.51)
$U \cdot t$	——	——	−.0118	(−6.32)
MARRIED	.1622	(4.50)	.1667	(4.68)
EXPER	−.0044	(−1.67)	−.0043	(−1.64)
EDUC	.0551	(11.93)	.0534	(11.65)
HEALTH	−.1162	(−4.45)	−.1089	(−4.20)
SMSA	.1475	(6.46)	.1411	(6.24)
EAST	−.0719	(−2.17)	−.0707	(−2.16)
MIDWEST	−.0874	(−2.64)	−.0884	(−2.70)
SOUTH	−.1673	(−4.82)	−.1676	(−4.88)
R^2	.341		.355	

a Key to additional variables: *SMSA* = 1 if individual lives in SMSA; *EAST, MIDWEST, SOUTH* = 1 if individual lives in respective region. These regressions also hold constant a set of industry dummies at the one-digit level.

interactions. Ignoring the interaction terms gives a union wage effect of approximately 4.9 percent. Bringing in the interaction yields:

$$\partial \ln W/\partial U = .231 - .012T \tag{4}$$

and raises R^2 from .341 to .355. Therefore, unionization significantly *reduces* the rate of growth of earnings within the firm. Consequently, the wage effect of unions diminishes with tenure and actually becomes negative past 19.3 years.[26]

This finding can be used to explain the rather peculiar results presented in Table 4 where the union-tenure effect on satisfaction becomes more negative if wages are *not* held constant.[27] In particular, suppose that job

26 It could be argued that if the earnings profile within the job is concave, and since unions increase tenure, the union-tenure interaction may be proxying for tenure squared. If this variable is introduced into the regressions in Table 5, the partial effect of unions on the logarithm of wages is unaffected. Moreover, running earnings functions within the two groups (union and nonunion) yields the same result, namely, that there exists a negative correlation between unionization and wage growth in the job.

27 Although an explanation of why the union wage effect declines with tenure is beyond the purview of this study, it is interesting to note that it is consistent with a hypothesis that has unions as wanting to reduce wage inequality in the firm (see Freeman [12]). Alternatively, it may simply be that unionized jobs tend to be in industries that have flatter earnings profiles. It can even be interpreted as the result of selectivity bias. That is, it may be that in this age range the best long-tenure workers advance into nonunion supervisory positions. Thus, the interaction terms in both Tables 4 and 5 could be interpreted as indicating that

satisfaction depends in part on the worker's perception of his relative standing in the income distribution. Clearly, at low levels of tenure the union is extremely effective in raising the relative wage of unionized workers, and satisfaction should not, at the very least, be lower in unionized firms. As job tenure increases, however, the relative advantage of being in a union diminishes, lowering the individual's relative standing in the earnings distribution and hence creating an increase in dissatisfaction. Thus when wages are not held constant in the satisfaction regression, the interaction term $U{\cdot}t$ is capturing both the exit-voice effect and the relative wage effect and will be more negative than when wages are held constant. The fact that inclusion of the wage rate in the satisfaction equation in Table 4 does not drive the interaction coefficient down to zero is, of course, an indication of the importance of the exit-voice hypothesis. Conversely, the fact that omitting the wage from the satisfaction equation makes the interaction coefficient more negative highlights the fact that, ceteris paribus, high-tenure unionized workers are more likely to be dissatisfied with their jobs than low-tenure unionized workers due to their falling relative wage.

The importance of distinguishing between these two effects lies in their ability to explain the *observed* behavior of workers—namely, quits. It is well known that unions diminish quit rates. The question that arises from the empirical results presented in this paper is whether this effect holds for all levels of tenure. It is in explaining this relationship that the two hypotheses differ in their predictions. Although the exit-voice hypothesis predicts more dissatisfaction at higher levels of tenure, this dissatisfaction is not genuine. That is, even though workers express their grievances openly, the quit rate is in fact reduced, and thus there would be no reasons to expect high-tenure workers (who express more dissatisfaction) to quit more often. On the other hand, suppose that the increase in dissatisfaction with tenure is *partly* due to the fact that the union relative wage has fallen; thus opportunities elsewhere have improved, and we would expect that the union effect on quit rates should diminish (absolutely) with tenure and may, in fact, be positive for high levels of tenure. In other words, the importance of the falling union relative wage can be measured by the behavior of the quit rate in unionized firms as tenure increases. Table 6 presents quit-rate regressions using the linear probability model. The dependent variable is the probability of quitting the job between 1969 and 1971. It should be pointed out that since the sample is restricted to individuals who are in the labor force as of 1971, the quit was not into retirement. The quit probability is regressed on a set of variables measured as of 1969. The coefficients of interest for this study are

the average "quality" of a unionized worker falls with tenure. In any case, it should be pointed out that the results in Table 5 are not entirely consistent with those of Johnson and Youmans [19], but are in line with the more recent findings of Bloch and Kuskin [6].

TABLE 6
UNION EFFECTS ON QUIT PROBABILITIES[a]

Regression	U	t	$U \cdot t$
(1)	−.0486	−.0028	—
	(−4.22)	(−5.83)	
(2)	−.1034	−.0043	.0040
	(−6.01)	(−7.25)	(4.27)

a The following variables were held constant in the regression: education, labor force experience, marital status, wife's education, number of children, family income, health status, and industry dummies at the one-digit level.

reported in Table 6, which also contains notes on the standardizing set of variables. It can be seen from Table 6 that the effect of unionization on quits is strongly negative at low tenure levels, but becomes weak and turns positive after 25 years of tenure.[28] This finding suggests that the falling relative wage effect is an important consequence of unionization. Thus, the increase in dissatisfaction with tenure (when wages are not held constant) is not entirely fictitious, but is partly due to the fact that the union wage effect favors long-tenure workers the least.[29]

V. MONETARY MEASURES OF JOB SATISFACTION

The National Longitudinal Survey is particularly useful in obtaining information on monetary measures of satisfaction since individuals were asked to place a monetary value on their jobs. This measure was obtained from answers to the question: "Suppose someone in this area offered you a job in the same line of work you are now in. How much would the new job have to pay for you to be willing to take it?" The answer to this question is, to a large extent, the individual's reservation wage or supply price to other jobs, W^*. Presumably it includes monetary and nonmonetary returns on the current job as well as the expected returns to investments in search. If unionization reduces nonpecuniary job components, then clearly the effect

28 If the wage rate is introduced in Table 6, we still find that unions decrease quits at lower levels of tenure, but that this effect turns around after 20 years. Although the results when the wage is included are weaker, the coefficients are still statistically significant. A discussion of turnover in microdata is found in Bartel and Borjas [2].

29 It is interesting to note that a recent study by Farber and Saks [10] finds additional evidence of the importance of the relative wage effect of unions. In particular, they find that workers with high levels of tenure are less likely to vote affirmatively to form a union than workers with low levels of tenure.

of the union on W^* should be smaller than the effect of the union on the money wage rate.

Unfortunately, only 40 percent of the sample responded to this question with a numerical answer. Instead, individuals often replied that they would not take a new job at any wage (i.e., that their reservation wage was *very* high).[30] Thus confining ourselves to numerical answers truncates the sample and leads to biased estimates of the union effect. Fortunately, a simple statistical technique can be used to obtain consistent estimates. In particular, let I_i be an index measuring the willingness of individual i to accept another job at some wage. Without loss of generality, assume that if $I_i > 0$, then W_i^* is observed; thus the index is positively correlated with the willingness to accept a money wage offer, and could be interpreted as an *inverse* measure of satisfaction. We can write the model as:

$$I_i = X_{1i}\beta_1 + \epsilon_{1i} \tag{5a}$$

$$W_i^* = X_{1i}\beta_2 + \epsilon_{2i} \tag{5b}$$

where W_i^* is observed if $\epsilon_{1i} > -X_{1i}\beta_1$. Heckman [16] has shown that, in general, this sample censoring leads to the disturbance ϵ_2 not having zero mean and depending on the probability that an individual is included in the sample. Thus, the censoring is formally equivalent to an omitted variable problem. In particular, if the disturbances are normally distributed, Heckman has shown that consistent estimates of β_2 can be obtained by estimating:

$$W_i^* = X_{2i}\beta_2 + \gamma\lambda_i + v_i \tag{6}$$

where $\lambda_i = \phi(Z_i)/[1 - \Phi(Z_i)]$, ϕ and Φ are the density function and the cumulative function of the normal distribution, respectively; $Z_i = -X_{1i}\beta_1/\sigma_1$; σ_1 is the standard deviation of ϵ_1; and v_i is a random disturbance. Clearly, estimates of λ_i can be obtained from probit estimates of equation (5a) giving us the probability that the individual is included in the sample that answered a numerical W^*. In summary, the method is a two-stage estimation procedure. First, we estimate the probability that an individual was willing to accept a job at some wage. Second, we use the probit estimates to calculate λ_i which we then add as a regressor in (6) for the sample that *did* give a numerical wage.[31]

30 In fact, a few individuals responded that they would accept a steady job at the same pay. To simplify the analysis, I set W^* equal to the money wage rate for these men. All other individuals who did not answer the question are combined with the sample who said that they would not take a job at any wage.

31 In fact, to increase efficiency it is preferable to use generalized least squares in the second stage. The use of ordinary least squares does yield coefficients of (6) which are consistent, but their standard errors are not. To maintain simplicity, this refinement is not pursued in this paper.

TABLE 7
RESERVATION WAGE EQUATIONS

Dependent: Willingness to Accept Another Job			Dependent: ln(Reservation Wage)				
Variable	Coeff.	*t*-ratio	Row	*U*	*t*	*U·t*	λ
U	−.1852	(−2.66)	(1)	−.0521	.0102	—	—
W	−.0298	(−2.08)		(−.72)	(3.48)		
MARRIED	.0666	(.51)	(2)	.0455	.0160	—	−.6479
WIFED	−.0247	(−2.40)		(.84)	(2.28)		(−1.13)
EXPER	−.0270	(−3.38)	(3)	.1852	.0164	−.0184	—
EDUC	.0008	(.05)		(1.80)	(4.69)	(−3.20)	
CHILD	.0647	(2.16)	(4)	.2192	.0185	−.0163	−.3692
HEALTH	.0588	(.79)		(1.93)	(2.63)	(−2.83)	(−1.04)
INCOME	.0131	(3.24)					
TENURE	−.0159	(−5.86)					
Log likelihood	−1199.9						

Table 7 presents both the probit regression and the union and tenure coefficients from earnings functions which used the reservation wages as the dependent variable. In keeping with the interpretation of I_i as an inverse measure of job satisfaction, the specification of (5a) is similar to the regressions shown in Table 1, while the specification of (5b) is similar to the earnings functions presented in Table 5. The probit regression is somewhat surprising. For instance, we find that individuals who are in a union are less willing to accept another job offer at some wage. At the same time, however, for those who answer a reservation wage, we find that union members do *not* give a higher reservation wage. In any case, Freeman [13] has shown that the use of the variable I_i does not predict future quit behavior as well as the standard satisfaction variables discussed earlier. Thus, to get a correct picture of how unions affect reservations wages, it is best to concentrate on the estimates of equation (6).

As can be seen from Table 7, the coefficient of λ_i is negative but weak, and it does not change the results substantially from the standard least-squares regressions. However, the selectivity bias is such that introducing λ_i slightly increases the effect of unions on the reservation wage.[32] We find that unions have both a positive effect on reservation wages and a strong negative

32 It can be shown that the sign of the coefficient of λ_i depends on the correlation between ϵ_{1i} and ϵ_{2i}.

interaction effect. It is interesting to contrast these results with those presented in Table 5 using actual wages. The comparison of these point estimates is somewhat suggestive. Although the level effect is about the same in both reservation and money wages, the interaction effect is stronger (absolutely) in the reservation wage. Note that this is consistent with the earlier finding that dissatisfaction is increasing for union members as tenure increases, since the reservation wage (i.e., the value of the job) is falling at a faster rate than the money wage itself. Thus the union effect on reservation wages is never higher than the union effect on money wages so that the direction of the results is consistent with the findings presented earlier.

VI. SUMMARY

This paper presented an empirical analysis of the relationship between trade unions, wages, and job satisfaction. The major empirical finding is that, on average, union members report significantly *lower* levels of job satisfaction in the Mature Men NLS. Moreover, this result holds within occupational categories and across types of unions.

It was also reported that although this effect may be due to the fact that unpleasant jobs lead to union creation, accounting for this simultaneity did not affect the results that unions have a direct effect on job satisfaction. Moreover, it was also found that the union effect on job satisfaction was highly dependent on job tenure. In particular, union members expressed more dissatisfaction at higher levels of tenure. Thus, surprisingly, it is the older workers in the firm who report (as a result of unionization) low levels of job satisfaction.

This interesting result follows from the exit-voice effect of unions. That is, one way that unions become effective is to politicize the firm's labor force, leading to higher *expressed* levels of dissatisfaction. The empirical evidence presented in this paper showed that the increase in dissatisfaction with tenure was larger when the wage was *not* held constant in the satisfaction equation. Moreover, the results indicated that the union wage effect is strongest at the early years of job tenure and diminishes substantially past this point. Thus, the increasing dissatisfaction of union men with tenure is due to two factors: the politicization of the unionized labor force and the relative fall in the gains from being unionized as tenure rises. The importance of this latter effect is obvious since it implies that the rising dissatisfaction of union workers with tenure is partly real and should therefore affect quit probabilities. In particular, the empirical evidence suggested that unionization had a strong negative effect on quit probabilities at low levels of tenure, but that the effect diminishes (absolutely) as tenure increases.

Finally, some effort was made to quantify the measures of job satisfaction by using the concept of the reservation wage. Despite serious problems encountered in obtaining a clean measure of this variable, the union effect on the reservation wage (which presumably includes nonpecuniary components) is somewhat weaker than its effect on actual wages, and again the effect was a negative function of job tenure.

The role of unions in the labor market is currently becoming a fertile area for research in labor economics. The analysis in this paper, of course, is far from complete. Clearly, the results discussed here should be replicated in other samples and in different time periods to gain more understanding of the labor market effects of unions. Moreover, several questions deserve further scrutiny. In particular, it would be worthwhile to conduct a detailed analysis of the simultaneous relationship between job satisfaction, union membership, and wages since this study would lead to substantially improved estimates of the union effect on money and real wages. Hopefully, this paper represents an attempt to return to the problems considered by institutional labor economists with the added insights provided by the human capital approach. Perhaps this marriage of two frameworks can lead to new insights into the role of unions in the labor market.

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