BASEMENT

THE DEMAND FOR UNION REPRESENTATION

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ABSTRACT

    A model of the determinants of worker preferences for union
representation as distinct from their actual union status is developed and
estimated using data from the Quality of Employment Survey. However, data on
worker preferences were available only for the nonunion workforce, and two
solutions to this censored data problem are implemented with comparable
results. After correcting for this censoring, it was found that the demand
for union representation was higher among the workforce in general than among
the nonunion workforce. In addition, among the nonunion workforce only race
and seniority have statistically significant effects on worker preferences,
while among the workforce as a whole additional sharp distinctions arise in
the dimensions of region, occupation, and sex-marital status. This relative
flatness in preferences among the nonunion workforce is likely due to sorting
based on worker preferences.

I. Introduction

A relatively neglected area of research on labor unions is the determination of the union status of workers. In order to understand the process through which labor unions developed, what their future holds, and what their effects on workers, the workplace, and compansation are, it is crucial that a thorough understanding of this issue be gained. It is argued here that there is an excess demand by workers for vacancies in existing union jobs so that the actual union status of workers is determined by a combination of worker demand for union representation and the decisions of union employers as to who to hire from the queue of workers. This makes union status an unreliable indicator of worker preferences, and a unique set of data is utilized in the study which isolates worker preferences for union representation as opposed to their actual union status. Thus, the analysis addresses an important piece of the larger and as yet unsolved problem.

The analysis of the determination of the union status of workers is complicated by the distinction which must be drawn between the union status of workers and the union status of jobs. Nonunion jobs are unionized by convincing enough (a substantial minority) of workers to sign cards petitioning the National Labor Relations Board (NLRB) for an election and subsequently having a plurality of votes cast in the election favorable to repesentation by the particular union. The union status of these jobs remains with the jobs and does not inhere in the workers. If a worker who originally voted for union representation leaves the job, his future union status is determined by the union status of his future jobs. Similarly, a new worker hired to fill the vacancy left by the worker who left the union

job would automatically be represented by the union.[1] Since the number of workers voting in NLRB representation elections in any year is relatively small, this implies that most workers' union status is determined by the union status of the jobs they take.[2]

Given this process it is important to ask about the nature of a market equilibrium for union representation. Clearly there are nontrivial costs associated with the effort required to organizing nonunion jobs. Union organizers and sympathizers must convince enough workers of the correctness of their view, often in the face of sophisticated opposition from the employer, both to hold and to win an election. The benefits of union representation, the value of which varies across workers, can take the form of increased wages and fringe benefits, improved job security, the protection of a formal grievance settlement mechanism, and other nonpecuniary changes in the terms and conditions of employment. It is reasonable to argue that, given cross-sectional variation in the costs and benefits, organization of nonunion jobs is attempted to the point where the expected costs are roughly equal to the expected benefits.

While this discussion implies that there is an equilibrium in the process of creation of new union jobs through organization, the fact noted above that union status inheres in a job after the worker who "organized" it leaves combined with the fact that the job becomes available almost costlessly to other workers suggests that there will be excess demand for vacancies in existing union jobs.[3] In this disequilibrium context the union status of most workers is determined in two stages. First, workers decide whether or not they want a union job. Second, because in most industrial

settings the employer has control over hiring, the employer decides which of the workers in the queue it wishes to hire. Only if the worker desires a union job and the employer selects the worker will the worker be "unionized".

In most studies which develop structural models of the determination of the union status of workers (e.g., Ashenfelter and Johnson, 1972; Lee, 1978) it is argued that union status is determined strictly as a result of worker preferences for unionization. However, in light of the above discussion this model is inadequate because it neglects the inherent excess demand for union jobs and the resultant employer selection problem. An exception is Abowd and Farber (1982) who attempt an analysis of the determination of a worker's unions status in which a distinction is drawn between worker preferences and employer choice criteria. However, their analysis is hampered by the fact that only the final outcome (uinion status) is observed, and it is impossible to determine whether nonunion workers did not want a union job, could not get a union job, or both. This difficulty is compounded by the problem that a worker's current preference for a union job, given accrued seniority, may be different from what it was at the time the worker took his current job. For example, a nonunion worker with ten years seniority may not want a union job even if one were offered at that point. However, the worker may have preferred a union job ten years earlier but was not offered one.

In this study, a rather unique data set is employed which can be used to identify for nonunion workers their preferences for unionization <u>holding seniority fixed</u>. Thus, one-half of the process through which worker union status is determined can be investigated. However, these data do not permit

investigation of the employer selection process, and a satisfactory solution
of the entire puzzle must await future research.

In the next section, a simple empirical model of worker preferences for
unionization is developed based on utility maximization by individual
workers. Section III contains a discussion of the data set from the 1977
Quality of Employment Survey (QES) developed by the Survey Research Center at
the University of Michigan. Particular attention is paid to the
interpretation of the question, "If an election were held with secret
ballots, would you vote for or against having a union or employees'
association represent you?" The response to this question (VFU) serves as
the basis for the analysis in this study.

Empirical implementation of the model developed in Section II is
hampered by the problem that the crucial question (VFU) was asked only of
nonunion workers. Analysis of the responses in this context is interesting
in that insight can be gained into the characteristics of nonunion workers
which lead them to desire union representation conditional on their nonunion
status.[4] However, for an analysis of preferences for union representation
unconditional on union status, the data are censored on the basis of an
obviously related variable. Two potential solutions to this problem are
developed in Section IV. The first relies on the assumption that all union
workers would have answered the VFU question in the affirmative at the time
they started their job. The second approach is to specify an econometric
model which takes account of the censoring and yields consistent estimates of
unconditional preferences for union representation.

The model is estimated in Section V both with and without accounting for the censored sample problem. The two sets of unconditional estimates derived using the different approaches outlined above yield remarkably similar results which contrast sharply with the estimates derived conditionally on nonunion status. The estimates suggest that preferences for union representation among the nonunion workforce are relatively flat across most individual characteristics, while among the workforce as a whole there are sharp distinctions. This result is likely due to sorting in the sense that many workers who desire union representation on the basis of both their observed and their unobserved characteristics are already working on union jobs, leaving a group of nonunion workers whose preferences for union representation show little systematic variation with characteristics. Among the nonunion workforce only race, age, and seniority have a statistically significant effect on worker preferences.

The empirical results are interpreted in Section VI, and the final section of the paper contains a synthesis of the results along with conclusions which can be drawn from the analysis.

II. A Simple Empirical Model of Worker Preferences for Union Representation

At its simplest level, a worker's preference for union representation on a job versus no such representation can be modeled as a comparison by the worker of the utilities yielded to him by his job in each case. His preference will be for that case which yields him the largest utility. A worker's utility on the job is affected by many factors, including wages, fringe benefits, safety, job security, comfort, etc. In addition, there are

subjective factors such as satisfaction with supervision, perceived fairness

of treatment, equitable comparisons with others, perceived chances for

promotion, etc. In the absence of explicit measures of most of these factors

for each worker in both a union and a nonunion environment, it is argued that

these utilities vary across workers as functions of their personal and

occupational characteristics as well as the measurable characteristics of the

union and nonunion jobs.

In order to examine this argument more carefully, assume, as is done

below, that the only explicit job characteristic which can be measured for

each worker in both a union and a nonunion context is the wage rate. More

formally, it is argued that

$$(1) \quad V_u = V_u(Z, W_u)$$

and

$$(2) \quad V_n = V_n(Z, W_n)$$

where V_u and V_n represent the worker's union and nonunion utilities

respectively, Z represents a vector of personal and occupational

characteristics, and W_u and W_n represent the worker's union and nonunion wage

rates respectively. The worker's preference for union representation can be

expressed by computing $y = V_u - V_n$. If this difference in utilities (y) is

positive, then the worker will opt for union representation. If y is

negative, then the worker will not opt for union representation. It is clear

that this difference in utilities can be expressed as

$$(3) \quad y = y(Z, W_u, W_n),$$

where variations in Z capture variations in the difference between the

worker's union and nonunion utilities.

In order to derive the empirical analog of this model, a specific functional form must be selected for y. This is

(4) $y = Z\gamma + \delta(\ln W_u - \ln W_n) + \varepsilon_1$

where γ is a vector of parameters, δ is a parameter expected to be positive, and ε_1 represents unmeasured components of the utility difference and is assumed to be distributed normally with zero mean. Clearly, if y, Z, W_u, and W_n were observable for all workers, the parameters in equation (4) could be estimated using straightforward linear methods. Unfortunately not all of these quantities are observed for all workers.

Two important variables required for all workers are the union and nonunion wage rates. However, at best only one wage or the other is observed, and a question arises as to how to handle this problem. The difficulty is that it is likely that the union-nonunion wage differential, and hence the union and nonunion wage rates, are important determinants of ultimate union status as well as of worker demands for union representation. This raises potentially serious problems of selection bias in estimating union and a nonunion earnings functions. In addition, the fact that the observed wage is likely to be correlated with union status makes it improper to estimate the model conditional on this wage rate. The correlation must be accounted for.

One approach toward solution of this problem is to specify union and nonunion individual earnings functions as

$\ln W_u = X\beta_u + \varepsilon_u$ and

(15)

$\ln W_n = X\beta_n + \varepsilon_n$

where X is a vector of exogenous characteristics, β_u and β_n are vectors of parameters, and ε_u and ε_n are normally distributed unobserved elements. The difficulties stem from the possibility that ε_u and/or ε_n are correlated with ε_1 and ε_2. If the representations of $\ln W_u$ and $\ln W_n$ from equation (5) are substituted into equation (4), the resulting union-nonunion utility difference can be expressed as

(6) $y = Z\gamma + \delta X(\beta_u - \beta_n) + \varepsilon_3$

where $\varepsilon_3 = \varepsilon_1 + \delta(\varepsilon_u - \varepsilon_n)$ and is distributed normally unconditional on union status. If consistent estimates of β_u and β_n were available then estimates of γ and δ could be obtained conditional on these values of β_u and β_n. While there are well known techniques to derive theoretically consistent estimates (Heckman, 197 ; Lee, 1978, 1979), recent work by Freeman and Medoff (1981) has documented convincingly the problems in applying these techniques in the context of the estimation of union wage effects.[5]

Given the difficulties in deriving consistent estimates of β_u and β_n from data on wage rates and union status, one could attempt to estimate γ, δ, β_u and β_n directly from equation (6). However, the structure of equation raises some serious identification problems because only for those elements of Z which are not included in X can the associated γ's be estimated. Similarly, only the product of $\delta(\beta_u-\beta_n)$ can be estimated, and then only for elements of X which are not included in Z. What can be identified is a reduced form version of equation (6) which is specified by substituting $Z*\gamma*$ $= Z\gamma + \delta X(\beta_u - \beta_n)$ into equation (6) which yields

(7) $y = Z*\gamma* + \varepsilon_3$.

the vector $Z*$ contains all of the variables which are in either Z or X, and

γ* is the vector of reduced form parameters. The penalty for specifying this reduced form model is that wage and nonwage variation in the union-nonunion utility difference as they vary by individual characteristics cannot be distinguished. Nonetheless, in the face of the problems outlined above the analysis proceeds using this reduced form specification.

Another problem which must be solved before the parameters of the model can be estimated is that y, which represents the difference in a worker's utility between union and nonunion status on the job, is not observed. All that is observed in the sign of y through the worker's response to the VFU question (y > 0 <=> VFU=1, y < 0 <=> VFU=0). The limited nature of the dependent variable implies that Pr(VFU=1) = Pr(y > 0), yielding from equation (7) that

(8) $Pr(VFU = 1) = Pr(\varepsilon_3 > - Z*\gamma*)$.

If ε_3 is assumed to be distributed normally with zero mean and unit variance, then equation (7) implies a probit specification for a likelihood function.[6] The contribution of any individual to the log-likelihood function is

(9) $L = VFU \ln[\Phi(Z*\gamma*)] + (1 - VFU)\ln[1 - \Phi(Z*\gamma*)]$,

where $\Phi(\cdot)$ represents a standard normal cumulative distribution function.[7] Before this model is estimated the discussion turns to an examination of the data and their limitations.

III. The Data

The data used are from the 1977 cross-section of the Quality of Employment Survey (QES) developed by the Survey Research Center of the University of Michigan. The QES contains data for approximately 1500

randomly selected workers (both union and nonunion) on their personal
characteristics and job attributes.[8] The crucial bit of information for this
study is the response to the question asked only of nonunion workers, "If an
election were held with secret ballots, would you vote for or against having
a union or employee's association represent you?" This variable is called
"Vote-for-union?" or VFU. It is interpreted here as the preference of a
worker for union representation on his current job. Thus, it holds all job
characteristics fixed except those which the worker expects the union to
affect. It is assumed that the worker's response is based on his current
utility as compared with what the worker's utility is expected to be were the
job to be unionized.[9]

A sample of workers was derived from the QES by selecting those workers
for whom the survey contained valid information on the variables listed in
Table 1. Self-employed workers, managers, sales workers, and construction
workers were deleted from the sample. The remaining sample contains 915
workers. Table 1 contains descriptions of the variables used in the study as
well as their means and standard deviations for both the entire sample and
the union and nonunion subsamples. The base group for the dichotomous
variables are white, nonsouthern, single, male blue collar workers with
twelve years of education. On average, the 37 percent of the sample who are
unionized are slightly older and are more likely to be male, married,
nonwhite, nonsouthern, and in a blue collar occupation.

Thirty-seven percent of the nonunion sample expressed a preference for
union representation (VFU = 1). It is unfortunate that the analogous
question was not asked of union members (If an election were held by secret

Table 1. Means (Standard Deviations) of Data
Quality of Employment Survey, 1977

Variable	Description (Dichotomous variables = 0 otherwise)	Combined Sample (n=915)	Union Sub-Sample (n=337)	Non-Union Sub-Sample (n=578)
U	= 1 if works on union job	.368	--	--
VFU	= 1 if desires union represent.	--	--	.370
Age	age in years	36.8 (13.1)	38.2 (12.6)	35.9 (13.3)
Sen	firm seniority in years	6.90 (7.49)	9.48 (8.18)	5.40 (6.60)
Age_o	Age-Sen	29.9 (10.8)	28.7 (9.28)	30.5 (11.5)
Age_n	Age - U(Sen)	33.3 (12.5)	28.7 (9.28)	35.9 (13.3)
Sen_n	Sen(1-U)	3.41 (5.86)	0	5.40 (6.60)
Fe	= 1 if female	.419	.329	.471
Marr	= 1 if married w/spouse present	.640	.709	.600
Marr*Fe	= 1 if Fe = 1 and Marr = 1	.198	.181	.208
NW	= 1 if nonwhite	.137	.160	.123
South	= 1 if worker resides in South	.353	.237	.420
Ed < 12	= 1 if <12 years education	.223	.258	.202
12<Ed<16	= 1 if >12 years & <16 years educ.	.212	.166	.239
Ed > 16	= 1 if >16 years education	.201	.202	.201
Cler	= 1 if occupation is clerical	.205	.116	.258
Serv	= 1 if occupation is service	.156	.119	.178
Prof&Tech	= 1 if occupation is professional or technical	.234	.211	.247

ballot, would you vote to continue union representation?). This would make information available for all workers about worker preferences for unionization of their current job at the current time. The lack of this information poses some important problems of econometrics and interpretation. It is to this and related problems that the next section is addressed.

IV. Censored Sample Problems and Options

If the question of interest is the estimation of a model of nonunion worker's preferences for union representation conditional on their nonunion status and if the assumption is made that ε_3 has a standard normal distribution conditional on the workers being nonunion, then maximum likelihood estimation applied to the likelihood function implied by equation (9) over all of the nonunion workers in the sample will lead to consistent estimates of γ and δ. However, these estimates cannot be interpreted, a priori, as those which describe a model applicable to all workers regardless of union status.

Two approaches to this problem can be identified. First, it can be assumed that all union workers would have answered VFU = 1 at the time of hiring.[10] As long as the values of the appropriate variables in Z* are adjusted to the time of hiring for union workers, the probit likelihood function implied by equation (9) can be maximized over the entire sample of workers. These maximum likelihood estimates of γ^* can be interpreted as describing a model of worker preferences unconditional on union status.

The variables in Table 1 which obviously change over time are age and seniority.[11] In order to adjust these values to the time of hire for union

workers, seniority is subtracted from each measure for union workers creating

two new measures: Sen_n and Age_n. The seniority measure is in fact nonunion

seniority because all union workers have $Sen_n = 0$. This is appropriate in

light of the fact that seniority on a union job and seniority on a nonunion

job are qualitatively different and are likely to have different effects on

worker demands for union representation. Note that if current data were

available on current VFU for union workers than a separate coefficient could

be estimated for union seniority. However, such data are not available, and

even if it were assumed that <u>current</u> VFU = 1 for all union members then the

coefficient on union seniority would be unidentified.

The second approach to the problem of the lack of data on VFU for union

workers is to specify an econometric model of the joint determination of VFU

and union status which can account for the censoring of the data. A simple

empirical model of the determination of the union status of workers can be

specified of the form

(10) $S = C\alpha + \varepsilon_2$.

where S is an unobservable latent variable determining union status, C is a

vector of worker and job chartacteristics at the time of hiring, α is a

vector of parameters, and ε_2 is a random component with a standard normal

distribution which captures unmeasured aspects of the union status decision.

If S is positive then the worker works on a union job (U=1), and if S is

negative then there is no union representation on the job (U=0). Thus,

$Pr(U=1) = Pr(S > 0)$, which implies that

(11) $Pr(U=1) = Pr(\varepsilon_2 > -C\alpha)$.

Given the normality assumption regarding ε_2, the contribution to the log-

likelihood function regarding union status is a probit of the form

(12) $L = U \ln[\Phi(C\alpha)] + (1 - U) \ln[1 - \Phi(C\alpha)].$

In light of the introductory discussion of the process by which union status is determined through separate decisions by workers and employers, the behavioral underpinnings of this probit model are deliberately left vague. It is to be interpreted as a reduced form empirical relationship describing the union status of a worker. However, a note of caution is required. It is difficult (if not impossible) to think of a structural model of the determination of the union status of a worker where both the worker and employer make separate decisions which would have a reduced form which could be described as a simple univariate probit. In fact, this study is motivated, at least in part, by a desire to model the union status determination of workers in a manner which is consistent with separate worker and employer decisions in order to move away from the behaviorally naive structural model which is implicit in the simple probit model described here.

The statistical problem with making inferences about worker demand's for union representation unconditional on union status by analyzing the nonunion subsample exclusively is made clear through the following argument. Assume that the random components in the VFU and U functions (ε_3 and ε_2) are correlated ($\rho_{32} \neq 0$ where ρ_{32} represents the correlation between ε_3 and ε_2). In particular, ε_3 was assumed to have a normal distribution with zero mean unconditional on union status, but the ε_3 are observed only for nonunion workers. The condition for a worker being nonunion from equation (8) is that $\varepsilon_2 < -C\alpha$ so that ε_3 is observed only if $\varepsilon_2 < -C\alpha$, and the likelihood must be written in terms of this conditional distribution. Using Bayes' Rule and

assuming joint normality of ε_3 and ε_2, the conditional distribution of ε_3 given $\varepsilon_2 < -C\alpha$ is

$$(13) \quad f(\varepsilon_3 | \varepsilon_2 < -C\alpha) = \frac{\int_{-\infty}^{-C\alpha} b(\varepsilon_3, \varepsilon_2; \rho_{32})d\varepsilon_2}{\Phi(-C\alpha)}$$

where $b(\cdot, \cdot; \rho)$ represents a standardized bivariate normal density function. This conditional distribution is non-normal and involves the parameters α and ρ_{32}.

Two points are worth noting here. First, if ε_3 and ε_2 are uncorrelated and so that $\rho_{32} = 0$, then $\varepsilon_3 | \varepsilon_2 < -C\alpha$ is distributed as a standard normal and the likelihood function on VFU implied by equation (9) and estimated over the sample of nonunion workers can be interpreted correctly as that relevant to all workers unconditional on their union status. However, it seems likely that unmeasured determinants of the union status of a worker and of the worker's preference for union representation are correlated so that it is unlikely that $\rho_{32} = 0$. This potential restriction will be tested in the next section. The second point to note is that where $\rho_{32} \neq 0$, the conditional interpretation given to the probit VFU likelihood function derived from equation (9) and the conditional likelihood function for VFU derived from the bivariate normal model in equation (13) are inconsistent with each other because $f(\varepsilon_3 | \varepsilon_2 < -C\alpha)$ is non-normal. Nonetheless, estimates from the simple probit VFU likelihood function maximized over the nonunion subsample will be interpreted as estimates of worker preferences for union representation conditional on being nonunion.

While consistent estimates of the parameters of the unconditional model could be derived from maximization over the nonunion subsample of the

appropriate conditional likelihood function derived from equation (13), the
parameters of the union status model (α) would also have to be estimated. It
is obviously asking a lot of the data and the functional form specification
to identify the parameters of the union status determination model strictly
from data on nonunion workers. A much more efficient approach is to use
information from the whole sample to write the joint likelihood of
preferences for union representation and union status while noting that VFU
is censored for union workers. For nonunion workers who desire union
representation the appropriate contribution to the likelihood function is
$\Pr(\varepsilon_3 > -Z^*\gamma^*, \varepsilon_2 < - C\alpha)$. Given the distributional assumption, this is

$$(14) \quad \Pr(VFU=1, U=0) = \int_{\infty}^{-C\alpha} \int_{\kappa}^{\infty} b(\varepsilon_3, \varepsilon_2; \rho) d\varepsilon_3 d\varepsilon_2$$

where $\kappa = -Z^*\gamma^*$. Similarly, for nonunion workers who do not desire union
representation, the appropriate contribution is $\Pr(\varepsilon_3 < -Z^*\gamma^*, \varepsilon_2 < -C\alpha)$,
which yields

$$(15) \quad \Pr(VFU=0, U=0) = \int_{-\infty}^{-C\alpha} \int_{-\infty}^{\kappa} b(\varepsilon_3, \varepsilon_2; \rho) d\varepsilon_3 d\varepsilon_2.$$

Finally, for union workers no information regarding VFU is known so that ε_3
is integrated out, and the contribution of these workers to the likelihood
function is a univariate normal CDF representing $\Pr(\varepsilon_2 > -C\alpha)$, which yields

$$(16) \quad \Pr(U=1) = \Phi(C\alpha).$$

Using these results, the contribution of a worker to the log-likelihood
function accounting for the sample censoring is

$$(17) \quad L = (VFU)(1 - U) \ln [\Pr(VFU=1, U=0)]$$
$$+ (1 - VFU)(1 - U) \ln [\Pr(VFU=0, U=0)] + U \ln [\Pr(U=1)]$$

where the relevant probabilities are defined in equations (14) - (16).

The two approaches to the censored data problem ought to lead to similar estimates of the parameters of the VFU model ($\gamma*$) if the assumptions which they are based on are valid. In the next section these models are estimated along with the model conditional on nonunion status.

V. Estimation

The vector, $Z*$, of variables included in the reduced form worker preference relationship includes all variables which appear in either the structural union preference function (Z) or in the earnings functions (X). This set of varaibles includes three dichotomous variables for various levels of education, nonunion seniority, age (adjusted to start of job for union workers), dichotomous variables for nonwhite, southern residence, female, married with spouse present, and the interaction of female and married, and three dichotomous variables representing broad occupational classifications. All variables are defined in Table 1. The base group for the sample consists of white, nonsouthern, single males with twelve years of education working in a blue collar occupation.

The estimates derived for $\gamma*$ by maximizing the likelihood function implied by equation (9) over the 578 nonunion workers are contained in the first column of Table 2. The only variables which have a statistically significant effect at the five percent level on a nonunion worker's preference for union representation are race and seniority. It is also true that nonunion married females have a significantly lower (at the .05 level) probability of desiring union representation than do single females. The

Table 2:   Estimates of Pr(VFU=1)

	Pr(VFU=1|U=0) Univariate (1)	Pr(VFU=1) Univariate (2)	Pr(VFU=1) Bivariate (3)
Constant	.0248 (.218)	1.03 (.168)	.890 (.218)
Ed < 12	.0696 (.166)	.0442 (.135)	.0968 (.132)
12 < Ed < 16	-.101 (.146)	-.187 (.122)	-.145 (.123)
Ed $\geq$ 16	.118 (.197)	.0988 (.161)	.148 (.163)
Sen_n	-.0263 (.0103)	-.0838 (.00880)	-.00930 (.0120)
Age_n	-.00791 (.00494)	-.0129 (.00427)	-.0133 (.00436)
NW	.832 (.174)	.840 (.150)	.763 (.166)
South	.00669 (.115)	-.232 (.0960)	-.281 (.128)
Fe	.198 (.189)	.179 (.157)	.162 (.158)
Marr	-.102 (.176)	.204 (.143)	.100 (.158)
Marr*Fe	-.199 (.237)	-.328 (.197)	-.226 (.199)
Cler	-.132 (.170)	-.437 (.139)	-.468 (.159)
Serv	.0811 (.176)	-.158 (.146)	-.234 (.170)
Prof & Tech	-.169 (.189)	-.320 (.151)	-.399 (.169)
P_{32}	--	--	.923 (.198)
n	578	915	915
ln L	-351.9	-507.3	-897.6

Notes: Numbers in parentheses are asymptotic standard errors. The estimates in column (3) are derived jointly with Pr(U=1) in column (2) of Table 5.

flatness of this relationship is corroborated by the fact that the hypothesis that all of the coefficients equal zero except those for the constant, race, and seniority cannot be rejected at even the .25 level of significance using a likelihood ratio test.[12]

In order to investigate how sensitive nonunion worker preferences for union representation are in a number of dimensions, Table 3 contains values of $Pr(VFU=1|U=0)$ at the point estimates of the parameters contained in the first column of Table 2 for 30 year old single males with twelve years of education, zero seniority, and various occupations, races, and regions.[13] It is clear that among the nonunion workforce for any occupation and region, nonwhites are substantially more likely to desire union representation than are whites. On the other hand, region has a trivial effect while occupational variation is moderate.

Though not evident in Table 3 the estimates also suggest that more senior nonunion workers are less likely to desire unionization of their job. To get some idea of the magnitude of the seniority effect, note that on the basis of the estimates a forty year old nonsouthern nonunion white single male blue collar worker with twelve years of education and no seniority has $Pr(VFU=1)$ = .39 while for a similar worker with ten years seniority $Pr(VFU=1)$ = .29 and with twenty years seniority $Pr(VFU=1)$ = .21

The first technique for deriving estimates of the worker preference model unconditional on union status was to assume that all union workers had VFU=1 at the time they were hired. These estimates were obtained by maximizing the likelihood function implied by equation (9) over the entire sample of 915 workers, and they are contained in column (2) of Table 2. As

Table 3. Predicted Pr(VFU=1) for Workers
Varying by Race, Region, and Occupation at Zero Seniority

	Pr(VFU=1|U=0)		Pr(VFU=1) Univariate		Pr(VFU=1) Bivariate	
	Non-South	South	Non-South	South	Non-South	South
Blue Collar						
white	.416	.418	.740	.659	.688	.583
nonwhite	.732	.734	.931	.894	.895	.835
Clerical						
white	.365	.368	.582	.489	.509	.398
nonwhite	.687	.689	.852	.792	.784	.693
Service						
white	.448	.450	.686	.560	.601	.490
nonwhite	.758	.760	.907	.863	.846	.770
Professional and Technical						
white	.351	.354	.627	.536	.537	.425
nonwhite	.674	.676	.878	.824	.804	.717

Computed for 30-year-old single males with 12 years education and no
seniority. Pr(VFU=1|U=0) computed from estimates in column (1), Table 2.
Pr(VFU=1), univariate, computed from estimates in column (2), Table 2.
Pr(VFU=1), bivariate, computed from estimates in column (3) of Table 2.

with the nonunion subsample, race and nonunion seniority have statistically significant effects on worker preferences for union representation. However, in contrast to the nonunion workforce, sharp distinctions between workers' preferences arise on the base of age, region, and occupation. Older workers, southern workers, clerical workers, and professional and technical workers are signficantly less likely to desire union representation at the .05 level.

Somewhat weaker statistical relationships are found between the sex and marriage variables and preferences for union representation among the workforce as a whole. Married males and single females are significantly more likely (at the .08 and .13 levels respectively) to prefer union representation than are single males. In contrast, married females are indistinguishable from single males in their preferences for union representation.

Table 3 contains values of Pr(VFU=1) at the point estimates of γ^* contained in the second column of Table 2 for thirty year old single males with twelve years of education and no seniority. The hypothetical worker's occupation, race, and region are varied in order to investigate the sensitivity of Pr(VFU=1) to these factors. It is clear that race and occupation have large effects on workers' preferences for union representation. Nonwhites are substantially more likely to prefer union representation. Professional and technical workers are the least likely occupation to prefer union representation, while blue collar workers are most likely. Southern workers are somewhat less likely than nonsouthern workers to desire union representation.

It is interesting to contrast the preferences of nonunion workers for union representation to workers' preferences unconditional on union status. The calculated probabilities contained in Table 3 facilitate this comparison. It is clear that workers in general are more likely to desire union representation than nonunion workers. For example, for the four types of blue collar workers listed in Table 3 (combinations of race and region), the probability that workers in general desire union representation is on average 46 percent higher than the probability that workers desire union representation conditional on being nonunion. The difference is positive but less pronounced for the other listed occupations. Two other differences, apparent from Table 3, are that region plays a much greater role in determining overall preferences for unionization than it does among the nonunion workforce. In addition, the occupational distinctions are much greater among the workforce in general than among the nonunion workforce.

The second technique for obtaining estimates of γ^* unconditional on union status was to maximize the joint likelihood function on VFU and U implied by equations (14) - (17). These estimates are contained in column (3) of Table 2.[14] There is remarkable agreement between the estimates obtained using the two techniques. With one important exception (seniority) the coefficients of all the variables are of comparable magnitude and statistical significance. Once again, among the workforce as a whole younger, nonwhite, nonsouth, blue collar workers are significantly more likely to desire union representation. This is illustrated with the predicted probabilities contained in Table 3. In addition, the sex-marriage

pattern is similar to that obtained with the first method although these results do not approach statistical significance at reasonable levels.

The chief difference between the results obtained using the two techniques is that nonunion seniority has no effect on worker preferences using the bivariate approach while it is estimated to have a large negative influence using the univariate model assuming that VFU=1 for all union members. Technically, this difference probably arises because $Sen_n = 0$ for all union members which results in many observations with $Sen_n = 0$ and VFU=1 in the univariate model while these observations never appear in the VFU portion of the bivariate model due to the censoring of the data. The importance of the difference in the estimates is highlighted in Table 4 which contains estimates of Pr(VFU=1) using the two techniques for various combinations of age and nonunion seniority. Using the univariate estimates Pr(VFU=1) falls off rapidly to very small values (< .10) while it falls only slowly using the bivariate estimates. While there is no clear way to reconcile these results, the inclination is to accept the bivariate results because they use less information by not relying on the assumption which introduced $Sen_n = 0$ for all union workers. However, a firm conclusion ought to await clearer evidence.

Table 5 contains the maximum likelihood estimates of the determination of union status function. In column (1) are the univariate probit results based on the likelihood function implied by equation (12) while in column (2) are the bivariate probit results based on the likelihood function implied by equations (14) - (17). These results are virtually identical which is not unexpected given that they are consistent estimates of the same reduced form

Table 4:   Predicted Pr(VFU=1) for Blue Collar Workers
Varying by Age and Nonunion Seniority

	Pr(VFU=1) Univariate	Pr(VFU=1) Bivariate
Age = 20		
Sen_n = 0	.780	.734
Age = 30		
Sen_n = 0	.740	.688
= 10	.442	.655
Age = 40		
Sen_n = 0	.696	.640
10	.373	.604
20	.123	.568
Age = 50		
Sen_n = 0	.650	.589
10	.325	.552
20	.0984	.516
30	.0166	.478
Age = 60		
Sen_n = 0	.601	.537
10	.280	.500
20	.0778	.463
30	.0120	.426
40	.000981	.390

Computed for white single males with twelve years education living in the
nonsouth and working in a blue collar occupation. Pr(VFU), univariate,
computed from estimates in column (2), Table 2. Pr(VFU), bivariate, computed
from estimates in column (3), Table 2.

empirical relationship described earlier, and, given the ambiguity regarding its behavioral underpinnings, not much space will be allocated to evaluation of these results. Suffice it to say that nonwhites and nonsoutherners are more likely to be union members, as are younger workers. In addition, there are rather sharp occupational breaks which imply that blue-collar workers are most likely to be unionized, while clerical workers are least likely, holding other factors fixed. Finally, married males are significantly more likely at the .05 level to be unionized than single males. The same pattern exists for females though at slightly larger significance levels. The propensity of males to be unionized is indistinguishable from that for females.

Note that the univariate probit model explaining worker preferences conditional on nonuion status combined with the univariate probit model explaining union status is a constrained version of the bivariate model where $\rho_{32} = 0$. The estimated value of ρ is .923 with an asymptotic standard error of .198. The positive value estimated for ρ_{32} suggests that unobserved factors which make workers more likely to work on union jobs also make these workers more likely to prefer union representation. The hypothesis that $\rho_{32} = 0$ can be rejected at the 5 percent level of significance using a two-tailed asymptotic t-test. A likelihood ratio test of the same hypothesis can be performed by summing the log-likelihoods for the first column of Table 2 and the first column of Table 5 and comparing the constrained log-likelihood to the unconstrained value. Using this test, the hypothesis that $\rho = 0$ cannot be rejected at conventional levels of significance.[15] The lack of agreement of these tests is surprising given their asymptotic equivalence.

Table 5: Estimates of Pr(U=1)

	Pr(U=1) Univariate (1)	Pr(U=1) Bivariate (2)
Constant	.364 (.181)	.386 (.180)
Ed < 12	.0922 (.126)	.0899 (.125)
12 < Ed < 16	-.156 (.125)	-.181 (.123)
Ed > 16	.145 (.172)	.182 (.170)
Age_o	-.0141 (.00472)	-.0141 (.00461)
NW	.316 (.134)	.321 (.134)
South	-.542 (.0965)	-.554 (.0956)
Fe	-.0269 (.159)	-.0244 (.160)
Marr	.272 (.136)	.259 (.137)
Marr*Fe	-.0571 (.197)	-.0706 (.198)
Cler	-.689 (.140)	-.689 (.140)
Serv	-.509 (.137)	-.503 (.138)
Prof & Tech	-.506 (.168)	-.559 (.165)
n	915	915
ln L	-546.3	-897.6

Notes: Numbers in parentheses are asymptotic standard errors. The estimates in column (2) are derived jointly with Pr(VFU=1) in column (3) of Table 2.

However, in light of the important changes in γ^* between the conditional and unconditional models, it is clear that allowing ρ to differ from zero makes a difference from a substantive standpoint.

VI.   Interpretation and Implications of the Analysis

The most striking empirical result described above is that there is relatively little variation in workers' demands for union representation among the nonunion workforce while there is substantial variation among the workforce as a whole. This suggests that prior sorting into union and nonunion jobs has taken place on the basis of worker preferences so that those in any observable subgroup who were most likely to desire union representation are, in fact, working on union jobs. The evidence that the unobservable characteristics which affect worker preferences and actual union status are positively correlated ($\rho_{32} > 0$) is consistent with this

interpretation. The implication is that workers in nonunion jobs are relatively homogenous in their (lack of) desire for union representation. For example, among the workforce as a whole nonsouthern workers are significantly more likely than southern workers to desire union representation. Over time, perhaps partly as a result of this fact, more organization has taken place outside the south so that a higher proportion of jobs are unionized outside the south. At the same time, within each region the union jobs are filled with the workers who are most likely to desire union representation. The result is that, while the nonunion workers in the south are relatively more numerous than those outside the south, they are individually no less favorably disposed toward union representation than their nonunion workers outside the south. This could explain in part the

relatively low level of unionization in the south which coexists with
currently comparable levels of new organization in the two regions.[16] A
similar interpretation can be given to the contrast in the effect of
occupational distinctions on preferences for union representation between the
entire workforce and the nonunion workforce.

The two dimensions along which preferences do vary among the nonunion
workforce are seniority (and less strongly age) and race. Taking the first
of these, the fact that more senior and older nonunion workers are less
likely to desire union representation is evidence that workers perceive that
the benefits of unionization decline with age and seniority. It is crucial
to note here that the decision which is analyzed here (VFU) does not involve
sacrificing seniority as would be the case if the worker had to change jobs.
In this light, the result is not consistent with the hypothesis that more
senior and older workers ought to prefer union representation because unioni-
zation is associated both with improved fringe benefits such as pensions and
with more effective seniority rules as a criterion for job security and
promotion.[17] However, there is evidence (Abowd and Farber, 1982; Bloch and
Kuskin, 1978) which suggests that the union-nonunion wage differential is
inversely related to seniority and experience. This standardization of wage
rates, noted by the Webbs (1920), will give the older, more senior workers a
smaller incentive to desire union representation. However, that this is not
the entire explanation is clear from the analysis of Farber and Saks (1980)
who find a negative relationship between age and worker preference for
unionization even after controlling for the standardization of wage rates.

There are a number of reasons to expect that nonwhites are more likely
to desire union representation. Chief among these are a potential reduction

in the white-black wage differential associated with the standardization of wage rates, though the evidence on this is mixed (Ashenfelter, 1972; Abowd and Farber, 1982; Johnson and Youmans, 1971), and the "web of rules" which unions impose to limit the discretionary power of employers to discipline employees. It is not hard to envision that blacks may be less sanguine than whites about the fairness of treatment they can expect from their employers so that they will value such a web of rules imposed by unions more highly. What may be surprising about the empirical result is that the same type of sorting described above has not taken place on the basis of race so that there would be no apparent difference in this dimension among the nonunion workforce. The fact that there is a large difference on the basis of race among the nonunion workforce suggests that such sorting has not been carried as far as it could, perhaps because enough nonwhites are not able to translate their preferences for union representation into union jobs. The answer to this puzzle must await an analysis of the hiring criteria of union employers.

The additional relationships which emerge among the workforce as a whole involve regional, occupational, and sex-marriage distinctions in worker preferences for union representation. The regional distinction, that southern workers are less likely to desire union representation, has been noted above and is counter to evidence which suggests slightly larger union-nonunion wage differentials among southern workers (Abowd and Farber, 1982). This raises the question as to whether workers in the south are "anti-union" in some fundamental sense. Conceptually, this question could be answered with some measure of prior disposition toward unions in general.

Unfortunately, no objective measures exists, and subjective measures of attitudes toward unions are contaminated with perceptions of the worker's current situation.

The occupational breakdown is that clerical and professional-technical workers are less likely to desire union representation than are blue collar workers. This may be because these are predominantly white collar occupations and the workers perceive that unions, as historically blue collar organizations, are not appropriate to their situation. Additionally, unions of white collar employees may not provide the same sorts of advantages as do blue collar unions. This will be the case if nonunion employers of these workers provide "union-like" benefits in order to avoid unionization.

An interesting result which emerges from the sex-marriage breakdown is that single females and married males are indistinguishable in their preferences for union representation as are single males and married females. In addition, the former pair of groups is more likely than the latter pair of groups to desire union representation. What is interesting is that the group of workers which has the highest labor force participation (i.e., committment to the labor force), married males, is more likely to desire union representation than the group of workers with the lowest labor force participation rate, married females. To the extent that the benefits from union representation are in the form of fringe benefits such as medical insurance which may be duplicated on a spouse's job or pensions which are of more value to workers who expect to remain on the job, married females may value the services of unions less highly.

VII. Synthesis and Conclusions

A model of the determination of worker preferences for union representation was developed which was based on a comparison of utilities in union and nonunion settings. A distinction was drawn between the observed union status of workers and their current preferences for union representation which was based on costs of job mobility and the existence of queues for union jobs. A rather unique set of data, the Quality of Employment Survey, was used because it contained the response to a question which indicated directly a worker's preference for union representation. Unfortunately, this particular bit of information was available only for nonunion workers.

Two solutions to the censored data problem were implemented. First, it was assumed that all union workers desired union representation at the time of hiring. Second, a reduced form empirical model to explain union status and hence the censoring was developed under the assumption of joint normality of latent variables determining union status and preference for union representation. These distinct approaches yielded comparable results. The union preference function was also estimated using the model conditional on nonunion status. A comparison of the general nature of the results both yields some insight into the determination of the extent and locus of unionization and has important implications for prospects for organizing currently unorganized workers.

Overall, worker preferences for unionization among nonunion workers are rather flat in that, with the exception of race and seniority, there is little variation across workers with different characteristics. On the other

hand, a number of relatively sharp additional delineations in worker preferences for union representation along such dimensions as region, occupation, and sex-marital status occur in the model unconditional on union status. This suggests that many workers with those characteristics (both observed and unobserved) which make them likely to desire union representation are, in fact, union workers. The nonunion workers who are left are relatively homogeneous in their (lack of) interest in union representation. This interpretation is supported by the positive correlation estimated between the unobservable factors affecting preference for union representation and actual union status. With regard to the prospects for union organizing, this sorting suggests both that current nonunion workers will be less receptive to organizing efforts and that effective targeting of campaign efforts on the basis of gross characteristics such as region will not be terribly useful in light of the flatness of preferences.

In closing, two cautions are necessary. First, all of the results presented here were derived under the assumption of joint normality largely for computational convenience. The results may be sensitive to alternative distributional assumptions. Second, as was discussed earlier, the reduced form empirical probit model used to explain union status and hence to correct for sample censoring has rather ambiguous behavioral underpinnings. Indeed, part of the reason for carrying out this study was to improve our understanding of union status determination. Some progress has been made but more is yet to be done.

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FOOTNOTES

[1]Jobs can revert from union status to nonunion status through an NLRB
supervised "decertification" election. However, these are quite rare.

[2]Evidence from the National Labor Relations Board (1974) and the U.S. Bureau
of Labor Statistics (1975, 1978) indicates that approximately 1.0 percent of
nonunion workers were eligible to vote in NLRB-supervised representation
elections in 1974. Eligibility refers to working in a potential bargaining
unit where an election was held.

[3]Raisian (1981) presents evidence which suggests that union dues offset only
a small portion of the measured union-nonunion wage advantage.

[4]See Farber and Saks (1980) for an analysis which focuses on the preferences
of nonunion workers for union representation.

[5]Using the 1977 Quality of Employment Survey, Farber (1982) derives two sets
of theoretically consistent estimates of union and nonunion wage equations,
first using the so called Mill's Ratio or selectivity regressor technique and
second using the maximum likelihood switching regressor technique. These two
approaches yield vastly different estimates with an average union-nonunion
wage differential ranging from 30 percent to -65 percent. This lack of
agreement between consistent estimates of the same quantity lends support to
Freeman and Medoff's (1981) suggestion that these techniques as they are
currently applied have limited usefulness for the union-nonunion wage
problem. Farber (1982) and Freeman and Medoff (1981) suggest different
explanations for this empirical problem and hence different solutions.

[6]The assumption of unit variance is a normalization required by the
dichotomous nature of VFU in order to fix the scale of γ.

[7] The identity that $\Phi(a) = 1 - \Phi(-a)$ is used in deriving this expression.

[8] See Quinn and Staines (1979) for a detailed description of the survey design.

[9] The question of how workers form their expectations about what unions do is interesting and important. However, it is left to future research. Kochan (1979) presents an analysis of worker perceptions of unions based on the QES.

[10] It is fallacious to argue that since union workers are in fact union workers voluntarily, they desire union representation currently. While it is true that they desired union representation when they took the job in the sense that it was part of a package of job characteristics which was preferred to any other package, the accumulation of seniority can reduce mobility so that a union worker may desire to retain his job but eliminate unionization. This does not mean that the worker will desire to quit.

[11] Other variables which can change over time within a given job are marital status and education. The remaining variables (occupation and region) can also change but are not likely to do so. Changes in these variables are ignored and the current (1977) values are used.

[12] The constrained log-likelihood is -358.0, while the unconstrained log-likelihood is -351.9. The test statistics is $-2(-358.0 - (-351.9)) = 12.2 < 13.7 = \chi^2_{.25}(11)$.

[13] There is redundant information in these tables because there is not a complete analysis of variance. In other words, a complete set of interaction variables was not included. The analysis is done because of the nonlinearity inherent in the relationship between $Z^*\gamma^*$ and $Pr(VFU=1|U=0)$ and the resulting difficulty in interpreting the parameter estimates from probit models.
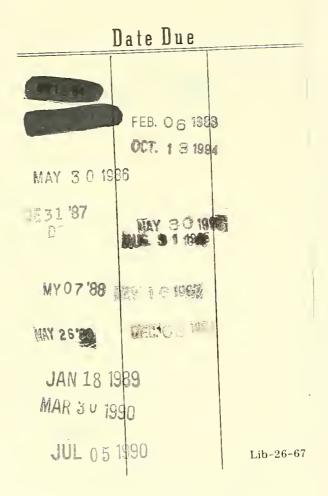
[14]The estimates of α which determine $Pr(U=1)$ are discussed below.

[15]The constrained log-likelihood is -898.2, while the unconstrained log-likelihood is -897.6. The test statistic is $-2(898.2 - (-897.6)) = 1.2 <$
$1.32 = \chi^2_{.25}(1)$.

[16]Evidence from the U.S. Bureau of Labor Statistics (1975, 1978) and from the National Labor Relations Board (1974) indicate that 1.2 percent of nonunion workers in the south were eligible to vote in NLRB-supervised representation elections in 1974. Outside the southern region, only 0.9 percent of nonunion workers were eligible to vote in such elections. (Eligibility refers to working in a potential bargaining unit where an election was held.) Of those workers who voted, 46 percent of workers in the south voted for union representation compared with 50 percent of nonsouthern workers. Similarly, union representation rights were won in 46 percent of the southern elections and in 51 percent of the nonsouthern elections. In both regions approximately .3 percent of nonunion workers were newly organized in 1974 as a result of NLRB-supervised elections.

[17]See Freeman (1981) for a discussion of the effect of unionism on fringe benefits. Medoff and Abraham (1980) investigate the relative importance of seniority for layoff and promotion in union and nonunion establishments.

Date Due