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THE DETERMINATION OF THE UNION STATUS OF WORKERS

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ABSTRACT

The Determination of the Union Status of Workers

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A model of the determination of the union status of workers is developed which incorporates the separate decisions of workers and potential union employers in a framework which recognizes the possibility of an excess supply of workers for existing union jobs. This theoretical framework results in an empirical problem of partial observability because information on union status is not sufficient to determine whether nonunion workers are nonunion because they do not desire union representation or because they were not hired by union employers despite a preference for union representation. The problem is solved by using data from the Quality of Employment Survey which have a unique piece of information on worker preferences which allows identification and estimation of the model.

The empirical results yield some interesting insights into the process of union status determination which cannot be gained from a simple logit or probit analysis of unionization. Chief among these relate to the unionization of nonwhites and southerners. The well-known fact that nonwhites are more likely to be unionized than otherwise equivalent whites is found to be due to a greater demand for union representation on the part of nonwhite workers which is partially offset by a lower probability of being hired by a union employer. The equally well-known lower propensity to be unionized among southern workers is found to be due to a combination of a lower demand for union representation on the part of southern workers and a supply of union jobs which is more constrained relative to demand than in the North.

1. INTRODUCTION

A source of much confusion in the analysis of labor unions regards the process by which the union status of workers is determined. In most cases the union status of individual workers has been modeled as being the result solely of utility maximizing decisions by workers. (See, for example, Ashenfelter and Johnson, 1972; Lee, 1977; and Schmidt and Strauss, 1976). On the other hand, it has been argued that any real effect of unions on compensation or other aspects of employment could be partially or even completely offset by employers' ability to hire better workers. This argument, that union workers might be "better" than observationally equivalent nonunion workers, has led to the recent outpouring of research attempting to measure the "true" impact of unions in the United States.[1] It is clear that union employers must have some control over whom they hire in order for the true effect of unions to be offset by this mechanism, and such employer control is not consistent with the worker choice model of union status. Indeed, it is a major weakness of this literature that either a worker choice model or no explicit model is offered while the implicit reasoning suggests that employers are making relevant decisions. Given the centrality to these analyses of the process by which union status is determined, one must question any conclusions which are drawn in this context.

In this study it is argued that the union status of workers is determined as the result of separate decisions by workers and potential union employers. Workers decide whether they would prefer union or nonunion jobs based on the utilities that these jobs yield to them. At the same time, union employers are deciding which of the workers who want union jobs to hire given that workers differ in their productive characteristics and that these characteristics are compensated differently in the union and nonunion sectors. Essentially union employers are assumed to hire the workers who enable them to produce at minimum cost.

The presumption that union employers have some discretion in hiring results from the likelihood of queues for vacancies in existing union jobs.[2] These queues result from the facts that it is unlikely that dues and initiation fees completely offset the advantages of unionization for all workers and that it is expensive to create new union jobs by organizing nonunion jobs.[3] More fundamentally, the queues result from a distinction, arising from the process of unioniztion, which must be drawn between the union status of workers and the union status of jobs. Nonunion jobs become unionized through organization of the workers who hold them. This is a costly and uncertain process which can involve the holding of an election supervised by the National Labor Relations Board (NLRB).[4] These elections are often preceded by intense and closely monitored campaigns, and they may involve appeals by either or both sides to the NLRB regarding such issues as illegal campaign tactics and determination of the appropriate bargaining unit. However, once the jobs are successfully unionized, their union status is preserved even if the workers who made the investment in organization leave.[5] In addition, new jobs created through expansion of unionized establishments are unionized by definition. Union employers can hire whomever they wish to fill any vacancies, but all new hirees will be unionized.[6] Thus, unless dues or initiation fees are sufficiently large, there will be workers who desire vacancies in existing union jobs but who are not willing to undertake investment in new unionization. For these workers the benefits of unionization are larger than the costs of union membership but smaller than the costs of organizing nonunion jobs. The results are queues for union jobs.

In general, empirical analysis of a model of the determination of the union status of workers of the sort proposed here is hampered by the fact that only the outcome (union status) is observed so that it is impossible to discern whether nonunion workers did not desire union representation or desired union representation but were not selected from the queue by a union employer. Abowd and Farber (1982) carry out with some success an empirical analysis of union status determination which is consistent with a queuing model, but they are hampered by just this partial observability problem. Poirier (1980) presents an econometric approach to identification and estimation of such models. Unfortunately, his technique is heavily dependent on functional form for identification and to date has not proven very useful in applications. More successful are studies which use data from such sources as the Quality of Employment Survey (QES) and surveys of workers participating in NLRB-supervised representation elections to focus on worker preferences for union representation as distinct from actual union status. These include studies by Farber and Saks (1980) and Farber (1982a, 1982b). The drawback of these studies is that they can shed no light on employer selection criteria, and as a result they cannot address the full question of the determination of the union status of workers.

The approach to estimation taken in this study is to utilize data from the QES on both the union status of workers and on the explicit preferences of nonunion workers for union representation. The crucial bit of information is the response elicited from nonunion workers as to whether or not they would vote for union representation on their current job were a secret ballot election to be held. While these data present some problems of their own, it is argued below that they provide enough information to allow identification of the queue and estimation of the full model of union status determination including both worker and employer decision criteria.

In the next section an explicit model of the determination of the union status of workers conditional on the locus of union jobs, incorporating both the worker and potential union employers as decision makers, is developed. Econometrically, the model is bivariate in nature which reflects the fact that there are two decision makers.

In section III the data from the QES are discussed. Particular attention is paid to the interpretation of the crucial question regarding nonunion worker preferences for union representation in the context of the problem of interest here. The data are censored with regard to this variable on the basis of the process of union status determination modeled in the previous section. It is argued that the censored QES information reflects current preferences for union representation while the model suggests that union status is a reflection of preferences for union representation at the time the worker began his current job. It is further argued that the structure of the workers' preference function for union representation does not change over time and that actual preferences will differ over time only to the extent that the measured and unmeasured characteristics of workers or their jobs change. In other words, age or seniority will vary over time and affect worker preferences, but the effect of a given level of age or seniority on preferences will not vary over time. In addition, unmeasured factors such as on-the-job relationships with co-workers or supervisors and unobserved factors which affect compensation can vary over time resulting in changes in preferences. An econometric framework which exploits this fixity of structure while accounting for the censored nature of the data is developed. Section IV contains the empirical analysis of the resulting trivariate discrete data model.

In section V the substantive results are discussed in the context of the theoretical framework derived in section II. Important insights into well known relationships between union status and such characteristics as race, sex, marital status, region, and occupation are gained from the results through the decomposition of these relationships separately into components due to workers and to employers. For example, it is found that the low probability of working on union jobs for southern workers is the result of a combination of a somewhat lower worker demand for union representation combined with a supply of union jobs which is more constrained relative to demand than in the North. On the other hand, the relatively high probability for nonwhite workers of working on union jobs, even after standardizing for education and occupation, is found to be due to a substantially higher demand for union representation among nonwhite workers which is partially offset by a somewhat lower probability of being hired by a union employer.

The final section contains a summary of the results along with a discussion of their implications both with regard to the process of unionization and with regard to analysis of the "true" effects of labor unions.

II. A MODEL OF UNION STATUS DETERMINATION

Assume that firms produce output through a production technology defined by a continuous twice differentiable production function with two inputs: effective labor (E) and capital. Labor is heterogeneous in that different workers embody different skill levels (S). The effective labor services available from any worker in one hour is defined to be a function of the skill level of that worker (A(S)). This function is monotonically increasing in S and has a negative second derivative. In this framework the total effective labor services available to an employer are

$$(1) \quad E = A(S)L$$

where L represents the number of man-hours employed.[7]

It is clear that holding the rate of compensation fixed employers would like to hire the workers with the highest S. However, in a competitive labor market workers with more skill will be able to demand higher compensation. For the purposes of this analysis it is assumed that the nonunion labor market is competitive and that the market equilibrium relationship between skill and total compensation can be described by the function $C_n(S)$ which is monotonically increasing in S. The market equilibrium has the properties both that all workers are employed and that all workers yield identical costs per effective labor unit. The latter quantity is the ratio of $C_n(S)$ to $A(S)$, and in logarithmic terms the equilibrium compensation relationship is defined by

$$(2) \quad \ln C_n(S) = \ln A(S) + \ln R_n$$

where R_n represents the equilibrium value of unit effective labor cost. The function $A(S)$ is assumed to be technologically determined and the distribution of skills is assumed to be exogenously determined so that $C_n(S)$ is free to adjust to these factors.[8]

In the union sector, compensation is determined through the collective bargaining process where market and other factors serve as constraints. The relationship between the skill level and compensation in the union sector can be expressed as the function $C_u(S)$ which is increasing in S. It is beyond the scope of this study to model the determinants of this schedule, though a major factor along with labor market forces is likely to be the internal political processes of the unions.[9] It is sufficient here to specify the skill/compensation relationship facing a given worker.

The decision of an individual worker to desire union representation is based on a comparison of the worker's utilities in the two sectors. The worker will desire employment in the sector which yields the highest level of satisfaction. A straightforward parameterization of the difference between a worker's union and nonunion utilities (M(S)) is

$$(3)\quad M(S) = D(S) + F(S)$$

where D(S) represents the proportional union-nonunion compensation differential ($\ln C_u(S) - \ln C_n(S)$) and F(S) represents the relative nonpecuniary benefits of union representation. If the quantity M(S) is positive, then the individual will desire union representation. Otherwise the individual will not.

The decision process of workers outlined above determines the composition of the labor force willing to work in the union sector. At the same time, union employers must decide which of the workers available to them they wish to hire. Given that they are cost minimizers, union employers will attempt to hire the workers with skill levels which allow the production of effective labor at least-cost, conditional on the skill/compensation relationship determined through the collective bargaining process. In this context a set of factors including the union compensation schedule, the skill distribution of workers who desire union jobs, and the derived demand schedule for union labor define the total quantity of effective labor demanded by union employers and the maximum unit cost for effective labor that union employers will have to pay in order to hire this quantity of effective labor. Denote this maximum cost level by R^*.

Analytically, the cost of a unit of effective labor is a function of the skill level, and it can be expressed as the ratio of compensation per man-hour to effective labor units per man-hour. This is expressed in logarithmic terms as

(4) $\ln R_u(S) = \ln C_u(S) - \ln A(S)$.

Given the assumption that union and nonunion firms face the same set of production technologies, substituting for A(S) from equation (2), and noting that $D(S) = \ln C_u(S) - \ln C_n(S)$ yields the relationship that

(5) $\ln R_u(S) = D(S) + \ln R_n$.

Define the difference between the logarithm of a worker's unit effective labor cost in the union sector and the logarithm of the maximum union cost level as $H(S) = \ln R_u(S) - \ln R^*$. Using equation (5) and rearranging terms, this relationship can be expressed as

(6) $H(S) = D(S) - K$

where K represents $\ln R^* - \ln R_n$. Any worker for whom H(S) is negative (i.e., $R_u(S) < R^*$ so that his unit effective labor cost is less than the maximum) will be hired by a union employer while no union employer will hire a worker for whom H(S) is positive. Another interpretation of this relationship is that a worker will be hired by a union employer only if the proportional union-nonunion compensation differential for that worker (D(S)) is smaller than the proportional difference (K) between the maximum unit effective labor cost in the union sector (R^*) and unit effective labor cost in the nonunion sector (R_n). The latter difference represents the maximum cost disadvantage of union firms in a given labor market and is a function of the relative demand for union labor in that market. It will be called the maximum cost differential.

On the basis of this theoretical framework, an individual will be observed in a union job if and only if $M(S) > 0$ and $H(S) < 0$. In other words, a worker will be unionized if and only if he both wants a union job and is hired by a union employer. The discussion now turns to the specification of empirically tractable forms for M(S) and H(S) in order to derive the econometric model.

Both the decision of union employers and the decision of workers regarding their union status depend on the union-nonunion compensation differential. A first-order approximation to this quantity is

$$(7) \quad D(S) = XB_1 + e_1$$

where X is a vector of observable individual characteristics, B_1 is a vector of parameters, and e_1 accounts for unobserved characteristics which affect D(S). The worker's decision also depends on the nonpecuniary benefit of unionization which has as a first-order approximation a linear function of the same observable factors. This is

$$(8) \quad F(S) = XB_2 + e_2$$

where B_2 is a vector of parameters and e_2 accounts for unobserved characteristics which affect F(S).

The final construct which needs to be specified is the maximum cost differential (K). This is a function of all of the factors which determine the supply of union jobs, including determinants of the extent of organization as well as the factors which determine the union-nonunion compensation differential and the supply of workers to the union and nonunion sectors. This quantity will vary across workers to the extent that they compete in different geographic or occupational labor markets and will be directly related to the relative demand for union labor in the particular market. Given that the relevant variables are subsumed in X, a first-order approximation to K is

$$(9) \quad K = XB_3 + e_3$$

where B_3 is a vector of parameters and e_3 accounts for unobserved factors which affect K.

Using these empirical constructs and substituting into equation (3), the worker criterion function for preferring union representation is

$$(10) \quad M(S) = XG_1 + u_1$$

where $G_1 = B_1 + B_2$ and $u_1 = e_1 + e_2$. Similarly, the criterion function which determines whether or not an individual will be hired by a union employer is derived by substituting into equation (6) as

$$(11) \quad H(S) = XG_2 + u_2$$

where $G_2 = B_1 - B_3$ and $u_2 = e_1 - e_3$.

The unobserved components of the model (e_1, e_2, e_3) can be assumed to be random variables which may be correlated for any particular individual but are distributed independently across different individuals. These random variables have zero mean and covariance matrix[10]

$$(12) \quad V = \begin{bmatrix} V_1^2 & & \\ V_{12} & V_2^2 & \\ V_{13} & V_{23} & V_3^2 \end{bmatrix}.$$

The moments of the distribution of the unobservables in the criterion functions (u_1 and u_2) can be derived in a straightforward fashion from the distribution of the underlying unobservables. The random variables u_1 and u_2 have zero mean and covariance matrix

$$(13) \quad P = \begin{bmatrix} P_1^2 & \\ P_{12} & P_2^2 \end{bmatrix}$$

where $P_1^2 = V_1^2 + V_2^2 + 2V_{12}$, $P_{12} = V_1^2 + V_{12} - V_{13} - V_{23}$, and $P_2^2 = V_1^2 + V_3^2 - 2V_{13}$.

In order to understand how this model can be implemented, it is useful to express formally what can be inferred from data on union status alone. If a worker reports that he is working on a union job, then it can be inferred that at the time he took the job he both desired a union job and was hired by a union employer. However, it cannot be inferred either that he currently desires union representation or that he would currently be hired by a union employer. Consider the worker's preference first. It is possible that a union worker may no longer desire union representation but not be willing to quit his union job and sacrifice the nonportable benefits of seniority in order to take a nonunion job. A similar argument can be made for nonunion workers. Now consider the ability to be hired by a union employer. A nonunion worker may have desired a union job but was not hired by a union employer at the time he started his current job. He may now be able to be hired by a union employer but not be willing to sacrifice his nonunion seniority to take the job. These examples suggest that both worker and employer decisions can change over time and inferences based on the union status of workers must be restricted to preferences of workers and employers at the time of hire.

In the context of the model developed here, the probability that a worker is observed in a union job is the joint probability that he desired a union job at the time of hire ($M_0(S) > 0$) and he was hired by a union employer ($H_0(S) < 0$). The "$_0$" subscript denotes that the relevant quantities are measured at the time of hire. On this basis, the probability of observing a worker on a union job is written in terms of the random variables as

$$(14) \quad \Pr(U=1) = \Pr(u_1 > -X_0G_1, \ u_2 < -X_0G_2) \ .$$

Similarly, the probability of observing a worker in a nonunion job is $1 - \Pr(U=1)$, which can be expressed as

$$(15) \quad \Pr(U=0) = \Pr(u_1 > -X_0G_1, \ u_2 > -X_0G_2) + \Pr(u_1 < -X_0G_1)$$

where the first term represents the probability that the worker desired a union job at the time he took his current job but was not hired by a union employer while the second term represents the probability that the worker did not desire a union job at the time he took his current job. The exogenous variables are time-subscripted to reflect conditions at the start of the job, and the random components (u_1 and u_2), while not subscripted, are considered to be specific to the time of hire. The crucial point to note is that the structural parameters (G_1 and G_2) are not time-subscripted and are assumed to be stable over time.

In order to implement the model a functional form must be selected for the random variables. Therefore, it is assumed that e_1, e_2, and e_3 are distributed as trivariate normal with zero mean and covariance matrix as defined in equation (12). This implies that u_1 and u_2 are distributed as a bivariate normal with zero mean and covariance matrix as defined in equation (13).

It is important to note that not all of the parameters of the model are identified. The likelihood function is specified in terms of the reduced form determinants of M(S) and H(S) rather than the structural variables D(S), F(S), and K. Thus, only the coefficients of this reduced form model (G_1, G_2, and the elements of the covariance matrix (P)) are estimable. It is not possible to recover from these estimates values for B_1, B_2, or B_3 except in the rare instances where it can be plausibly argued that a particular variable is excluded from at least one of the three vectors (i.e., that there are restrictions on the B vectors). Given the difficulty in imposing such restrictions in a convincing fashion, no attempt will be made to recover the structural parameters. Instead, the empirical analysis will focus on the estimates of the reduced form parameters, and these are of sufficient interest in their own right.

Not all of the parameters of the covariance matrix of the reduced form errors (P) are estimable. Due to its discrete choice nature, the model is identified only up to the ratio of the parameter vectors to the standard deviations of their respective errors. For this reason the variances of u_1 and u_2 are normalized to one. Thus, the only element of the covariance matrix which is estimable is the correlation between the reduced form errors (p_{12}). In addition, the probabilities in equations (14) and (15) become standardized normal probabilities.

The model is theoretically identified and can be estimated using data on union status alone where the probability of a worker being unionized is defined as Pr(U=1) in equation (14). However, the two distinct elements in Pr(U=0) in equation (15) highlight the fundamental partial observability problem which stems from not knowing whether nonunion workers are nonunion because they desired a union job but were not hired by a union employer or because they did not desire a union job. Poirier (1980) discusses estimation of partial observability bivariate probit models of this sort and argues that the model is identified and estimable. However, identification relies heavily on nonlinearities in the functional form of the probability distribution, and this is not terribly satisfactory. In addition, some experience with estimation of partial observability models in this context suggests that there are convergence problems and that where convergence is reached the parameters are not estimated with useful precision.[11] In view of these factors, the empirical analysis proceeds using a different approach: additional information on worker preferences, available from the Quality of Employment Survey, is used to aid in the identification and estimation of the model. The discussion turns now to a description of the data and the development of the appropriate econometric framework for estimation of the model utilizing the auxiliary information on worker preferences.

III. THE DATA AND ECONOMETRIC FRAMEWORK

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.[12] The particular sample for use in this study 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 due to the fact that the union status of these workers is determined by a different process than that outlined in the previous section. For example, self-employed workers will not be unionized by definition, while union employment in the construction industry is characterized by hiring halls where the union effectively makes the hiring decisions for employers. 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 the entire sample and the union and nonunion subsamples. The base group for the dichotomous variables consists of white, nonsouthern, unmarried, 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. Unionization is defined as working on a job which is covered by a collective bargaining agreement. This is appropriate in light of the fact that it is collective bargaining as opposed to union membership which alters the employment relationship.

The crucial bits of information for this study are data on the union status of the jobs held by the individuals and the response to the question

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_c	age in years	36.8 (13.1)	38.2 (12.6)	35.9 (13.3)
Sen_c	firm seniority in years	6.90 (7.49)	9.48 (8.18)	5.40 (6.60)
Age_o	$Age_c - Sen_c$	29.9 (10.8)	28.7 (9.28)	30.5 (11.5)
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

asked only of nonunion workers, "If an election were held with secret ballots, would you vote for or against having a union or employee association represent you?". This latter variable, called VFU, is the piece of information which is unique to this data set, and it will serve as the basis for identification of the queue for union jobs. It is interpreted here as the current preference of a worker for union representation on his current job. Thus, it holds all job characteristics fixed, including seniority, except those which the worker expects the union to affect. Fully 37 percent of the nonunion sample answered this question in the affirmative so that there is substantial variation in the response.

It was noted in the previous section that the partial observability problem is the cause of difficulty in identifying and estimating the model strictly from data on union status. The information on VFU can be used to solve this problem in a rather straightforward fashion. Note that the probability that a worker currently desires union representation on his job (Pr(VFU=1)) is a result of the same decision calculus derived in the previous section. This probability is $Pr(M_c(S) > 0)$ where the subscript "c" refers to the current time. In terms of the underlying random variables, the probability that a worker currently desires union representation is

(16) $Pr(VFU=1) = Pr(u_3 > -X_c G_1)$

where X_c represents the exogenous variables measured at the current time and u_3 represents the random component in the worker preference function measured at the current time.[13]

If the data on VFU were available for all workers it would be straightforward to estimate G_1 from a simple probit likelihood function derived from equation (16) under the assumption that u_3 was normally distributed. However, data on VFU are available only for nonunion workers so that the data are censored on the basis of a variable which is obviously related. The standard approach to estimating a censored data model is to specify the censoring process along with the joint stochastic structure of the censored and censoring processes. The model can then be estimated jointly using maximum likelihood techniques. In the case at hand, the censoring process is the model of union status determination derived in section II and expressed probabilistically in equations (14) and (15). Assuming that u_3 is normally distributed with zero mean and using the earlier assumption regarding the joint normality of u_1 and u_2, the implication is that u_1, u_2, and u_3 have a trivariate normal distribution with zero mean and covariance matrix

$$(17) \quad \begin{bmatrix} 1 & p_{12} & p_{13} \\ p_{12} & 1 & p_{23} \\ p_{13} & p_{23} & 1 \end{bmatrix}$$

where the variances are normalized to one as required for identification of this class of discrete data models.

Three distinct events are possible in this framework. The first is that the worker is unionized, in which case there is no information regarding current preferences for union representation. The probability of this event is the probability that at the time the worker started his union job he desired a union job ($M_0(S) > 0$) and he was hired by a union employer ($H_0(S) < 0$). From equation (14) this is

$$(18) \quad Pr(U=1) = Pr(u_1 > -X_0 G_1 ,\ u_2 < -X_0 G_2) .$$

The second event is that the worker is nonunion and currently desires union representation. The probability of this event is derived from equations (15) and (16) as

$$(19)\quad \Pr(U=0,\ VFU=1) = \Pr(u_1 > -X_0G_1,\ u_2 > -X_0G_2,\ u_3 > -X_cG_1) + \Pr(u_1 < -X_0G_1,\ u_3 > -X_cG_1)\ .$$

The first term represents the joint probability that the worker is nonunion because he desired a union job but was not hired and that the worker currently desires a union job. The second term represents the joint probability that the worker is nonunion because he did not desire a union job at the time he started his job and that he currently desires a union job. The final event is that a worker is nonunion and currently does not desire union representation. The probability of this event is derived from equations (15) and (16) as

$$(20)\quad \Pr(U=0,\ VFU=0) = \Pr(u_1 > -X_0G_1,\ u_2 > -X_0G_2,\ u_3 < -X_cG_1) + \Pr(u_1 < -X_0G_1,\ u_3 < -X_cG_1)\ .$$

The first term represents the joint probability that the worker is nonunion because he desired a union job but was not hired by a union employer and that he currently does not desire union representation. The second term represents the joint probability that the worker is nonunion because he did not desire union representation at the time he started his job and that he currently does not desire union representation.

The three probabilities defined in equations (18) through (20) appropriately account for the union status of a particular worker along with his current preference for union representation where it is observed. Identification is clearly aided by the assumption that the parameters of the model which determines worker preferences at the start of the job are the same as the parameters of the model which determines current preferences (G_1). This is a prior theoretical restriction which provides "real" identification of the model and does not rely unduly on the functional form of the probability distribution. It is interesting to note that censored data models are generally estimated in order to obtain consistent estimates of the parameters of the censored process, while in this case the censored data are used to help identify and estimate the parameters of the censoring process.

Although the parameters of the model are fixed over time, the framework allows considerable flexibility in preferences over time. This comes from two sources. The first is that the unobserved components in worker preferences at the start of the job (u_1) and currently (u_3) can and likely do differ while the real possibility of correlation is allowed for. The second source of flexibility comes from the fact that the exogenous variables can change over time. In the empirical work which follows, the major time-varying variables are age and seniority.[14] Overall, the framework allows fluctuations over time in both the measured and unmeasured characteristics of workers and their jobs to have effects on worker preferences for union representation. These effects are consistent with the theoretical framework while at the same time preserving the fundamental identification of the model.

IV. ESTIMATION

The log-likelihood function for the trivariate censored data model is defined using equations (18) through (20) as

$$
\begin{aligned}
(21)\quad L = \sum_{i=1}^{n} \{ & U_i \ln \Pr(U_{1i} > -X_{0i}G_1,\ u_{2i} < -X_{0i}G_2) \\
& + (1-U_i)VFU_i \ln[\Pr(u_{1i} > -X_{0i}G_1,\ u_{2i} > -X_{0i}G_2,\ u_{3i} > -X_{ci}G_1) \\
& \qquad + \Pr(u_{1i} < -X_{0i}G_1,\ u_{3i} > -X_{ci}G_1)] \\
& + (1-U_i)(1-VFU_i)\ln[\Pr(u_{1i} > -X_{0i}G_1,\ u_{2i} > -X_{0i}G_2,\ u_{3i} < -X_{ci}G_1) \\
& \qquad + \Pr(u_{1i} < -X_{0i}G_1,\ u_{3i} < -X_{ci}G_1)]\} ,
\end{aligned}
$$

where i indexes observations. The dichotomous variable U_i equals one for union workers and is zero otherwise, and the dichotomous variable VFU_i equals one if the worker responded to the VFU question affirmatively and is zero otherwise. The likelihood function and its derivatives are composed of univariate, bivariate, and trivariate normal cumulative distribution functions which, while they cannot be evaluated in closed form, can be approximated numerically to the required accuracy. The likelihood function was maximized numerically with respect to G_1, G_2, and the three correlations between u_1, u_2, and u_3 using the algorithm described by Berndt, Hall, Hall, and Hausman (1974). This was a process which consumed large amounts of computational resources but was not marked by any particular difficulty in convergence. Various starting values were used to ensure convergence to a consistent set of parameters.

The maximum likelihood estimates of the parameters are contained in Table 2. The value of the log-likelihood function at the maximum is -897.2. This is compared to a log-likelihood value for a constrained model with two parameters which represent constant probabilities of observing a worker in each of the three possible states of -983.3. This model embodies twenty-eight constraints on the structural model and can be rejected using a likelihood ratio test at any reasonable level of significance. This suggests that the model explains a significant portion of the variation in the data.

Table 2 also contains estimates of a simple univariate probit model of the union status of workers using the same variables as the queuing model. The time dependent variables are measured at the start of the workers' current jobs. These estimates are included simply as an illustration of the conventional approach to estimating models of union status determination, and they are best interpreted as indicative of the partial correlations between the exogenous variables and union status.

Table 2. Estimates of Univariate Union Status Model

	Queue Model G_1	Queue Model G_2	Simple Probit
Constant	.526 (.275)	-1..31 (2.65)	.364 (.181)
NW	.771 (.220)	.148 (1.70)	.316 (.134)
Fe	.252 (.164)	.345 (.780)	-.0269 (.159)
Marr	.118 (.135)	-.290 (.270)	.272 (.136)
Marr*Fe	-.264 (.195)	-.0713 (.702)	-.0571 (.197)
South	-.224 (.105))	.735 (.271)	-.542 (.0965)
Cler	-.444 (.150)	.742 (.702)	-.689 (.140)
Serv	-.148 (.152)	.782 (.290)	-.509 (.138)
Prof & Tech	-.420 (.166)	.506 (.748)	-.506 (.168)
Ed < 12	.0441 (.125)	-.179 (.234)	.0922 (.126)
12 < Ed < 16	-.138 (.119)	.149 (.323)	-.156 (.125)
Ed ≥ 16	.174 (.161)	-.0900 (.444)	.145 (.172)
Age	-.0112 (.00434)	.0146 (.0209)	-.0141 (.00472)
Sen	-.0257 (.0174)	---	---
P_{12}		-.220 (2.88)	
P_{13}		.765 (.287)	
P_{23}		.241 (2.48)	
n	915		915
ln L	-897.2		-546.3

(The numbers in parentheses are asymptotic standard errors.)

It is clear from the estimates in Table 2 that two of the three estimated correlations are estimated very imprecisely. These are the correlation (p_{12}) between the errors in the start-of-job worker preference equation and in the employer selection equation, and the correlation (p_{23}) between the errors in the current worker preference equation and in the employer selection equation. It is obvious that the likelihood function is very flat in these dimensions, which suggests that there is little information in the data regarding whether workers who are more likely on the basis of their unobservable attributes to desire union representation are more or less likely to be hired by union employers. Further evidence for this is that when two versions of the model which constrain these correlations were estimated, the results did not change substantially. The first special case was to impose the constraint that $p_{12} = p_{23}$ so that the correlation between the unobservables affecting worker and employer preferences are time invariant. The maximum log-likelihood value of this model was -897.3 which implies that, using a likelihood ratio test, it is not possible to reject the constraint at any reasonable level of significance. The second special case was to impose the double constraint that $p_{12} = p_{23} = 0$ so that the unobservables affecting worker and employer preferences are uncorrelated. The maximum log-likelihood value for this model was -897.3 which again implies using a likelihood ratio test that the constraint cannot be rejected at any reasonable level of significance. The estimates of the other parameters of the model are virtually unchanged, although the precision with which they are estimated is improved somewhat by the imposition of the constraints. Nonetheless, to be conservative, the discussion of the results will focus on the estimates obtained for unconstrained model and contained in Table 2.

The remaining correlation (p_{13}) between the unobservable factors affecting worker preferences at different points in time is asymptotically significantly greater than zero at conventional levels. This is consistent with the expectation that there are unmeasured attributes of jobs and workers which affect preferences for union representation and which persist over time.

V. ANALYSIS OF RESULTS

It was concluded that in the absence of constraints imposed on _a priori_ grounds it is impossible to use the parameter estimates in Table 2 to examine the separate variation in the union-nonunion compensation differential, the nonpecuniary costs and benefits of unionization, and the maximum unit cost per effective labor unit in the union sector. However, it is clear that the estimates of G_1 reflect variation in worker preferences for unionization. In particular, the probability that a worker desires union representation is $Pr(u_1 > -XG_1)$ so that a positive coefficient on a variable in XG_1 implies that workers with higher values of that variable are more likely to desire union representation. Similarly, the estimates of G_2 reflect variation in the propensity of union employers to hire particular workers. The probability that a given worker will be hired by a union employer is $Pr(u_2 < -XG_2)$ so that a positive coefficient on a variable in XG_2 implies that workers with higher values of that variable are less likely to find union employment.

The estimates of the simple probit model of union status determination contained in Table 2 highlights a number of interesting empirical relationships. Chief among these are that nonwhites and married workers are more likely while southern workers less likely to be union workers. In addition, older workers are less likely to be unionized while blue collar workers are significantly more likely to be unionized than any of the other three occupational groupings. These results, while typical, are not easily interpreted with regard to the behavior of workers or employers. For example, the fact that southern workers are less likely to be unionized does not provide any information regarding the extent to which this is a result of less preference for union representation on the part of workers as opposed to a relative lack of supply of union jobs.

The estimates of the queuing model of union status determination can be used to resolve these behavioral issues. The important quantities are the probability that a worker desires union representation (Pr(DES=1)), the probability that a worker will be hired by a union employer (Pr(HIRE=1)), the probability that a worker who desires union representation will be hired by a union employer (Pr(HIRE=1|DES=1), and the probability that a worker is unionized (Pr(U=1)). These probabilities are easily constructed from the parameter estimates as

$$\begin{aligned} \Pr(\text{DES}=1) &= \Pr(u_1 > -XG_1) \\ \Pr(\text{HIRE}=1) &= \Pr(U_2 < -XG_2) \\ \Pr(U=1) &= \Pr(\text{DES} = 1, \text{HIRE} = 1) \\ &= \Pr(u_1 > -X_0G_1,\ u_2 < -X_0G_2), \text{ and} \qquad (22) \\ \Pr(\text{HIRE}=1|\text{DES}=1) &= \frac{\Pr(U=1)}{\Pr(\text{DES}=1)} \end{aligned}$$

where the last relationship follows from application of Bayes' Law. The quantity Pr(HIRE=1|DES=1) is particularly interesting in that it reflects (inversely) the extent to which there are queues for vacancies in existing union jobs.

The parameter estimates will be discussed considering the effect of one variable at a time for a 30 year old worker in the base group consisting of white single male blue collar nonsouthern workers with 12 years of education and zero seniority. Based on the parameter estimates contained in the first two columns of Table 3, workers with these characteristics have $Pr(DES=1) = .58$, $Pr(HIRE=1) = .81$, and $Pr(HIRE=1|DES=1) = .85$. The resultant $Pr(U=1) = .49$. The fact that $Pr(HIRE=1|DES=1)$ is substantially less than one suggests that workers cannot necessarily translate a preference for union representation into a union job so that there are queues for vacancies in existing jobs.

Nonwhites are significantly more likely than other workers in the base group to desire union representation ($Pr(DES=1) = .83$).[15] At the same time, their probability of being hired by a union employer is not significantly different than that of whites ($Pr(HIRE=1) = .77$). However, the probability that a nonwhite worker will be hired by a union employer conditional on desiring a union job is smaller than for white workers ($Pr(HIRE=1|DES=1) = .79$). In other words, since the higher probability of desiring a union job for nonwhites is not accommodated by union employers through a higher probability of being hired, nonwhites who desire union jobs have a lower probability than whites who desire union jobs to realize that desire. Simply put, nonwhites are overrepresented in the queue for union jobs. Overall, the resultant $Pr(U=1) = .66$ and the observed positive correlation between nonwhite and unionization is the result of the greater desire on the part of nonwhites for union representation which is partially offset by their lesser success in being hired for union jobs. This result must be interpreted with some care due to the imprecision with which the coefficient on nonwhite in the employer selection equation is estimated.

Southern workers are significantly less likely than other workers in the base group to desire union representation ($Pr(DES=1) = .49$). In addition, they are significantly less likely to be hired by union employers ($Pr(HIRE=1) = .56$). More importantly, the probability of a southern worker being hired by a union employer conditional on desiring a union job is substantially smaller than that for other workers in the base group ($Pr(HIRE=1|DES=1) = .63$). These results suggest that the relatively low probability of unionization in the south ($Pr(U=1) = .31$) is due both to less demand for union services by workers and to more severe supply constraints on union jobs relative to even this lower demand. In other words, in spite of the relatively low overall demand for unionization in the South, the queues for vacancies in existing union jobs are long relative to those outside the South. This may be due to a social and legal climate (typified by Right-to-Work Laws common in the South) which makes union organizing and administration in the South more difficult and expensive.

It is clear from the simple probit results that the three occupational groupings including clerical, service, and professional and technical workers are each less likely than blue collar workers to be unionized. While no distinction can be drawn among the first three groups based on the probit model, some interesting distinctions can be drawn using the queuing model. In order to facilitate this discussion, Table 3 contains the values of the relevant probabilities computed for workers in the base group for the various occupations. For example, clerical workers and professional and technical workers are significantly less likely than blue collar workers to desire union representation, but this is not true of service workers. On the other hand, service workers are significantly less likely than blue collar workers to be hired by a union employer both unconditionally and conditionally on desiring a union job. This is true to a lesser extent for clerical workers, but it is not true of professional and technical workers. The latter group

Table 3. Predicted Probabilities for Various Occupations

	Pr(DES=1)	Pr(HIRE=1)	Pr(HIRE=1|DES=1)	Pr(U=1)
Blue collar	.58	.81	.85	.49
Clerical	.40	.55	.64	.26
Service	.50	.50	.57	.29
Professional and Technical	.41	.65	.72	.30

These probabilities are computed from equations (27) using the estimates in the first two columns of Table 3. The workers are 30 year old white single male nonsouthern workers with 12 years education and zero seniority.

Table 4. Predicted Probabilities for Various Marital Status/Sex Categories

	Pr(DES=1)	Pr(HIRE=1)	Pr(HIRE=1|DES=1)	Pr(U=1)
Males				
Single	.58	.81	.85	.49
Married	.62	.88	.91	.56
Females				
Single	.67	.70	.75	.50
Married	.62	.81	.85	.53

These probabilities are computed from equations (27) using the estimates in the first two columns of Table 3. The workers are 30 year old blue collar white nonsouthern workers with 12 years education and zero seniority.

has an unconditional probability of being hired by a union employer which does not differ significantly from the base group. What these occupational distinctions suggest is that the low observed extent of unionization among professional and technical workers reflects largely a lower demand for union representation among this group. At the other extreme, service workers are less unionized largely as a result of inability to be hired by union employers in spite of a demand for union jobs. This may reflect higher costs of creating new union jobs as a result of market conditions or employer resistance. The low extent of unionization among clerical workers is due to a mixture of the two factors.

The simple probit model suggests both that married workers are more likely to be unionized and that there is no significant sex differential. Once again, interesting distinctions arise using the queuing model. In order to facilitate this discussion, Table 4 contains the relevant probabilities for the various marital status and sex combinations. These probabilities must be interpreted with caution due to the relatively imprecise parameter estimates. Nonetheless, the results suggest that married males are more likely to be hired by a union employer than single males both unconditionally and conditionally on desiring a union job. There is somewhat weaker evidence that married males are more likely to desire union representation than are single males. These two factors in combination explain the observed higher propensity for unionization among married males.

Analysis of the sex pattern in the process of union status determination is more complicated. It is important to note, given the well known occupational segregation of women, that the effects of sex measured here are derived after controlling for occupation. In this context, single females are significantly more likely than single males to desire union representation while they are weakly less likely to be hired by a union employer both unconditionally and conditionally on desiring union representation. These factors offset each other, resulting in the zero net effect of sex on union status of single workers noted in the simple probit. The evidence suggests that married females are not significantly different from single females in either their desire for union representation or in their probability of being hired by a union employer.[16] Finally, married females are virtually identical to married males in their desire for union representation, while married females are weakly less likely than married males to be hired by a union employer.[17]

Older workers are significantly less likely to desire union representation to the extent that Pr(DES=1) falls from .58 for a 20 year old member of the base group to .49 for a 50 year old member. However, the probability of being hired for a union job is not significantly related to age. The result is that the probability of being hired by a union employer conditional on desiring union representation falls slightly with age. Thus, the inverse relationship noted between age and union status in the simple probit model is largely due to an inverse relationship between age and the demand for union representation. On its face, this result seems to contradict the notion that unions provide more fringe benefits such as pensions, which ought to be valued more by older workers than do nonunion employers.[18] However, this result is consistent with evidence presented by Farber and Saks (1980), based on an entirely different data set, which shows a similar inverse relationship between age and worker preferences for union representation. Again, caution is necessary in interpreting these results due to the imprecision with which the coefficient of age in the employer preference function is estimated.

Nonunion seniority can affect only the desire for union representation in this model. Workers with more nonunion seniority are significantly less likely to desire union representation than are workers with less nonunion seniority. To illustrate this, the probability that a worker in the base group with no nonunion seniority at age 40 desires union representation is .53, while the same probability for an otherwise equivalent worker with 10 years seniority is .43. Note that the result refers to the effect of seniority on the desire for union representation on the current job so that it is not caused by a reluctance of high seniority nonunion workers to quit their jobs in order to take union jobs.

The remaining set of variables relates to the educational attainment of workers. No systematic pattern emerges from the estimates regarding the relationship between education and the process by which the union status of workers is determined.

VI. SUMMARY AND CONCLUSIONS

In this study a model of the determination of the union status of workers was developed which differs substantially from the standard worker choice model. The decisions of both workers and potential union employers were incorporated in the model, recognizing the possibility of an excess supply of workers for existing union jobs. In this context, workers make explicit decisions regarding their desire for union representation which do not necessarily result in employment on a union job. Only if the worker is hired by a union employer out of the queue of workers who desire union representation will the worker's preference actually result in unionization. This theoretical framework results in an empirical problem of partial observability because data on union status are not sufficient to determine whether nonunion workers are nonunion because they do not desire union representation or because they were not hired by a union employer despite their preference for such a job.

In order to solve this problem without relying on distributional assumptions for identification, a rather unique data set from the Quality of Employment Survey (QES) was used. These data contain information which, for nonunion workers, provides information on their current preferences for union representation. Using these data, a trivariate econometric model which accounts for the censored nature of these data as well as the union status of workers was derived explicitly from the theoretical framework. This empirical specification embodies the separate decisions of workers and potential union employers regarding the determination of the union status of workers.

The empirical results yield some interesting insights into the process of union status determination which cannot be learned from a simple probit or logit analysis of unionization. Chief among these relate to unionization of nonwhites and southerners. The well-known fact that nonwhites are more likely to be unionized compared to otherwise equivalent whites was found to be the result of a greater preference for union representation which is tempered somewhat by a lower probability of being hired by a union employer conditional on desiring union representation. The equally well-known lower propensity to be unionized among southern workers was found to be due to a combination of a lower demand for union representation on the part of workers and a supply of unionized jobs which is more constrained than outside the South relative to demand. The longer queues in the South for vacancies in existing union jobs implied by the latter result are attributed to higher costs of organization and administration of labor unions in the South. Other dimensions along which the results interpreted in the context of the model yield behavioral insights include occupational status, sex and marital status, and age.

The model and estimates presented here have important implications for measuring the true effect of unions (as opposed to the union-nonunion differential) on such quantities as wages, turnover, and productivity. The wealth of studies (surveyed and critiqued by Freeman and Medoff (1981)) that attempt to estimate this true effect rely on econometric techniques which posit that union status is determined through a single equation/single decision-maker process. To the extent that this process is inadequately modeled, the estimates of the true effects of unions which rely on them will be misleading.

To be more explicit, consider the example of the widely used Mills' ratio technique presented by Heckman (1979) to correct for sample selection bias. This technique proceeds on the assumption that the log of wages, for example, is distributed normally and that union status can be modeled as determined by a simple probit. Under the assumption of joint normality of the errors, estimates can be derived for the mean of the error(s) in the wage equation(s) conditional on union status as a function of the reduced form probit estimates on union status. These estimated conditional means are the basis of the correction of the union-nonunion differential to yield estimates of the true effect of unions. This correction is crucially dependent on a range of assumptions, not the least of which is that union status can be modeled correctly as a simple univariate probit. If this particular assumption fails, then the conditional means of the wage functions will have a different form from that derived from a simple probit so that the correction will be unreliable.

It should be clear from the results of this study that the determination of union status cannot be modeled adequately as a simple probit and that an approach to estimating the true effects of unions consistent with the model developed here would be preferable. Unfortunately, the data problems outlined above make implementation of this model for such purposes difficult. As far as can be determined, only the QES has the data required to estimate the model, and previous experience with estimating union and nonunion wage equations using these data is not typical of similar experience with more widely used data sources such as the Current Population Survey or the Panel Study of Income Dynamics.[19] A topic for future research is the development of techniques for estimating models of the sort presented here which use data solely on union status and which do not rely to an undue extent on the functional form of the error distribution for identification.

NOTES

1. See Freeman and Medoff (1981) for an interesting summary of this literature as well as a critique from a unique perspective.
2. This analysis is not applicable to industries, such as construction, where hiring is controlled by the union through a hiring hall. Workers in such industries are excluded from both the theoretical and empirical analyses throughout.
3. Raisian (1981) investigates the issues of the magnitude of union dues relative to the union-nonunion wage differential.
4. The particular set of institutions described here refer to private sector nonagricultural and nonmanagerial workers in the United States who are covered by the National Labor Relations Act (NLRA). Organization of workers not covered by the NLRA proceeds along different, but equally costly and uncertain lines.
5. It is possible for union jobs to revert to nonunion status through an NLRB-supervised decertification election. However, these are relatively rare and can safely be ignored in this analysis. For example, according to the NLRB (1979), during fiscal 1979 7266 certification elections involving 538,404 workers were officially decided while only 777 decertification elections involving 39,538 workers were officially decided.
6. In states with Right-to-Work laws, new hirees cannot be forced to join the union or pay dues, but they do share in any benefits of unionization. This issue will be raised again in interpreting the empirical results.
7. Johnson (1970) discusses this production framework in more detail.

8. The foregoing analysis is considerably complicated by recognition that an individual's skill level is determined at least in part through investment decisions made by the individual on the basis of $C_n(S)$. However, explicit consideration of this factor is beyond the scope of this study, and the current assumptions are sufficient for the problem at hand.

9. See the Webbs (1920), Ross (1948), and Dunlop (1950) for early discussions of market and political forces in the determination of union bargaining goals. Farber (1978) develops and estimates a simple voting model of union wage determination.

10. The assumption of a zero mean is neutral due to the presence of constant terms in the parameter vectors which capture the mean unobserved effect.

11. These models have been estimated in this context using samples from the Panel Study of Income Dynamics in excess of 1500 observations and from the Current Population Survey in excess of 19,000 observations.

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

13. A more cumbersome notation would define u_3 as u_{1c} and u_1 and u_2 in equations (14) and (15) as u_{10} and u_{20} respectively.

14. Other variables, such as marital status, which can change over time are assumed not to vary due to lack of information on such variation.

15. The term "significance" is used in this discussion to indicate coefficient estimates which are significantly different from zero at conventional levels. This does not translate directly into statements about significantly different probabilities due to the nonlinearity of the probability function.

16. The sum of the married and the married*female coefficients in G_1 is -.146 with an asymptotic standard error of .140. The sum of the married

and the married*female coefficients in G_2 is -.36 with an asymptotic standard error of .55.

17. The former relationship comes from the sum of the coefficients of female and female*married in G_1. This is -.012 with an asymptotic standard error of .130. The latter relationship is based on the sum of the married and the married*female coefficients in G_1, which is .276 with an asymptotic standard error of .268.

18. See Freeman (1981) for an empirical analysis of the relationship between unionization and fringe benefits.

19. See Farber (1982a).

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