Job Queues and the Union Status of Workers

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JOHN M. ABOWD and HENRY S. FARBER*

This paper develops a model of the determination of the union status of
workers that allows for the possibility of queuing for union jobs. The em-
pirical results derived, using a sample from the University of Michigan *Panel
Study of Income Dynamics*, are supportive of the queuing hypothesis. The
no-queue model can be rejected using a likelihood-ratio test. This suggests
that a simple probit or logit model for union status is misspecified because it
is not based on any consistent behavioral theory. An important implication
of the model is that because most new entrants to the labor market prefer un-
ion jobs but cannot get them and because accrual of nonunion seniority
makes workers progressively less likely to desire union jobs, the union status
of workers is largely determined by their success in being selected from the
queue early in their working life.

M OST recent empirical research on the
union status of individuals is based
on an underlying model that assumes that
workers determine the union status of jobs
through individual, utility-maximizing
decisions.[1] In this view, workers compare

[1]See, for example, Peter Schmidt and Robert P.
Strauss, "The Effect of Unions on Earnings and Earn-
ings on Unions," *International Economic Review*,
Vol. 17, No. 1 (February 1976), pp. 204–12; Lung-Fei
Lee, "Unionism and Wage Rates: A Simultaneous
Equations Model with Qualitative and Limited De-
pendent Variables," *International Economic Review*,
Vol. 19, No. 2 (June 1978), pp. 415–33; Lung-Fei Lee,
"On the Estimation of Probit Choice Model with
Censored Dependent Variables and Amemiya's Prin-

the utility of union and nonunion jobs, and
the observed union status is the result. Ash-
enfelter and Johnson and Leigh recognize
that this individual utility comparison de-
termines only the demand for union services
and that there is a supply function for union
services that is derived from some model of
employer behavior and the cost of union-
ization of jobs.[2] The demand for and supply

ciple" Discussion Paper 78–99, Center for Economic
Research, Department of Economics, University of
Minnesota, 1978; and James J. Heckman and George
R. Neumann, "Union Wage Differentials and the
Decision to Join Unions" mimeo, University of Chi-
cago, 1977.

[2]Orley Ashenfelter and George Johnson, "Union-
ism, Relative Wages, and Labor Quality in U.S. Man-
ufacturing Industries," *International Economic
Review*, Vol. 13, No. 3 (October 1972), pp. 488–507.
See also Duane E. Leigh, "An Analysis of the Inter-
relation Between Unions, Race, and Wage and Non-
wage Compensation," final report submitted to the
Employment and Training Administration, U.S.

of union services are implicitly assumed to equilibrate through a market mechanism, yielding an equilibrium relationship that determines the observed union status of individuals.

It is argued here that while an explicit market mechanism may be operational in the determination of the size of the union sector—where on the margin the costs of unionization of an additional job will equal the benefits of unionization of that job—the allocation of *existing* union jobs is not mediated through such a market. The major costs of unionization are incurred when the bargaining unit is organized, and the benefits of unionization are generally not capitalized in initiation fees or recovered through dues payments. Since, in general, union jobs are not sold and the benefits of the union jobs are fixed through collective bargaining, there will be excess demand for job vacancies in existing union establishments.

These considerations suggest that there may be a queue of workers for union jobs. In most industries the employer has discretion in filling vacancies, and a profit-maximizing employer faced with a queue of workers and the fixed compensation rules imposed by the collective bargaining process will systematically select workers from the queue so as to minimize production costs.[3] Thus, the observed union status of individuals is the product of distinct decisions systematically made by both the worker and the potential employer.

The existence of a queue raises serious questions about the interpretation of the results of the earlier studies of the union status of workers. Since a simple unionization function does not necessarily reflect a market in equilibrium, it is not clear whether it is a demand function, a supply function, or some hybrid of the two. A worker's union status is determined by both a desire for a union job and the employer's selection criteria. An individual will be working on a union job if that individual both wants a union job and is selected from the queue. If an individual is not working on a union job, however, the observer can not know whether that is because the individual did not want a union job or whether he or she wanted a union job but could not get one.

The Composition of the Queue for Union Jobs

Assume that individual i has a utility function of the form

$$(1) \qquad V_i = V(w_i, U_i \phi_i)$$

where w_i represents the wage of individual i, U_i is a zero-one dummy variable denoting the union status of individual i, and ϕ_i is a factor that represents the pecuniary and nonpecuniary costs and benefits of unionization that are not included in earnings. It is assumed that $\frac{\partial V}{\partial w} > 0$ and $\frac{\partial V}{\partial U \phi} < 0$.

Intuitively, the latter inequality implies that individuals with positive ϕ's find that the costs of union membership outweigh the benefits net of any wage differential. Conversely, individuals with negative ϕ's find that the benefits of union membership outweigh the costs net of any wage differential. If individual i is a union member, utility is

$$(2) \qquad V_{ui} = V(w_{ui}, \phi_i)$$

while if the worker is not a union member, utility is

$$(3) \qquad V_{ni} = V(w_{ni}, 0),$$

where w_{ui} and w_{ni} represent the union and nonunion wages, respectively, of individual i.

It is assumed that the utility function has the particular form

$$(4) \qquad V_i = \text{ln} w_i - U_i \phi_i.$$

Thus, the difference between the utility to individual i of a union job and of a nonunion job is[4]

[3]In industries such as construction, where there are strong union hiring halls, employers do not exert such control over hiring.

[4]Our specification assumes that the net benefit (or cost) of a unionized job is independent of the union that organized the job being considered.

(5) $y_{1i} = \ell n w_{ui} - \ell n w_{ni} - \phi_i.$

Individual i will prefer a union job if $y_{1i} > 0$ and a nonunion job if $y_{1i} < 0.$[5]

Assume that each individual has union and nonunion earnings functions of the form

(6) $\ell n w_{ui} = x_{ui} \beta_u + \varepsilon_{ui}$ and

(7) $\ell n w_{ni} = x_{ni} \beta_n + \varepsilon_{ni}$

where x_{ui} and x_{ni} are row vectors of individual and labor market characteristics pertaining to individual i and ε_{ui} and ε_{ni} represent random errors. The parameter vectors β_u and β_n are assumed to be constant across individuals and to be given exogenously.[6]

The union-nonunion wage differential facing any worker is

(8) $\Delta \ell n w_i = \ell n w_{ui} - \ell n w_{ni}$

$= x_{wi} (\beta_u - \beta_n) + \varepsilon_{wi}$

where $\varepsilon_{wi} = \varepsilon_{ui} - \varepsilon_{ni}$, x_{wi} represents a vector of all the variables in x_{ui} or x_{ni}, and $\beta_u - \beta_n$ is the vector of coefficients for elements that correspond to the appropriate variables in x_{wi}.

The variable ϕ_i represents the pecuniary and nonpecuniary costs and benefits of union membership net of any wage differential. The obvious elements of the costs of union membership are dues payments and initiation fees. The pecuniary benefits of unionization other than higher wages accrue as higher levels of fringe benefits, such as insurance and pension plans, in union jobs relative to nonunion jobs. While data concerning relative provision of fringe benefits for individuals in union and nonunion jobs are rare, the U.S. Bureau of Labor Statistics reports that in 1972 "pay for time worked as a percent of total compensation was about 6 percentage points lower for union than for nonunion workers in all situations."[7] Thus, even if unions confer no wage advantage on their members, on average union members receive more fringe benefits.[8]

The nonpecuniary costs and benefits of union membership flow from union-induced changes in job and workplace characteristics. Examples of such changes include changes in the time spent at work actually working and the introduction of formal grievance-handling procedures. Workers may also realize nonpecuniary costs or benefits simply from holding a union job as opposed to a nonunion job. These aspects of union employment are notoriously difficult to evaluate, but one attempt to attach a value to the impact of unionization on the way time is spent on the job was made by Duncan and Stafford.[9] They concluded that differences in the structure of the use of time at work offset approximately one-third of the measured union-nonunion wage differential.

Although these aspects of union membership (ϕ_i) are not measured directly, they are assumed to vary across individuals as a linear function of individual and market

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[5]Up to this point in the analysis, union membership and employment in a union job have been used interchangeably. They are not in fact identical. Union membership denotes actual membership in the organization while employment in a union job denotes simply that the job is covered by a collective bargaining agreement. Unless there is an explicit agreement in the contract concerning a membership requirement, an individual may work on a job that is covered by a collective bargaining contract without actually becoming a member of the union. However, such individuals are "covered" by the contract in the sense that their terms of employment are those agreed to by the union and the employer. In light of this, it is assumed for the purposes of this study that union membership and employment in a union job are identical.

When we turn to the empirical work below, coverage by a collective bargaining agreement will be the measure of union "membership" used. Implicit in this convention is the assumption that an individual receives no utility or disutility from actually belonging to a union that differs from that received by a nonmember who works on a "covered" job.

[6]The assumption of exogenous wage structures is not strictly accurate in the context of a broader view of union behavior. The union wage structure is clearly a result of the collective bargaining process and it will feed back upon the nonunion wage structure, thereby modifying it as well.

[7]U.S. Bureau of Labor Statistics, *Employee Compensation in the Private Nonfarm Economy,* Bulletin 1873 (Washington, D.C.: G.P.O., 1975), p. 4.

[8]Richard B. Freeman, "The Effect of Unionism on Fringe Benefits," *Industrial and Labor Relations Review,* Vol. 34, No. 4 (July 1981), pp. 489–509.

[9]Greg J. Duncan and Frank P. Stafford, "Do Union Members Receive Compensating Wage Differentials?" *American Economic Review,* Vol. 70, No. 3 (June 1980), pp. 355–371.

characteristics $(x_{\phi i})$ with a random component, $\varepsilon_{\phi i}$.

Formally,

$$(9) \qquad \phi_i = x_{\phi i}\,\beta_\phi + \varepsilon_{\phi i}$$

where β_ϕ is a vector of parameters. Substitution of Equations 8 and 9 into Equation 5 yields

$$(10)\, Y_{1i} = x_{wi}\,(\beta_u - \beta_n) + x_{\phi i}\,\beta_\phi + \varepsilon_{1i} > 0$$

as the condition for individual i wanting a union job where $\varepsilon_{1i} = \varepsilon_{wi} + \varepsilon_{\phi i}$.

It is clear from Equation 10 that the effect of individual characteristics on the desire for a union job is only in part a function of $\beta_u - \beta_n$, which represents the difference between the pecuniary returns to individual characteristics in the union and nonunion sectors. It is argued here that unions standardize wage rates, resulting in a reduction in skill differentials within the union sector. This notion can be traced back to at least the Webbs.[10] Recent empirical work by Bloch and Kuskin and by Freeman supports this hypothesis.[11]

The implication of this standardization of rates within the union sector is that the propensity of an individual to desire a union job is inversely related to his or her skill level. In addition, since there seems to be less wage discrimination against blacks in the union sector, blacks will be more likely to desire union jobs.[12] These hypothesized relationships may be modi-

fied by any systematic relationships between the characteristics of workers and the nonwage costs and benefits of unionization embodied in ϕ_i.

Standardization of wage rates by unions implies that worker skill characteristics will be negatively related to the desire for a union job; however, there are four potential relationships between worker characteristics and nonwage aspects of unionism that might reverse this prediction. First, initiation fees and annual dues could be structured to favor more senior workers over less senior workers.[13] Since there do not appear to be any negative relationships between actual union dues and seniority or skill groups, this effect is unlikely to reverse the standardization implication. Second, the demand for grievance procedures to protect union members' job rights should differ systematically across groups. We expect minorities to benefit relatively more from grievance procedures. Third, to the extent that formalization of the work environment and job progressions within a union firm benefit some skill groups more than others, the insulation of high-skill jobs from some forms of within-firm competition could mitigate the standardization implications. Finally, fringe benefits—in particular, the division of compensation between wages and pensions—are a nonwage factor that could reverse the standardization implications. Since older workers benefit more from pensions than younger workers, the desire to remain in a union job could increase with age because of the anticipated pension benefits, despite union wage benefits that may favor younger workers.

[10]Sidney Webb and Beatrice Webb, *Industrial Democracy* (New York: August M. Kelley, 1965). Reprint of 1920 edition.

[11]Farrell E. Bloch and Mark S. Kuskin, "Wage Determination in the Union and Nonunion Sectors," *Industrial and Labor Relations Review*, Vol. 31, No. 2 (January 1978), pp. 183–92 and Richard B. Freeman, "Unionism and the Dispersion of Wages," *Industrial and Labor Relations Review*, Vol. 34, No. 1 (October 1980), pp. 3–23.

A wage-function analysis using the same data we analyze in this paper shows that our data also exhibit the standardization of wage rates within the union sector. The union-wage equation is flatter than the nonunion equation in virtually every dimension. These results are available in an appendix, which will be supplied by the authors on request.

[12]Orley Ashenfelter, "Racial Discrimination and Trade Unionism," *Journal of Political Economy*, Vol. 80, No. 3 (May/June 1972), pp. 435–64.

[13]Initiation fees could, in principle, capitalize some of the advantage accruing to younger, less skilled workers. A large initiation fee could serve as a transfer from the new union entrants, whose gains are largest, to the older union members, whose wage growth may be slower in the union than in the nonunion sector. As a nonunion worker ages and becomes more skilled in the nonunion sector, the size of the initiation fee the worker would be willing to pay to gain access to the union declines. To the extent that initiation fees are larger than annual dues payments, this effect strengthens the prediction that standardization of wages leads to an inverse relationship between skill measures and the desire for a union job.

Due to the overlap in variables that will be included in x_w and x_ϕ, it will not be possible to estimate β_u, β_n, and β_ϕ from data on union status.[14] Thus, the condition for desiring a union job in Equation 10 can be simplified to

$$(11) \qquad y_{1i} = x_{1i}\beta_1 + \varepsilon_{1i} > 0$$

where x_{1i} represents a vector of individual characteristics that includes all the elements of x_{wi} and $x_{\phi i}$. The parameter vector β_1 includes β_u, β_n, and β_ϕ and its elements can be interpreted as the overall wage and nonwage effect of the relevant characteristics on the desire for a union job.

If there is no queue and workers can thus translate their desire for a union job into the fact of a union job, then Equation 11 determines the union status of workers directly. Assuming that ε_{1i} is distributed as a standard normal random variable, there is a probit specification for the probability of observing union status given by:

$$(12) \qquad Pr[U_i = 1] = Pr[y_{1i} > 0]$$
$$= Pr[\varepsilon_{1i} > -x_{1i}\beta_1]$$

where U_i is a dichotomous variable that equals one if individual i works on a union job and zero otherwise.

If, on the other hand, there is a queue for union jobs, workers cannot necessarily translate their desire for a union job into the fact of a union job, and Equation 11 merely determines whether or not an individual is in the queue.[15] More formally,

$$(13) \qquad Pr[IQ_i = 1] = Pr[y_{1i} > 0]$$
$$= Pr[\varepsilon_{1i} > -x_{1i}\beta_1]$$

where IQ_i is a dichotomous variable that equals one if individual i is in the queue and zero otherwise. Unlike union status, however, queuing has no observable analogue. Therefore Equation 13 cannot be empirically implemented by itself; such imple-

mentation requires a model of the selection process used by employers.

The Process of Selection from the Queue

Conditional on the composition of the queue and on the earnings structure imposed through collective bargaining, it is reasonable to assume that union employers systematically select workers from the queue so as to minimize the costs of production. The model developed in this section implies that in order to minimize costs union employers attempt to hire the more highly skilled workers from the queue.

It was argued previously that the queue is likely to be composed of the less skilled workers due to the fact that the marginal wage return to an additional unit of skill is lower in the union sector than in the nonunion sector. The same union and nonunion wage functions can be interpreted from the employers' point of view to imply that the marginal cost of an additional unit of skill, holding number of workers fixed, is lower in the union sector than in the nonunion sector.

Given this interpretation of the relative shapes of the earnings functions, we assume: (1) that firms hire workers of various skill levels until their marginal product is no greater than their wage rate; (2) that the production functions exhibit declining marginal productivity for every skill level; and (3) that skill and number of workers are substitutes in production. It can be shown that in firms that are identical except for facing different prices for human capital, the firms facing a lower price for human capital will hire a smaller work force with a higher average human capital level in order to produce a given output. In other words, they will substitute skill for numbers.[16] Since it is hypothesized that unions lower the price of human capital, employers of union labor will want to hire a more skilled work force

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[14]It is theoretically possible to identify and estimate β_u, β_n, and β_ϕ in a known regime-switching regression model consisting of two wage equations and a union-status model. However, the implementation of such a model in a queuing context is quite complicated, and it is not attempted here.

[15]It is assumed that queuing is costless and that individuals can be working on nonunion jobs while waiting to be selected from the queue for union jobs.

[16]George E. Johnson, "The Demand for Labor by Educational Category," *Southern Economic Journal*, Vol. 37, No. 2 (October 1970), pp. 190–204. See Ashenfelter and Johnson, "Unionism, Relative Wages, and Labor Quality," for an analysis of the role of skill in the production process.

than employers in the nonunion sector. Thus, more skilled workers will have a higher probability of being chosen from the queue.

More formally, let

$$(14) \qquad y_{2i} = x_{2i}\,\beta_2 + \varepsilon_{2i}$$

where x_{2i} is a vector of individual characteristics, β_2 is a parameter vector, and ε_{2i} is a random error. For any individual i who is in the queue the probability of being chosen from the queue is

$$(15)\; Pr[CFQ_i = 1|IQ_i = 1] = Pr[y_{2i} > 0]$$
$$= Pr[\,\varepsilon_{2i} > - x_{2i}\,\beta_2\,]$$

where CFQ_i is a dichotomous variable that equals one if the individual is chosen from the queue and zero otherwise.

The unobservable latent variable, y_{2i}, which determines whether or not a worker is selected from the queue, is hypothesized to be positively related to the skill variables, such as education and work experience, included in x_{2i}. Variables other than skill measures are also included in x_{2i}, however, to capture differences in the supply of union jobs (across geographic regions) and in the preferences of union employers with regard to certain individual characteristics (such as race).

It is interesting to note that the model as it is developed here has a rather paradoxical implication. The queue will be composed of predominantly lower skilled workers while employers will be attempting to hire workers with just the opposite characteristics. This set of competing goals gives rise to bargaining strategies that could determine the union wage structure relative to the nonunion structure. Although we have not modeled the determination of the ultimate union sector wage, the use of constant coefficient models of that structure for cross-sectional data analysis is consistent with a bargaining determination of that structure.

The Determination of Union Status: An Econometric Framework

The probability that an individual will be observed in a union job is simply the probability that the worker is both in and selected from the queue. More formally,[17]

$$(16)\; Pr[U_i = 1] = Pr[IQ_i = 1 \ \& \ CFQ_i = 1]$$
$$= Pr[IQ_i = 1]Pr[CFQ_i = 1|IQ_i = 1]$$

Substitution from Equations 13 and 15 into Equation 16 yields

$$(17)\quad Pr[U_i=1] = Pr[y_{1i} > 0]Pr[y_{2i} > 0]$$
$$= Pr[\,\varepsilon_{1i} > - x_{1i}\beta_1\,]Pr[\,\varepsilon_{2i} > - x_{2i}\beta_2\,]$$

It is assumed that ε_{1i} and ε_{2i} have independent standard normal distributions, and this implies what can be called a partially observable bivariate probit model.[18]

Since the observable event is the union status of an individual and the queuing process is not directly observable, it is clear that if x_1 and x_2 contain the same set of variables, there will be no way to distinguish β_1 from β_2. In this case a union status likelihood function based on Equation 17 will have two distinct global optima. The second will be symmetric to the first with the maximum-likelihood estimates of β_1 and β_2 interchanged. While these two optima are well defined and have equal likelihood values, there is no way to discern which of the estimated parameter vectors corresponds to β_1 and which corresponds to β_2.

Poirier has argued that as long as there is at least one variable that is contained in one of the variable vectors (either x_1 or x_2) but not in the other, both parameter vectors are identified.[19] While the parameter vectors

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[17]Note that

$$Pr[U_i=0] = 1 - Pr[U_i=1]$$
$$= \left\{1 - Pr[IQ_i=1]\right\} + Pr[IQ_i=1]$$
$$\left\{1 - Pr[CFQ_i=1|IQ_i=1]\right\}$$

In other words, the probability that a worker is not working on a union job is the sum of the probability that the worker did not want a union job and the probability that the worker did want a union job but was not selected from the queue.

[18]The specification implicit in Equation 17 of ε_{1i} and ε_{2i} being independently distributed falls naturally out of the interpretation of the CFQ process as being conditional on being in the queue. This is explicit in Equation 16. In particular, this model does *not* require the joint IQ, CFQ process to be independent.

[19]The variables common to x_1 and x_2 include a constant, education, labor market experience and its square, and race and region dichotomous variables.

for the *IQ* and *CFQ* equations contain largely the same variables, there are two variables that affect the *IQ* decision but not the *CFQ* decision. Thus, by Poirier's criterion the model is identified.

The variables that are included in x_1 but not in x_2 are measures of seniority. Seniority is an important determinant of both the wage and nonwage benefits a worker receives from the job. How a worker's seniority affects the decision regarding union employment depends on whether any accrued seniority is related to a union job or a nonunion job. A worker employed on a union job has what is called union seniority (*USEN*). The union worker would have to give up this seniority to take a nonunion job; thus *USEN* will have a positive effect on the probability that the worker desires a union job. In contrast, a worker on a nonunion job has what is called nonunion seniority (*NUSEN*). The nonunion worker would have to give up this seniority to take a union job; thus *NUSEN* will have a negative effect on the probability that the worker desires a union job.[20] Clearly, being selected from the queue entails changing jobs; seniority should have no effect on the probability of being chosen from the queue.

Implicit in this discussion is the fact that the analysis is developed conditional on the worker's union status in the last period. Simply put, the hypothesis formulated in the last paragraph states that workers who held a union job last year are more likely to desire a union job this year, and workers who held a nonunion job last year are less likely to desire a union job this year. In addition, the magnitude of the effect of last year's union status is directly related to the seniority the worker has.[21]

The fact that the analysis is conditional on last period's job status has an important implication, which provides a more fundamental kind of identification than is provided by the fact that seniority does not affect $Pr[CFQ_i=1|IQ_i=1]$. Note that a worker who is observed to hold the same union job both last year and this year did not have to be selected from the queue this year. The worker held job rights to that union job, and desiring the union job ($IQ_i=1$) is a sufficient condition for finding him or her in the union job. Alternatively, if the worker voluntarily leaves a union job to take a nonunion job, it can be inferred that the individual did not want a union job, without reference to selection from the queue.

A worker who has job rights to a union job does not have to join the queue.[22] In the context of the model, $Pr[CFQ_i=1 \mid IQ_i=1 \ \& \ JR_i=1]=1$ where JR_i is a dichotomous variable that equals one if individual i has job rights to a union job and zero otherwise. Thus, from Equations 13, 16, and 17

(18) $Pr[U_i=1|JR_i=1]$

$= Pr[IQ_i=1]Pr[CFQ_i=1|IQ_i=1 \ \& \ JR_i=1]$

$= Pr[IQ_i=1]$

$= Pr[\varepsilon_{1i} > - \ x_{1i}\beta_1].$

Thus, for individuals with job rights the union-status decision is modeled as a simple univariate probit under the assumption

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The data are discussed in the next section. (For a discussion of technical details, see Dale J. Poirier, "Partial Observability in Bivariate Probit Models," *Journal of Econometrics*, Vol. 12, No. 2 (February 1980), pp. 209 – 18.)

[20]A referee correctly noted that most workers would be required to give up their union seniority if they took a job covered by a different union. We have deliberately chosen not to model multiple union queues since this appears to be econometrically untractable and is not in the spirit of the other unionization models we are extending (Lee, "Unionism and Wage Rates," and Schmidt and Strauss, "The Effect of Unions on Earnings and Earnings on Unions," for example). Nevertheless, the union-specific job-tenure variable certainly captures the essential notion that workers covered by a collective bargaining agreement negotiate a compensation package tied to seniority in ways that are not transportable to the nonunion sector. Nonunion workers may also receive compensation packages tied to seniority in ways that are not transportable to other employers (see Edward P. Lazear, "Why is There Mandatory Retirement?" *Journal of Political Economy*, Vol. 87, No. 6 (December 1979), pp. 1261 – 84). Enforceability of collective bargaining agreements allows substantial differences in the methods used by union and nonunion employers in tying compen sation to seniority.

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[21]A worker fired from or voluntarily leaving the job he or she held last year does not have any effective seniority. Hence, both *USEN* and *NUSEN* would be zero for such an individual.

[22]A worker who is fired from the union job held last year also does not have any job rights and must join the queue like any other worker without job rights.

that ε_{1i} is distributed as a standard normal. This provides sufficient information to identify the queuing process. Some care is required to interpret the resulting probabilities correctly, however.

As in Equations 15 – 17 above, the result in Equation 18 describes the conditional probability of holding a union job, given queue and job-rights information. Of necessity, union status, queue status, and job-rights status must possess a joint probability distribution in which they are structurally correlated. To construct the joint probabilities of union and job-rights status we require the marginal distribution of current job-rights status. This requires the marginal distribution of union status in the previous year, along with job separation information. Although our model allows calculation of this quantity given separation assumptions, we do not require it for estimation. The interested reader may calculate all the relevant probabilities by taking a new entrant into the labor force ($JR_i = 0$) and using the laws of probability to move from year to year.

Equation 18 thus uses an observable characteristic (JR_i) to distinguish two different functional forms for the structural probability of current union status, conditional on previous employment history. This is the sense in which job-rights information is the key to identification of the IQ and CFQ structural equations. Equation 17 still represents the probability of observing a worker in a union job for workers who have no job rights to a union job ($JR_i=0$).[23]

The likelihood function for union status, conditional on previous union status and job rights, can be derived in a straightforward fashion from Equation 17 for those workers without job rights, from Equation 18 for those workers with job rights, and from the assumption that ε_{1i} and ε_{2i} have

standard normal distributions.[24] We now turn to the description of the data used to estimate the parameters of this likelihood function (β_1 and β_2).

The Data and Empirical Analysis

The sample consists of 1341 males who, according to the 1976 probability sample of the University of Michigan's *Panel Study of Income Dynamics*, (a) were employed, but not self-employed, in 1976, (b) were not employed in the construction industry,[25] and (c) were head of the same household in 1975 and 1976.[26] An individual is assumed to be a union member if his primary job is covered by a collective bargaining agreement.[27] Under this definition, 439 individuals, representing 32.7 percent of the 1341 in the sample, were union members.

The definitions, means, and standard deviations of the variables used in the analysis are shown in Table 1. The in-queue (IQ) vector (x_1) includes a constant; dichotomous variables for education, race, and region; potential labor market experience and its square; union seniority and its square; and nonunion seniority and its square. The chosen-from-queue (CFQ) vector (x_2) includes all the same variables except the four seniority measures. Over 38 percent of the sample reported having exactly twelve years of education. Thus, education is entered in interval form, with the base group having exactly twelve years of education. Overall, the base group with reference to the dichotomous variables is

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[23]More formally,

$Pr[U_i=1|JR_i=0] = Pr[IQ_i=1]Pr[CFQ_i=1|IQ_i=1 \& JR_i=0]$

where $Pr[CFQ_i=1|IQ_i=1 \& JR_i=0] = Pr[\varepsilon_{2i} > -x_{2i}\beta_2]$. It should also be noted that workers who quit one union job to take another do not have job rights to their new union job. Thus, they need to be selected from the queue as if they did not previously hold a union job.

[24]The assumption that the variances of ε_1 and ε_2 equal one is an arbitrary normalization common to probit models necessary to fix the scale of β_1 and β_2.

[25]The mechanism by which workers are assigned to union jobs in the construction industry is substantially different from the mechanism modeled above.

[26]This criterion is used to ensure that the information on job and union status in 1975 relates to the same person as the analogous information for 1976. It is necessary to consider this problem because the *Panel Study of Income Dynamics* follows households and not individuals. Survey Research Center, *A Panel Study of Income Dynamics: Procedures and Tape Codes, 1977 Interviewing Year* (Ann Arbor: Institute for Social Research, University of Michigan, 1978).

[27]Primary job refers to the main job of the individual as opposed to second or part-time jobs.

Table 1. Means, Standard Deviations, and Definitions of Variables for 1976.

Variable	Definition	Mean	Standard Deviation
Union	Union job = 1	.3274	—
Education	Years of education	12.72	2.872
Education < 12	Education < 12 = 1	.2028	—
Education > 12	Education > 12 = 1	.4161	—
Experience *(EXP)*	*EXP*=Age – education – 6	19.18	12.99
Experience Squared	*EXP* squared	536.5	614.9
Union Seniority *(USEN)*	Years of union seniority	11.11[a]	9.312[a]
Union Seniority Squared	—	210.1[a]	271.8[a]
Nonunion Seniority *(NUSEN)*	Years of nonunion seniority	8.967[b]	8.441[b]
Nonunion Seniority Squared	—	151.7[b]	254.8[b]
Nonwhite	Nonwhite = 1	.1044	—
South	South = 1	.3020	—
Job Rights	Job rights = 1[c]	.2901	—

N = 1341

[a] Taken over the subsample of 408 workers who had positive *USEN*. These 408 workers were not fired or laid off from the union job they held in 1975.

[b] Taken over the subsample of 826 workers who had positive *NUSEN*. These 826 workers were not fired or laid off from the nonunion jobs they held in 1975.

[c] See text for definition of job rights.

white males from the non-South with twelve years of education.

The union job-rights variable (JR_i) discussed in the previous section is defined to be one if the individual held a union job in 1975, was not fired from that job, and did not quit that job to take another union job. Twenty-nine percent of the sample (389 workers) held job rights by this definition. Workers with job rights had a much higher probability of holding a union job in 1976 than those without job rights (.936 vs. .0788). In light of this fact, it is important to keep in mind that the analysis is conditional on job rights and hence on 1975 union status.

We first estimate the constrained version of the model where there is no queue. This is the special case of the queuing model where $Pr[CFQ_j=1|IQ_j=1] = 1$ for all individuals. More formally, assume that $Pr[U_i=1]$ is given by Equation 12 instead of Equations

17 and 18.[28] This implies a univariate probit likelihood function for union status. The results of both models are presented in Table 2. The first column of Table 2 contains the maximum-likelihood estimates of the parameters of the no-queue model. The maximum log-likelihood value is – 331.9 (column labeled "No-queue model") which compares to a log-likelihood value of – 847.9 (not shown in the table) for a constrained version of the model in which all the elements of β_1 except the constant are

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[28]This is a constrained version of the model in which all the parameters except the constant of the β_2 vector in Equation 15 are set equal to zero and the constant is set equal to some arbitrary, large, positive number. This yields an arbitrarily close approximation to

$$Pr[CFQ_j=1|IQ=1] = Pr[\varepsilon_{2i} > -x_{2i}\beta_2] = 1$$

for all i.

Table 2. Maximum Likelihood Estimates of the Parameters
of the Union Status Functions.[a]
(The numbers in parentheses are asymptotic standard errors)

Coefficient of	No-Queue Model	Queue Model	
	$Pr[U=1]$ (Equation 12)	$Pr[IQ=1]$ (Equation 13)	$Pr[CFQ=1\|IQ=1]$ (Equation 15)
Constant	− .2349 (.1724)	1.2810 (.3168)	− 1.1270 (.3279)
Education < 12	.2530 (.1596)	.6454 (.2740)	− .2478 (.3101)
Education > 12	− .4735 (.1374)	− .2697 (.2273)	− .5269 (.2533)
Experience	− .0224 (.0199)	− .0867 (.0344)	.1005 (.0410)
Experience Squared	$.3089 \times 10^{-3}$ $(.3748 \times 10^{-3})$	$.1457 \times 10^{-2}$ $(.6465 \times 10^{-3})$	$- .2099 \times 10^{-2}$ $(.8489 \times 10^{-3})$
Union Seniority	.4030 (.0233)	.2285 (.0357)	—
Union Seniority Squared	− .0110 $(.7458 \times 10^{-3})$	$- .6217 \times 10^{-2}$ $(.1054 \times 10^{-2})$	—
Nonunion Seniority	− .2148 (.0258)	− .2566 (.0408)	—
Nonunion Seniority Squared	$.5923 \times 10^{-2}$ $(.8316 \times 10^{-3})$	$.7294 \times 10^{-2}$ $(.1506 \times 10^{-2})$	—
Nonwhite	.3134 (.1576)	− .0613 (.2819)	.6068 (.3867)
South	− .3245 (.1316)	− .4596 (.2561)	.0526 (.2791)
Log-Likelihood	− 331.9	− 305.5	
N = 1341			

[a] The base group consists of white nonsouthern males with twelve years of education.

set equal to zero. A likelihood-ratio test of this set of constraints rejects the constrained version of the model at any reasonable level of significance. Thus, the no-queue model has a significant amount of explanatory power for union status.

The qualitative nature of the parameter estimates of the no-queue model accords well with the results of previous studies of the union status of individuals. Union status is negatively related to the education level of the individuals. Union seniority and nonunion seniority have the hypothesized effects (positive and negative, respectively) on union status, and nonwhites are more

likely than whites to hold a union job while southerners are less likely to hold a union job. It is interesting to note that neither the coefficient of experience nor the coefficient of its square is significantly different from zero. In addition, a Wald-test of the hypothesis that these two coefficients are zero fails to reject the hypothesis at conventional levels of significance.[29] This suggests that labor market experience is not a significant deter-

[29] The test statistic was 2.215, which is less than the critical value of a X^2 distribution with two degrees of freedom at the .25 level of significance of 2.77.

minant of union status in spite of the evidence cited above that the wage advantage of unionization is inversely related to labor market experience.

The last two columns of Table 2 contain the maximum-likelihood estimates of the parameters of the queuing model defined in Equations 17 and 18. The maximum log-likelihood value is − 305.5, which compares to a log-likelihood value of − 331.9 in the no-queue version of the model. The likelihood ratio test statistic of 52.8 (not shown in the table) is sufficient to reject the no-queue model at reasonable levels of significance.[30]

The results are generally consistent with the hypotheses derived previously. For the in-queue relationship, individuals with fewer than twelve years of education are significantly more likely, at the one percent level of significance, to desire a union job than workers with exactly twelve years' education. Individuals with more than twelve years of education are less likely to desire a union job than individuals with exactly twelve years' education, although this estimated effect has a relatively large standard error.

The estimates of the queue model imply that additional labor-force experience significantly reduces the probability of desiring a union job. Holding union and nonunion seniority constant at zero, the marginal impact of an additional year of labor-force experience on the probability of being in the union queue is negative over the relevant working life of the individual. The marginal effect of experience is dampened as experience increases. On the other hand, there is no statistically significant effect of experience on union status in the no-queue model. In addition, the probability of being in the queue is more sensitive to changes in experience, absolutely, than is the probability of being in a union job in the no-queue model.[31]

Union and nonunion seniority have the hypothesized effects on the probability of desiring a union job. Union seniority is positively related to $Pr[IQ=1]$, and nonunion seniority is negatively related to $Pr[IQ=1]$.

A result that is contrary to our expectation is that nonwhites are not more likely to desire a union job. The lower white-black wage differential generally found in the union sector suggested that blacks would be more likely than whites to desire a union job.[32] In addition, Farber and Saks have found that nonwhites were more likely to vote for union representation in National Labor Relations Board supervised representation elections, even after controlling for the wage effect.[33] The nonwhite coefficient in our queue equation is essentially zero. This is consistent with nonwage benefits of unions accruing primarily to whites. The effect would cancel the wage benefit effect modeled above.

A final result is that southern workers are less likely to desire a union job than non-southerners. It is possible to reject the hypothesis that the coefficient on south equals zero against the alternative that the coefficient is less than zero at the five percent level of significance.

Now consider the results of the chosen-from-queue equation (CFQ) reported in the last column of Table 2. Given the definition of the job-rights variable, the CFQ function may just be measuring the effect of the previous year's union status on the current year's union status, and it is trivial to say simply that nonunion workers last year are less likely to be union workers this year. Any explanatory power must come from the variables included in the CFQ function. As a first test, the hypothesis that all the coefficients in the CFQ function except the constant are zero can be rejected at the .05 level of significance using the usual likelihood-ratio test.[34] This suggests that the charac-

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[30]The no-queue model embodies seven constraints on the queuing model. These are described in footnote 28. The critical value of a X^2 distribution with seven degrees of freedom at the .005 level of significance is 20.3.

[31]These claims are supported statistically in the appendix, which is available from the authors.

[32]See Ashenfelter, "Racial Discrimination and Trade Unionism."

[33]Henry S. Farber and Daniel H. Saks, "Why Workers Want Unions: The Role of Relative Wages and Job Characteristics," *Journal of Political Economy*, Vol. 88, No. 2 (April 1980), pp. 349− 69.

[34]The log-likelihood of the constrained model is

teristics included in the *CFQ* function do indeed have a significant effect on the selection probability.

The hypothesized direct relationship between $Pr[CFQ=1|IQ=1]$ and education is not supported by the data. In fact, workers with more than twelve years of education are significantly less likely to be chosen from the queue than those workers with exactly twelve years of education. The estimated $Pr[CFQ=1|IQ=1]$, based on the coefficients in Table 2, for white non-southern workers with no experience or seniority (new entrants) and twelve years of education is .13, while otherwise identical workers with more than twelve years of education have an estimated probability of being chosen from the queue of .049.

Two potential explanations can be advanced for this result. First, union employers may perceive some negative characteristic in more highly educated workers that offsets their higher skill level. These workers may be too independent and ambitious to fit well in a structured union work environment where advancement is likely to be governed strictly by seniority. A second explanation is that the supply of union jobs suitable for highly educated workers may be relatively small. Thus, the education variables in the *CFQ* function may be picking up a supply effect.[35]

Since the supply of union jobs is not as

likely to be differentiated by appropriate experience levels as by appropriate education levels, the experience levels ought to reflect more closely the employer decision calculus described previously. Indeed, this is the case. The effect of experience on the probability of being chosen from the queue is significantly positive. For levels from zero to over twenty years of experience, the marginal effect of experience, although declining, is positive. The effect of additional experience is essentially zero for the most experienced workers. This is not an anomoly since the theoretical model does not predict that the derivative of the positive effect of experience should be of any sign.[36]

In order to summarize the effects of experience on union status, Table 3 contains the marginal effects of experience on $Pr[IQ=1]$, $Pr[CFQ=1|IQ=1]$ and $Pr[U=1]$. It is clear that in both the queue and no-queue models the effect of experience on $Pr[U=1]$ is relatively small. However, it is also clear that this masks the real relationships between experience and the probabilities of being in the queue and of being chosen from the queue. Columns 1 and 2 of Table 3 show that the relatively small overall effect of experience on union status is the result of a negative effect on the probability of being in the queue combined with a larger (in absolute value) positive effect on the probability of being chosen from the queue.

Nonwhites have a higher probability of being chosen from the queue. However, this coefficient is significantly different from zero only at the ten percent level of significance. This result suggests that union employers have a preference for nonwhite workers that is not explicitly captured elsewhere.

The queuing model has some interesting implications for the allocation of an individual's time to the union and nonunion sectors over the life cycle. New entrants to the labor force with a high school education have a probability of desiring a union job of .90.[37] However, these workers have a prob-

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− 312.7, which gives a likelihood-ratio statistic of 14.4 when compared to the log-likelihood of − 305.5 in the unconstrained model. The critical values of a X^2 distribution with 6 degrees of freedom at the .05 level is 12.6.

[35]This highlights a weakness of the assumption implicit in the formulation of the model that there is only one queue. A more realistic model would allow for the possibility that there are likely to be different queues for workers of different basic skill levels. For example, there may be distinct queues for white-collar and blue-collar workers or for workers with high versus low levels of schooling. In this case we might expect qualitatively different results for different stratifications of the sample. In fact, when our model is estimated using just the blue-collar workers, the results are qualitatively very similar to the results discussed in the text. Results for high school graduates and lower schooling levels are qualitatively similar but poorly determined. The blue-collar results are contained in a separate appendix available from the authors on request.

[36]These claims are supported statistically in the appendix available from the authors.

[37]These and all the following numbers used in this discussion are derived from Table 2 for white non-southern workers with twelve years of education.

Table 3. The Marginal Effect of Changes in Experience on
$Pr[IQ=1]$, $Pr[CFQ=1|IQ=1]$, and $Pr[U=1]$
in the Queue and No-Queue Models.[a]

Experience	Queue			No-Queue
	$\dfrac{\partial Pr[IQ=1]}{\partial EXP}$ [b]	$\dfrac{\partial Pr[CFQ=1|IQ=1]}{\partial EXP}$ [c]	$\dfrac{\partial Pr[U=1]}{\partial EXP}$ [d]	$\dfrac{\partial Pr[U=1]}{\partial EXP}$ [b]
0	− .01523	.02125	.01608	− .00871
5	− .01947	.02522	.01562	− .00729
10	− .01964	.02209	.00846	− .00592
20	− .01125	.00659	− .00570	− .00344

[a] Computed for white nonsouthern workers with twelve years of education and no seniority.

[b] $\dfrac{\partial Pr[IQ=1]}{\partial EXP} = (\beta_{1E} + 2\beta_{1ES}(EXP))f(x_1\,\beta_1)$ where β_{1E} and β_{1ES}
are the coefficients of experience and experience squared in the in-queue equation and $f(\,\bullet\,)$ is the standard normal density.

[c] $\dfrac{\partial Pr[CFQ=1|IQ=1]}{\partial EXP} = (\beta_{2E} + 2\beta_{2ES}(EXP))f(x_2\,\beta_2)$ where β_{2E} and β_{2ES}
are the coefficients of experience and experience squared in the chosen-from-queue equation and $f(\,\bullet\,)$ is the standard normal density.

[d] $\dfrac{\partial Pr[U=1]}{\partial EXP} = \dfrac{\partial Pr[IQ=1]}{\partial EXP}\,Pr[CFQ=1|IQ=1] + \dfrac{\partial Pr[CFQ=1|IQ=1]}{\partial EXP}\,Pr[IQ=1]$
by a simple application of the chain rule for differentiation.

ability of being selected from the queue of only .13. Thus, they are unlikely to get a union job $(Pr[U=1] = .12)$. If the worker does not get a union job for five years and holds one nonunion job for that time, the probability of desiring a union job falls to .41. After ten years on the nonunion job, the worker's probability of desiring a union job falls to .10, and after twenty years it falls to .02.

Table 4 contains $Pr[IQ=1]$ for various values of experience and nonunion seniority. Consistent with the above discussion, the numbers in this table suggest that $Pr[IQ=1]$ is relatively high for workers with no nonunion seniority, but it falls rapidly with the accrual of nonunion seniority. Given that the probability of being selected from the queue is relatively small at most experience levels, only those workers who are "lucky" enough to be selected from the queue while they are still in it (at low *NUSEN*) will be found in union jobs.[38]

Once a worker is selected from the queue, the accrual of union seniority makes it progressively less likely that he or she will voluntarily give up a union job for a nonunion job. Since these workers have job rights, they do not have to be selected from the queue again; so they will remain in their union jobs.

Intuitively, individuals will become long-run union workers if they are successful at being selected from the queue relatively early in their working life, before they build up too much nonunion seniority. As they build up nonunion seniority, they become less willing to sacrifice the benefits associated with this seniority in order to take a union job.[39] Since most high school gradu-

[38] The $Pr[CFQ=1|IQ=1]$ is .1299 for workers with $EXP=0$, .2494 with $EXP=5$, .3670 with $EXP=10$, and .5173 with $EXP=20$.

[39] Of course, workers who are fired or permanently laid off from their nonunion jobs lose nonunion seniority and may desire union jobs. However, the probability of losing one's job falls with seniority even in the nonunion sector. See Richard B. Freeman, "The Exit-Voice Tradeoff in the Labor Market: Unionism, Job Tenure, Quits and Separations," *Quarterly Journal of Economics*, Vol. 94, No. 4 (June 1980), pp. 643–73.

Table 4. Probabilities of Being in the Queue for Various Values
of Experience and Nonunion Seniority.[a]

Nonunion Seniority (in years)	Experience (in years)				
	0	*5*	*10*	*15*	*20*
0	.8999	.8116	.7121	.6210	.5516
5	—	.4142	.2942	.2140	.1658
10	—	—	.1008	.0632	.0439
15	—	—	—	.0287	.0188
20	—	—	—	—	.0185

[a] Computed for nonsouthern whites with a high school education from the estimates contained in Table 2.

ates would prefer a union job when they enter the labor force, it becomes the employer's decision as to who will make up the union labor force in the long run.

Summary and Conclusions

A model of the determination of the union status of individual workers that allowed for the possibility of queuing for union jobs was developed. It was hypothesized that more skilled workers would be less likely to desire a union job while union employers would be more likely to want to hire more skilled workers.

Overall, the empirical results provide evidence that is generally supportive of the queuing hypothesis. The no-queue model can be rejected, using a likelihood-ratio test, and the primary skill measure (experi- ence) is negatively related to the probability of desiring a union job and positively related to the probability of being chosen from the queue, as was hypothesized. These results suggest that a simple probit model for union status may be misspecified because it is not based on any consistent behavioral theory. Such a probit was shown to hide a number of interesting relationships and, to the extent it does so, it is misleading.

Finally, it was shown that the model has implications for the allocation of workers to the union and nonunion sectors. The results suggest that most new entrants prefer union jobs but cannot get them. As time goes by and workers accrue nonunion seniority, they become less likely to want union jobs. Thus, the union status of most workers is determined by their success in being selected from the queue early in their working life.