The Decline of Unionization in the United States: What can be Learned from Recent Experience?

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The Decline of Unionization in the United States: What Can Be Learned from Recent Experience?

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The dramatic decline in unionization over the last decade is investigated in the context of a supply/demand model of union status determination using data from surveys of workers conducted in 1977 and 1984 along with data from the National Labor Relations Board on representation elections. It is concluded that the decline in unionization since 1977 is accounted for largely by (1) an increase in employer resistance to unionization, probably due to increased product market competitiveness and (2) a decrease in demand for union representation by nonunion workers due to an increase in the satisfaction of nonunion workers with their jobs and a decline in nonunion workers' beliefs that unions are able to improve wages and working conditions.

I. Introduction

The last 15 years have seen a precipitous decline in the extent of unionization in the United States. Based on data from the May Current Popu-

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lation Surveys (CPS), summarized in table 1, the fraction of private nonagricultural employment made up of union members fell from 25.6% in 1973 to 14.1% in 1985. The reasons for this decline are not clear. Farber (1985) and Dickens and Leonard (1985) present evidence that shifts in the demographic, industrial, and occupational composition of the labor force away from traditionally heavily unionized types of workers and sectors accounted for a substantial fraction of the decline in unionization prior to the mid-1970s. However, evidence based on data from the Current Population Survey on the union status of workers and presented in the first part of this study suggests that only a small part of the decline since that period can be accounted for by these shifts.

In the remainder of this study, a queuing model of the determination of the union status of workers (Farber 1983) is used to decompose the decline in unionization into two components: (1) a decline in the demand by workers for union representation and (2) a decline in the supply of union jobs relative to this demand. While the usual data on the union status of workers are not sufficient to identify shifts in demand and supply separately, data from the 1977 Quality of Employment Survey (QES) (Quinn and Staines 1979) and a survey done by Louis Harris and Associates for the AFL-CIO in 1984 (AFL) contain information that are adequate for this decomposition. Data from the National Labor Relations Board (NLRB) on union organizing activity are also analyzed.

It is concluded that the decline in unionization since 1977 is accounted for largely by two factors. First, there has been an increase in employer resistance to unionization, probably due to increased product market competitiveness. Second, there has been a decrease in demand for union representation by nonunion workers due to an increase in the satisfaction of nonunion workers with their jobs and a decline in nonunion workers' beliefs that unions are able to improve wages and working conditions.

II. The Decline in the Fraction Unionized

The best available data for measuring the decline in the fraction of the work force that is unionized is the May Current Population Survey

Table 1
Union Membership as Fraction of Private Nonagricultural Employment, 1973–84

Year	Fraction	Year	Fraction
1973	.256	1980	.208
1974	.249	1981	.197
1975	.230	1982	. . .
1976	.226	1983	.159
1977	.218	1984	.150
1978	.209	1985	.141
1979	.220		

NOTE.—Based on tabulation of May CPSs, 1973–84.

(CPS).This survey is collected in a consistent fashion from year to year, and information on union membership is available for every year from 1973 to 1985 with the exception of 1982. The analysis focuses on 1977 and 1984 because these are the years for which the supplemental information in the QES and AFL survey, used later in this study, are available.

Samples of workers were derived from the two CPSs in a similar fashion. The May 1984 CPS has data on union status for only 25% of the overall sample while the May 1977 CPS has data on union status for the entire sample. A 25% random subsample of the May 1977 CPS was used along with all of the May 1984 CPS with data on union status. The final samples (9,912 workers in 1977 and 10,676 workers in 1984) consist of all non-managerial workers who were not self-employed and for whom complete information was available on the workers' demographic characteristics, industry, occupation, and union status.

Simple tabulation of the data confirm the dramatic decline in unionization between 1977 and 1984. Fully 26.8% of the workers in the May 1977 CPS sample, compared with only 21.4% of the May 1984 CPS sample, reported that they were union members.[1] In order to investigate the extent to which this decline can be accounted for by the standard explanations of shifts in the demographic, occupational, and industrial composition of the labor force, table 2 contains mean sample values for each year for a set of variables representing various dimensions of labor-force structure along with the fraction of workers in each group who report themselves to be union members.

Three clear patterns emerge from table 2. First, the results confirm the conventional wisdom regarding which types of workers and jobs are relatively heavily unionized: (1) males, nonwhites, and workers living outside the South; (2) jobs in manufacturing, construction, and the transportation, communication, and public utility industries; and (3) workers in blue-collar jobs. The second pattern is that there have indeed been shifts in employment (1) away from relatively heavily unionized jobs in manufacturing industries and (2) away from relatively heavily unionized blue-collar jobs. Finally, the fraction unionized fell between 1977 and 1984 in virtually all categories *and the decline was generally greatest among the most heavily unionized workers.* The conclusion is that shifts in labor-force structure cannot fully account for the decline in unionization.

In order to determine precisely how much of the decline in unionization can be accounted for by shifts in labor-force structure, a more formal model is needed. The simplest empirical model of the union status of workers is a univariate discrete-choice model. In this model a worker is unionized ($U_i = 1$) only if some latent variable, Y_{1i}, is positive. The interpretation generally given to this latent variable (Lee 1978) is that it rep-

[1] These numbers are higher than the tabulations presented in table 1 because the latter exclude workers in the relatively highly unionized public sector.

Table 2
Sample Proportions and Fraction Unionized, Broken Down by Labor-Force Structure, Using May CPS Data

	1977		1984	
	Sample Fraction	Fraction Unionized	Sample Fraction	Fraction Unionized
Total	1.0	.268	1.0	.214
Sex:				
Female	.432	.182	.474	.202
Male	.568	.334	.526	.262
Race:				
Nonwhite	.106	.337	.126	.259
White	.894	.260	.874	.208
Region:				
South	.287	.165	.309	.139
Nonsouth	.713	.309	.691	.245
Industry:				
Manufacturing	.300	.382	.259	.291
Construction	.068	.384	.066	.290
Transportation, communications, public utilities	.076	.497	.078	.426
Trade	.202	.129	.199	.081
Finance, insurance, real estate	.058	.045	.067	.021
Services	.296	.205	.331	.209
Occupation:				
Blue collar	.423	.420	.368	.331
Clerical	.202	.131	.213	.113
Service	.131	.181	.142	.155
Professional	.173	.217	.202	.220
Sales	.071	.040	.075	.026
Sample size	9,912		10,676	

resents the difference between worker i's utility on a union job and his or her utility on a nonunion job. Given that

$$Y_{1i} = X_i\beta_1 + \varepsilon_{1i}, \tag{1}$$

where X_i is a vector of worker characteristics, β_1 is a vector of parameters, and ε_{1i} is a random component, the probability of a worker being unionized is

$$\mathrm{pr}(U_i = 1) = \mathrm{pr}(Y_{1i} > 0) = \mathrm{pr}(\varepsilon_{1i} > -X_i\beta_1). \tag{2}$$

If ε_{1i} has a standard normal distribution, then this is the usual probit model so that $\mathrm{pr}(U_i = 1) = \Phi(X\beta_1)$ where $\Phi(\cdot)$ is the standard normal cumulative-distribution function.

The CPS data described above were used to estimate separate probit models of individual union membership for 1977 and 1984. The vector X_i includes a constant plus 19 dichotomous variables representing main effects for four educational categories, five age categories, and the characteristics in table 2.[2] While the parameter estimates are not presented here, their character is consistent with the differences presented in table 2. These estimates were used to decompose the change in the average probability of unionization into two components: (1) a piece due to changes in the structure of the labor force (ΔX) and (2) a piece due to changes in the within-sector union membership fractions ($\Delta\beta$).

The quantity $P(X_j, \hat{\beta}_k)$ is the average predicted probability of union membership using the characteristics of the year j sample with the parameter estimates from the year k sample. This is

$$P(X_j, \hat{\beta}_k) = \frac{1}{n_j} \cdot \sum_{i=1}^{n_j} \Phi(X_{ji}\hat{\beta}_k), \tag{3}$$

where X_{ji} is a vector of the characteristics of the ith individual in year j, $\hat{\beta}_k$ is the vector of probit parameters estimated from the year k data, and n_j is the number of individuals in the year j sample. Using this notation, the change in the estimated probability of union membership between 1977 and 1984 is $P(X_{84}, \hat{\beta}_{84}) - P(X_{77}, \hat{\beta}_{77})$, which can be decomposed as

$$\begin{aligned} P(X_{84}, \hat{\beta}_{84}) - P(X_{77}, \hat{\beta}_{77}) = {} & [P(X_{84}, \hat{\beta}_{77}) - P(X_{77}, \hat{\beta}_{77})] \\ & + [P(X_{84}, \hat{\beta}_{84}) - P(X_{84}, \hat{\beta}_{77})]. \end{aligned} \tag{4}$$

The first bracketed term is the change in the average probability of unionization that is due to the change in labor-force structure between 1977 and 1984, holding the within-sector probability of union membership (β) fixed at the estimated 1977 levels. The second bracketed term is the change in the average probability of unionization due to the change in the within-sector probability of union membership, holding labor-force structure (X) fixed at the 1984 sample values.

The overall estimated average change in the probability of union membership is −.0545 with an estimated asymptotic standard error of .00550.[3] The component due to the change in labor-force structure is −.01101 with an asymptotic standard error of .001079. The component due to change in the within-sector propensity for union membership is −.0429 with an asymptotic standard error of .005513. Thus, only approximately 20% of the decline in union membership (.01101/.0545 = .202) can be accounted

[2] The four educational categories are $ED < 12$, $ED = 12$, $12 < ED < 16$, and $ED \geq 16$. The five age categories are $AGE \leq 24$, $25 \leq AGE \leq 34$, $35 \leq AGE \leq 44$, $45 \leq AGE \leq 54$, and $AGE \geq 55$.

[3] The asymptotic standard errors are computed from first-order approximations to the asymptotic variances of the appropriate nonlinear function of the estimated parameters.

for by changes in labor-force structure. The rest is accounted for by declines in the within-sector probabilities of union membership.[4]

III. A Supply-Demand Model of the Extent of Unionization

A useful tool for the analysis of the decline of unionization is an economic model based on movements in the supply of and demand for union representation (Abowd and Farber 1982; Farber 1983). The supply of union jobs in a simple version of this model is the result of decisions by unions regarding the optimal stock of union jobs. These decisions are based on the relative costs and benefits of organization to the unions and their members. Where jobs are newly organized, the current workers may bear the full cost of organization. However, once a job is unionized, it remains unionized regardless of who holds that job. Thus, the long-run benefits of investment in organization accrue not only to those who organize a job but also to future holders of the job. For example, the investment in organizing the automobile industry in the 1930s is still paying returns to current workers in the industry.

A price mechanism through which a union can charge members (demanders of union services) a market price generally does not exist. Unions charge dues and entry fees, but these are not sufficient to offset the benefits of union membership (Raisian 1983). Neither can workers who organize jobs sell the rights to vacancies in these jobs. In most industries, despite requirements regarding eventual union membership, employers can hire whomever they desire. Thus, unions generally cannot ration membership in other ways.[5]

This market failure has two important implications. First, unions and workers will not invest enough in union organizing activity because they will not be able to recoup their investment fully by charging those who will benefit in the future. Second, there will be workers who would not want to undertake organization themselves but who would prefer either a vacancy in a union job or that their job be organized by others. Thus, there is excess demand for vacancies in existing union jobs.

With regard to the simple model of union status determination outlined in the previous section, the implication is that not all nonunion workers who would prefer their job to be unionized by the criterion outlined above

[4] Alternatively, this decomposition can be carried out using (1) the 1984 within-sector propensities to unionize (β_{84}) to compute share accounted for by shifts in labor-force structure, and (2) the 1977 characteristics (X_{77}) to compute the share accounted for by shifts in the within-sector propensity for union membership. This decomposition yields the even more extreme result that only approximately 10% of the decline in union membership can be accounted for by changes in labor-force structure.

[5] Industries that are characterized by hiring halls where unions effectively make decisions regarding who gets hired are exceptions.

($Y_{1i} > 0$) are willing to organize their current job. To the extent that this is true, $\text{pr}(U_i = 1) \neq \text{pr}(Y_{1i} > 0)$. There will be some nonunion workers for whom $Y_{1i} > 0$, and the univariate probit model of union status summarized in equations 1 and 2 loses its structural interpretation.

The set of workers who demand union representation is composed of two distinct groups. The first group consists of those workers who both demand union representation and were hired by a union employer. This group is easy to identify. They are simply the workers who report themselves to be union members. The second group consists of workers who demand union representation but are unable or unwilling either to find a union job or to organize their current job. This group is more difficult to identify since they are a subset of the workers who report themselves as not union members. The key to the analysis in this study is that data are available in 1977 and 1984 that can be used to split the group of nonunion workers into a group that prefers union representation and a group that does not. These data are described in Sections IV and V.

In this model the quantity of unionization is determined by the stock of union jobs. This is the result of organization undertaken by unions and the success of these efforts. Clearly unions' costs of organization and likelihood of success in organizing are affected by (1) worker demand for union representation and (2) employer resistance to union organizing efforts. Two measures are relevant for investigating the importance of these factors. The first is the level of worker demand for union representation measured as the fraction of the work force that demands union representation. This focuses on workers' decisions regarding the benefits of union membership relative to the (small) costs of filling an existing union vacancy. The second is the supply of union jobs relative to demand measured as the ratio of satisfied demand (unionized workers) to total demand (unionized workers plus excess demand). This focuses on the costs of union organization conditional on the level of demand.

Employer resistance to unionization can have two effects on these measures. First, employer resistance will reduce the demand for union representation among nonunion workers. This can occur as nonunion employers (*a*) pay workers higher wages and introduce union-like fringe benefits and personnel policies and (*b*) implicitly or explicitly threaten workers regarding unfavorable results of unionization. Second, employer resistance can make it more difficult to organize even workers with an underlying demand for union representation. This will reduce the optimal stock of union jobs both overall and relative to the level of demand.

The demand for union representation by an individual worker is modeled as a discrete choice problem where the worker compares his/her utility on union and nonunion jobs and selects the job with the higher utility. This utility difference is defined above (eq. [1]) as Y_{1i}. Thus, worker i will prefer union representation ($D_i = 1$) only if Y_{1i} is positive, and worker i

does not prefer union representation ($D_i = 0$) otherwise. On this basis, the probability that worker i prefers union representation is

$$\mathrm{pr}(D_i = 1) = \mathrm{pr}(Y_{1i} > 0) = \mathrm{pr}(\varepsilon_{1i} > -X_i\beta_1). \tag{5}$$

Given data on D_i for *all* workers and assuming a standard normal distribution for ε_{1i}, the parameter vector β_1 can be estimated using standard univariate probit maximum-likelihood techniques.

The relative supply of union jobs is measured by the ability of a worker who demands union representation to find a union job. This also is modeled as a discrete-choice problem. More formally, a worker who demands union representation will find a union job ($U_i = 1$) only if some index Y_{2i} is positive, and the worker will not find a union job ($U_i = 0$) otherwise. Let

$$Y_{2i} = X_i\beta_2 + \varepsilon_{2i}, \tag{6}$$

where β_2 is a vector of parameters, and ε_{2i} is a random component. Thus, the probability that worker i, who prefers union representation, is able to find a union job is

$$\begin{aligned} \mathrm{pr}(U_i = 1 \mid D_i = 1) &= \mathrm{pr}(U_i = 1 \mid D_i = 1), \\ &= \mathrm{pr}(Y_{2i} > 0) = \mathrm{pr}(\varepsilon_{2i} > -X_i\beta_2). \end{aligned} \tag{7}$$

Given data on U_i for *all* workers who demand union representation and assuming a standard normal distribution for $\varepsilon_{2i} \mid D_i = 1$, the parameter vector β_2 can be estimated using standard univariate probit maximum-likelihood techniques.[6]

Based on these relationships, a worker i will be unionized ($U_i = 1$) only if he or she both prefers a union job and is able to find one (Y_{1i} and Y_{2i} are positive). The probability of this event is

$$\begin{aligned} \mathrm{pr}(U_i = 1) &= \mathrm{pr}(D_i = 1) \cdot \mathrm{pr}(U_i = 1 \mid D_i = 1), \\ &= \mathrm{pr}(Y_{1i} > 0) \cdot \mathrm{pr}(Y_{2i} > 0). \end{aligned} \tag{8}$$

It is clear that with data only on union status it is not possible to estimate the determinants of demand (Y_{1i}) and supply (Y_{2i}) separately with any

[6] Farber (1983) develops and estimates a multivariate probit model where ε_{2i} has a standard normal distribution so that $\varepsilon_{2i} \mid D_i = 1$ has a distribution that depends on the joint distribution of ε_{1i} and ε_{2i}. Farber (1984) derives estimates of β_2 using both the univariate probit model and the multivariate probit model based on the 1977 data, and they are virtually identical. On this basis, the analysis proceeds using the more straightforward univariate probit model.

robustness. However, with additional data on demand such estimation is possible.[7]

IV. The Data and Sampling Issues in Using the QES and AFL Survey

Data from two surveys that contain information on nonunion worker demand for union representation independent of union status are used in the analysis. The first survey (QES) is the Quality of Employment Survey, which was carried out by the Survey Research Center at the University of Michigan in 1977 (Quinn and Staines 1979). The second survey (AFL) was carried out by Louis Harris and Associates, Inc., for the American Federation of Labor-Congress of Industrial Organizations in 1984 (Louis Harris and Associates 1984).

The QES was designed to yield a representative sample of American workers. However, the AFL survey is not representative of the work force in that, since its goal was to learn about the attitudes of nonunion workers toward unions in order to aid organizing efforts, union members were "quota sampled." Workers were contacted randomly by Harris, and all nonunion workers who met certain criteria (age over 18, employed, not self-employed) were administered the survey. Union members who were contacted and who met the criteria were administered the survey until a quota of 250 was reached. Twenty-eight union workers contacted after this point were counted but not administered the survey. Using this information, the probability that a randomly selected union worker was surveyed is 250/278 = 0.9. Given the central role that union membership plays as an endogenous variable in the analysis, this creates a classic choice-based sampling problem (Manski and Lerman 1977; Manski and McFadden 1981). This is accounted for in the analysis that follows in a very simple (though inefficient) way by randomly dropping 10% of the nonunion workers in the AFL sample.

Samples of workers were derived from the two surveys in an identical fashion to that used to derive the CPS samples used in Section II. These samples consist of all nonmanagerial workers who were not self-employed and for whom complete information was available on the workers' demographic characteristics, industry, occupation, union status, preference for union representation, attitudes about the general usefulness of unions, and job satisfaction. The QES sample has 961 observations while the AFL sample has 996 observations (after deleting 86 nonunion workers). The first two columns of table 3 contain sample proportions and fractions of

[7] In fact, it is possible to derive estimates of the determinants of demand and supply separately with data only on union status, but identification will depend crucially on the functional forms chosen for the distributions of ε_{1i} and ε_{2i}. See Poirier (1980).

Table 3
Sample Proportions and Fraction Unionized Broken Down by Labor-Force Structure Using QES and AFL Data

	1977 QES (Adjusted)		1984 AFL (Adjusted for Quota Sampling)		1977 QES (Adjusted for Difference with CPS)	
	Sample Fraction	Fraction Unionized	Sample Fraction	Fraction Unionized	Sample Fraction	Fraction Unionized
Total	1.0	.310	1.0	.218	1.0	.271
Sex:						
Female	.387	.239	.465	.155	.393	.208
Male	.613	.355	.535	.272	.607	.312
Race:						
Nonwhite	.118	.363	.114	.333	.117	.476
White	.882	.303	.886	.203	.883	.264
Region:						
South	.350	.200	.301	.120	.361	.180
Nonsouth	.650	.370	.699	.260	.639	.322
Industry:						
Manufacturing	.305	.437	.223	.257	.296	.387
Construction	.053	.451	.139	.167	.052	.405
Transportation, communications, public utilities	.084	.531	.093	.463	.081	.487
Trade	.158	.132	.141	.100	.162	.102
Finance, insurance, real estate	.044	.072	.056	0.0	.046	.072
Services	.356	.237	.348	.231	.363	.209
Occupation:						
Blue collar	.417	.448	.387	.312	.424	.397
Clerical	.176	.166	.215	.089	.179	.135
Service	.139	.202	.111	.271	.142	.170
Professional	.207	.252	.218	.208	.213	.232
Sales	.041	.129	.069	.044	.042	.105
Sample size	961		996		909	

union membership for the QES sample for the important-labor force structure measures that are used in the analysis. The next two columns of the table contain the same information for the AFL sample.

A potential problem with using surveys derived from different sampling frames and with different survey designs is that the resulting samples may not be comparable. To the extent that the samples are not representative in dimensions that are assumed exogenous to the analysis (the demographic and labor-force structure variables), this can be accounted for in a multivariate analysis. However, if the samples are choice based in the sense that the probability of inclusion in the samples depends on the endogenous variables (union membership, demand for union representation by nonunion workers), then there can be serious problems of statistical inference. The quota sampling of union workers in the AFL survey is a clear example of this, but both the problem and a solution are obvious because the degree of undersampling of union members is known.

The samples from the CPS used in Section II provide useful benchmarks for the fractions that are union members in the QES and AFL samples. However, there is no such benchmark available for the demand for union representation by nonunion workers, and it is assumed for the purposes of this study that there is no systematic sample selection problem in this dimension.

Since the CPSs in 1977 and 1984 are based on a consistent sampling frame and survey design and on very large samples, the analysis proceeds as if the sample union membership proportions from the CPSs are, in fact, the population proportions. The tabulations of the CPS data, contained in table 2, show that 26.8% of workers were union members in 1977 and 21.4% of workers were union members in 1984 for a drop of 5.4 percentage points. This contrasts with tabulation of the QES and AFL samples, contained in table 3, which show 31.0% union membership in 1977 and 21.8% union membership in 1984 for a drop of 9.2 percentage points. A chi-square test of the hypothesis that the fractions unionized are the same in the QES and May 1977 CPS samples can be rejected (p value = .005). The same test applied to the AFL and May 1984 CPS samples does not reject the hypothesis (p value = .799).

We are left with the problem that the QES sample overrepresents union workers. Because of this, the decline in unionization between 1977 and 1984 as measured by a comparison of the QES and AFL surveys will be an exaggeration of the true decline, and this will distort measurement of any decline in the demand for unionization and the supply of union jobs relative to demand.

It may be true that the difference in the fraction unionized between the QES and May 1977 CPS is due to differences in the distribution of measured demographic and structural characteristics. On comparing the distributions in tables 2 and 3, there are two key observations. First, the sample distri-

Table 4
Probit Model of Probability of Unionization, QES and May 1977 CPS Samples, Selected Parameters

Variable	(1)	(2)
Constant	−.4956	Not reported
	(.0423)	
CPS	−.1243	−.1212
	(.0444)	(.0473)
Labor-force structure	No	Yes
Log-likelihood (n = 10873)	−6353.0	−5353.6
Mean probability assuming all:		
QES	.310	.303
	(.0149)	(.0143)
CPS	.268	.269
	(.0045)	(.0041)
Difference	.0424	.0345
	(.0156)	(.0138)

NOTE.—Labor-force structure includes a set of 19 variables representing main effects for sex, race, age (five categories), education (four categories), industry (six categories), and occupation (five categories). The numbers in parentheses are asymptotic standard errors. The all-QES mean probability is computed using the actual values of the labor-force structure variables for the combined sample assuming $CPS_i = 0$ for all observations. The all-CPS mean probability is computed using the actual values of the labor-force structure variables for the combined sample assuming $CPS_i = 1$ for all observations.

butions for both the QES and AFL survey differ from their CPS counterparts. This is not surprising given the relatively small sizes of the QES and AFL survey. Individual cells can be especially small leading to the potential for substantial sampling variability.[8] The second observation is that the fractions unionized within categories are almost uniformly higher in the QES than in the May 1977 CPS. There is no such consistent difference between the AFL survey and the May 1984 CPS. These observations suggest that differences in sample composition cannot account for differences in the fraction unionized between the QES and the May 1977 CPS.

A multivariate probit analysis of the probability of union membership using both 1977 surveys yields the estimates in table 4. The estimates in the first column are for a specification that includes only a constant and a 0–1 dummy variable for being in the CPS survey. The estimates in the second column are for a specification that additionally includes 19 dummy variables representing main effects for the labor-force structure variables in tables 2 and 3 as well as education (4 categories) and age (5 categories). Based on the improvement in the log-likelihood value, it is clear that labor-force structure has significant explanatory power for the probability of

[8] For example, there are only 51 construction workers in the QES while there are 674 construction workers in the sample from the May 1977 CPS.

union membership. However, it is also clear that very little of the higher probability of unionization in the QES sample can be accounted for by shifts in labor-force structure. Even after accounting for differences in labor-force structure, the coefficient on the CPS dummy is still significantly less than zero (p value = .011).

The second panel of table 4 contains average predicted probabilities of union membership assuming first that all workers were in the QES survey (CPS = 0) and second that all workers were in the CPS survey (CPS = 1). The difference in these probabilities is also presented. The estimates in the first column verify the 4.2 percentage point (SE = 1.2) higher probability of union membership in the QES sample. If there were no difference in the within-sector probabilities of union membership between the two samples then the two averages in the second column (controlling for differences in labor-force structure) would be the same. However, their difference is 3.5 percentage points (SE = 1.4). Thus, approximately 20% of the difference between the two samples is due to measured differences in labor-force structure.

What this suggests is that for unknown reasons the QES has a higher fraction of union members than it should (using the CPS as the standard). There are two possible reasons for the difference. First, it might be that the QES sampling scheme oversampled union workers or some unmeasured characteristic correlated with union membership. This leads to a choice-based sampling problem like that induced by the quota sampling of union workers in the AFL survey. Second, it might be that the QES survey instrument was designed in such a way that some workers who were not union members responded that they were union members. This leads to problems of response error.

Each of these problems can be handled in a maximum-likelihood context, and they lead to identifiable differences in the likelihood function. However, without more information on the source of the difference between the samples, identification must depend heavily on untestable functional form assumptions. I proceed as if the problem is one of choice-based sampling, but the results are interpreted simply as a correction for the erroneously high probability of union membership in the QES sample.

A relatively straightforward approach is to use the May 1977 CPS as a standard and use a modification of the choice-based sampling estimator proposed by Hausman and Wise (1981) to adjust the probabilities for the QES sample so that they agree with the CPS. An overview of the procedure is that a modified probit model of union membership is specified using both the QES and May 1977 CPS data, but no sample indicator is included as a regressor. The model explicitly allows there to be a different probability of sampling a nonunion worker relative to a union worker in the QES than in the CPS. Differences between the QES and CPS in the probability of union membership are attributed to this differential sampling probability and are used to derive its estimate. This estimate is then used as if it were absolutely correct to determine how many union workers should be deleted from the QES sample so that its fraction union membership is representative (by the standard of the CPS). This last step is identical to the adjustment for the quota sampling of union workers in the AFL survey. The difference is that the degree of undersampling of union workers in the AFL survey is known, while the degree of undersampling of nonunion workers in the QES is unknown and estimated from a comparison with the May 1977 CPS.

Following Hausman and Wise (1981), consider a random variable y with density function $f(y)$. Assume that y is sampled with probability p_1 if y is less than zero, and that y is sampled with probability p_2 if y is greater than or equal to zero. In this case, the probability of sampling y if $y < 0$ relative to the probability of sampling y if $y > 0$ is $\theta = p_1/p_2$. The distribution of y in the resulting sample is

$$h(y) = \frac{\theta f(y)}{\theta \int_{-\infty}^{0} f(\omega)d\omega + \int_{0}^{\infty} f(\omega)d\omega} \quad \text{for } y < 0, \tag{9}$$

and

$$h(y) = \frac{f(y)}{\theta \int_{-\infty}^{0} f(\omega)d\omega + \int_{0}^{\infty} f(\omega)d\omega} \quad \text{for } y \geq 0. \tag{10}$$

Consider estimating the simple probit model of union status outlined in equations (1) and (2) so that workers are union members if $y > 0$. Using data from the QES and the May 1977 CPS, workers are undersampled if they are both from the QES sample and not union members. In this context,

$$\theta_i = 1 - \delta \cdot \text{QES}_i \tag{11}$$

for observation i, where QES_i is a dummy variable that equals one if the observation is from the QES sample and equals zero if the observation is from the CPS sample. The parameter δ represents the degree of undersampling of nonunion workers in the QES. If $\delta = 0$, then $\theta_i = 1$ for all observations. If $\delta > 0$, then $\theta_i < 1$ for the QES observations. The choice-based density function in equations (9) and (10) reduces to $f(y)$ for the CPS sample because θ equals one for these observations.

Assuming that $Y_i = X_i\beta + \varepsilon_i$, a standard normal distribution for ε_i, and the choice-based nature of the sample, the probability that a worker is unionized is

$$\mathrm{pr}(U_i = 1) = \frac{\Phi(X_i\beta)}{\theta_i[1 - \Phi(X_i\beta)] + (X_i\beta)}, \tag{12}$$

where θ_i is defined in equation (11). Similarly, the probability that a worker is not unionized is

$$\mathrm{pr}(U_i = 0) = \frac{\theta_i[1 - \Phi(X_i\beta)]}{\theta_i[1 - \Phi(X_i\beta)] + \Phi(X_i\beta)}, \tag{13}$$

and the log-likelihood function can be formed in a straightforward manner from these probabilities.

Maximum-likelihood estimates of the parameters (β and δ) of the choice-based probit model were derived using the May 1977 CPS and QES samples described in tables 2 and 3. As before, the vector X_i includes a constant plus 19 dichotomous variables representing main effects for four educational categories, five age categories, and the characteristics in tables 2 and 3. While the estimates of β are not presented here, their character is consistent with the differences presented in the tables. The key parameter for adjusting the QES probabilities is δ, which was estimated to be 0.174 with an asymptotic standard error of 0.0667.

Conditional on this estimate of δ, 17.4% (52) of the union workers in the QES sample were randomly selected and deleted. This adjusted sample is used in the remaining analyses, and it has 909 workers, of whom 27.1% are union members. This compares with 31.0% of the unadjusted QES sample and 26.7% of the May 1977 CPS sample. The sample characteristics of the adjusted QES sample are contained in the last two columns of table 3.

The decline in the probability of union membership between 1977 and 1984 computed using the adjusted QES and AFL surveys is 5.3 percentage points (p value = .007). This is virtually identical to the 5.4 percentage point decline computed using the CPS. The analysis now turns to decomposing this decline in union membership into components due to a decline in demand for union representation and to a decline in the relative supply of union jobs.

V. Identifying Demand for and Relative Supply of Union Jobs

The key information contained in both the QES and and the AFL survey is the response of nonunion workers to a question asking whether they would vote for union representation on their current job if a secret ballot election were held. The response to this question, called VFU here, is interpreted as an indication of the worker's demand for union representation. An affirmative response (VFU = 1) suggests that the worker feels he or she would be better off if the job were unionized ($Y_{1i} > 0$). Similarly, a negative response (VFU = 0) suggests that the worker feels he or she would be better off if the job were not unionized ($Y_{1i} < 0$).

The fraction of the nonunion sample that responded affirmatively to the VFU question fell 6.2 percentage points (p value = .011) from 38.6% in 1977 to 32.4% in 1984.[9] Thus, the demand for union representation among nonunion workers, $\mathrm{pr}(D = 1 | \mathrm{U} = 0)$, fell significantly between 1977 and 1984.

Neither survey asks the analogous question of unionized workers. However, since nonunion jobs are relatively freely available, it is assumed that all unionized workers are unionized because they prefer their jobs to be unionized and that they would answer the VFU question affirmatively ($Y_{i1} > 0$). This is clearly not completely accurate, and there are likely to be some unionized workers who would prefer their job not to be unionized. There is relevant evidence from the National Longitudinal Surveys of Young Men and Young Women that asks the VFU question of union members in 1980 and 1982 respectively. In samples constructed similarly to those used in this study, only about 11% of union members reported that they preferred their job not to be unionized.[10]

A worker is classified as demanding union representation ($D_i = 1$) if (1) he or she is unionized ($U_i = 1$) or (2) he or she is not unionized and responds affirmatively to the union preference question ($\mathrm{VFU}_i = 1$). The overall demand for union representation, $\mathrm{pr}(D = 1)$, is the sum of the fraction of the sample that is unionized and the fraction of the sample that is nonunion and responded affirmatively to the VFU question. On this basis, the demand for union representation fell 7.6 percentage points (p value = .011) from 55.2% in 1977 to 47.6% in 1984.

The supply of union jobs relative to demand, $\mathrm{pr}(U = 1 | D = 1)$, is measured by the fraction unionized of the workers who demand union representation. This fell only 3.2 percentage points (p value = .313) from 49.0% in 1977 to 45.8% in 1984.

These frequencies are summarized in table 5. With this information, the decline in union membership between 1977 and 1984 can be broken out into components due to the drop in demand and the drop in supply relative to demand. Taking the differential of equation (8) yields

[9] The 86 nonunion workers who were deleted from the AFL sample to correct for the quota sampling of union workers are included in this tabulation. Since union membership is not at issue in this analysis, there is no choice-based sampling problem. All analyses that involve strictly nonunion workers will continue to use the "full" nonunion sample.

[10] These data are not analyzed in detail here because they sample only a limited age range of workers (late twenties–late thirties). They are analyzed by Farber (1989).

Table 5
Summary Statistics for Adjusted QES and AFL Samples

	n	pr(U = 1)	pr(D = 1|U = 0)	pr(D = 1)	pr(U = 1|D = 1)
QES (1977)	909	.271	.386	.552	.490
AFL (1984)	996	.218	.330	.476	.458

$$\Delta \text{pr}(U_i = 1) = \Delta \text{pr}(D_i = 1) \cdot \text{pr}(U_i = 1 \mid D_i = 1) + \text{pr}(D_i = 1) \cdot \Delta \text{pr}(U_i = 1 \mid D_i = 1) + \Delta \text{pr}(D_i = 1) \cdot \Delta \text{pr}(U_i = 1 \mid D_i = 1). \tag{14}$$

The first term represents the effect of a change in demand, the second term represents the effect of a change in supply relative to demand, and the last term is a second-order term that can be ignored safely here. Based on the numbers in table 5 and using the average of the 1977 and 1984 levels, 3.6 points of the 5.3 percentage point drop in unionization are due to a decline in demand, and 1.6 points are due to a decline in supply relative to demand.

This analysis, which is essentially a comparison of means, ignores differences in labor-force structure across the two samples. The next step is to estimate multivariate probit models that can account for observable structural differences.

Table 6 contains estimates of a probit model of the probability that a worker demands union representation using the QES and AFL samples. This analysis is identical to an analysis of the probability of union membership except that the dependent dichotomous variable equals one not only for union members but also for nonunion workers who would vote for union representation.

The model in the first column of table 6 includes only a constant and the AFL dummy variable. The estimated coefficient on the AFL dummy is significantly less than zero, reflecting the fact that the demand for union representation was lower in 1984 than in 1977. The model in the second column additionally includes the 19 labor-force structure dummy variables used in the earlier analysis of union membership. Based on a likelihood ratio test, these labor-force structure variables clearly are significantly related to the probability that a worker demands union representation (p value $< .0001$). Even after controlling for labor-force structure, the estimated coefficient on the AFL dummy is significantly less than zero.

The second panel of table 6 contains average predicted probabilities of union membership assuming first that all workers were in 1977 (AFL = 0) and second that all workers were in 1984 (AFL = 1). The difference in these probabilities is also presented. The estimates in the first column verify the 7.6 percentage point (SE = 2.2) lower probability that workers

Table 6
Probit Model of Demand for Unionization, QES and AFL Samples, Selected Parameters

Variable	(1)	(2)
Constant	.131	Not reported
	(.0417)	
1984 (AFL)	−.192	−.173
	(.0576)	(.0618)
Labor-force structure	No	Yes
Log-likelihood (n = 1,905)	−1314.3	−1214.6
Mean probability assuming all:		
1977 (QES)	.552	.545
	(.0165)	(.0160)
1984 (AFL)	.476	.482
	(.0158)	(.0155)
Difference	−.0762	−.0632
	(.0221)	(.0227)

NOTE.—Labor-force structure includes a set of 19 variables representing main effects for sex, race, age (five categories), education (four categories), industry (six categories), and occupation (five categories). The numbers in parentheses are asymptotic standard errors. The all-QES mean probability is computed using the actual values of the labor-force structure variables for the combined sample assuming $AFL_i = 0$ for all observations. The all-AFL mean probability is computed using the actual values of the labor-force structure variables for the combined sample assuming $AFL_i = 1$ for all observations.

demanded union representation in 1984. If there were no difference in the within-sector probabilities of demand for union membership between the two samples, then the two averages in the second column (controlling for differences in labor-force structure) would be the same. However, their difference is 6.3 percentage points (SE = 2.3). Thus, only about 20% of the decline in demand for union representation between 1977 and 1984 can be accounted for by measured differences in labor-force structure.

Table 7 contains estimates of a probit model of the probability that a worker who demands union representation is, in fact, unionized. This analysis is identical to an analysis of the probability of union membership except that the model is estimated over the subsample of only those workers who either are union members or are nonunion but would vote affirmatively for union representation.

The model in the first column of table 7 includes only a constant and the AFL dummy variable. The estimated coefficient on the AFL dummy is not significantly different from zero, suggesting that the supply of union jobs relative to demand did not change significantly between 1977 and 1984. The model in the second column includes the 19 labor-force structure dummy variables. Based on a likelihood ratio test, these labor-force structure variables clearly are significantly related to the probability of that a worker who wants a union job is actually unionized (p value $< .0001$).

Table 7
Probit Model of Union Membership Conditional on Demand (Relative Supply), QES and AFL Samples, Selected Parameters

Variable	(1)	(2)
Constant	−.0250	Not reported
	(.0559)	
AFL	−.0810	−.0896
	(.0804)	(.0905)
Labor-force structure	No	Yes
Log-likelihood (n = 976)	−674.7	−560.9
Mean probability assuming all:		
1977 (AFL = 0)	.490	.489
	(.0223)	(.0202)
1984 (AFL = 1)	.458	.460
	(.0229)	(.0215)
Difference	−.0322	−.0292
	(.0320)	(.0304)

NOTE.—Labor-force structure includes a set of 19 variables representing main effects for sex, race, age (five categories), education (four categories), industry (six categories), and occupation (five categories). The numbers in parentheses are asymptotic standard errors. The all-QES mean probability is computed using the actual values of the labor-force structure variables for the combined sample assuming AFL_i = 0 for all observations. The all-AFL mean probability is computed using the actual values of the labor-force structure variables for the combined sample assuming AFL_i = 1 for all observations.

Controlling for labor-force structure does not change the conclusion that there is no significant difference in the relative supply of union jobs between 1977 and 1984.

The second panel of table 7 contains average predicted probabilities of union membership conditional on demand, assuming first that all workers were in 1977 (AFL = 0) and second that all workers were in 1984 (AFL = 1). The difference in these probabilities is also presented. The estimates in the first column verify the insignificant 3.2 percentage point (SE = 3.2) decline in the probability of union membership conditional on demand. The estimates in the second column again show an insignificant decline in the probability of union membership conditional on demand, this time controlling for changes in labor-force structure.

The analyses in this section suggest that most of the decline in union membership between 1977 and 1984 can be accounted for by a decline in demand that is not due to changes in labor-force structure. I turn now to alternative explanations for this decline.

VI. Increased Employer Resistance as an Explanation for the Decline in Unionization

The forms of employer resistance to unionization range from outright hostility to unions to the improvement of wages and/or working conditions so that workers will not feel they need unions. One key tactic is to hire labor-management consultants whose hallmark is defeating unions in representation elections. Freeman (1985) outlines three approaches that these consultants take. First, they can emphasize "positive labor relations" by having the nonunion employer provide a union-like environment that includes higher wages, better fringe benefits, workplace due process, and so on. Second, they can conduct a very active but legal campaign that includes much communication with workers regarding management's views of what unionization will mean for the workforce, gerrymandering of the unit of representation, and delay of the election itself. Finally, they can conduct an illegal election campaign by committing obvious unfair labor practices. There is evidence from data on individual votes in actual NLRB elections that very active legal campaigns and illegal campaigns have a significant influence on the outcomes of representation elections (Dickens 1983). In addition, there is evidence that simply delaying the election reduces the probability of union success significantly (Roomkin and Block 1981).

This set of behaviors by employers will have adverse effects on worker demand for union representation and, by extension, on the extent of unionization. The key direct evidence for an increase in employer resistance is the dramatic decrease in the quantity of union organizing success over the last decade coupled with the equally dramatic increase over the last decade in representation-election-related unfair labor practice charges filed by unions against employers.

Table 8 contains data on trends in union organizing activity. It is clear that election activity has declined in absolute terms. Both the number of NLRB-supervised representation elections and the number of nonunion workers who were eligible to vote in these elections has declined dramat-

Table 8
Union Representation Election Activity, Selected Years (1960–84)

Year	No. of Elections	No. of Workers in Elections (in thousands)	Nonunion Workers in Elections (%)	Elections Won by Union (%)	Unfair-Labor-Practice Complaints per Election
1960	6380	484.0	1.12	58.6	1.78
1970	8074	608.6	1.15	55.2	2.61
1975	8577	568.9	.97	48.3	3.64
1977	9484	570.7	.87	46.0	3.99
1978	8240	471.8	.67	46.0	4.81
1979	8043	577.9	.80	45.1	5.13
1980	8198	521.6	.71	45.7	5.37
1981	7512	449.2	.60	43.1	5.77
1982	5116	297.8	.40	40.3	7.45
1983	4405	209.9	.27	43.0	...

SOURCES.—Election and unfair labor practice data from various issues of National Labor Relations Board *Annual Report* (Washington, D.C.: U.S. Government Printing Office). Nonunion employment derived from (1) employment data from U.S. Bureau of Labor Statistics *Handbook of Labor Economics* (Bulletin no. 2070. Washington, D.C.: U.S. Government Printing Office, December 1980) and (2) membership data from Troy and Sheflin (1985).

ically since the mid-1970s.[11] The decline in organizing activity is even more extreme when measured relative to the size of the nonunion workforce. The percentage of nonunion workers who were eligible to vote in NLRB representation elections fell from 1.15% in 1970 to only 0.27% in 1983, the last year for which data are available.

This decline in organizing activity is consistent with increased employer resistance to union organizing for two reasons. First, to the extent that increased employer resistance takes the form of outright hostility to union organizing efforts, unions and workers will perceive a lower probability of success in organizing efforts. The result is that fewer elections will be undertaken. Second, to the extent that employer resistance takes the form of improved wages and working conditions and "positive labor relations," the measured demand for union representation among nonunion workers will be lower and there will be less election activity.

The penultimate column of table 8 contains data on union success in elections attempted as measured by the percentage of elections that are won by unions. This percentage fell from 55.2% in 1970 to 43.0% in 1983. This reflects the increased sophistication of employer responses to explicit organizing efforts. That employers are responding more aggressively to union organizing efforts is clear from the data contained in the last column of table 8, which shows that the number of employer unfair labor practice charges rose from 2.61 per election in 1970 to 7.45 per election in 1982. These unfair labor practices are a set of employers' activities that are proscribed under the National Labor Relations Act because they are felt to interfere with employees' rights to make free decisions regarding collective organization. Examples of these activities are threats, harassment, firing, and unduly pessimistic claims of what will result from unionization.

The reasons for this increased employer resistance to unionization are not clear. It may be that new and more effective tactics to resist union organization have been invented in a manner analogous to technological advancement in any production process. In particular, it has been argued casually that the advent of labor-management consultants is just this sort of event. However, it is more likely that the costs of unionization to firms have increased. This would provide firms with an incentive to resist unionization more strenuously than in the past, both by utilizing existing techniques for remaining nonunion and by "inventing" new and more effective techniques. Viewed in this context, the increased use of labor-management consultants is demand driven, and the discussion of increased employer opposition must start with a discussion of how the economic environment has changed with regard to the ability of unionized firms to compete successfully.

[11] Eligible workers are those who worked in potential bargaining units where elections were held.

The most obvious change in the U.S. economy over the past 2 decades is the increased level of foreign competition, particularly in the manufacturing sector that has formed the heart of the union movement in the United States. Some newly tabulated data on import penetration (Abowd 1987) illustrate this graphically. In 1958 only 2.5% of manufacturing sales in the United States were imports. This rose to 7.2% by 1977 and to 11.0% by 1984.

To the extent that unions raise production costs, some of this increase in imports is likely to be due to the unions themselves. However, it is also likely that other countries have rapidly developed industrial capacity that rivals (and in some cases surpasses) our own for reasons unrelated to unionization in the United States. In any case, it may be that, in the past, with no significant foreign competition, American firms could afford to accommodate higher costs associated with labor unions by sharing some of the gains of a relatively closed economy with their workers. However, the increased openness of the American economy results in a reduction of the gains available to be shared, because product prices that reflect the higher costs of unionized firms will not be borne by consumers who have attractive foreign alternatives.

Another structural change in the U.S. economy is the deregulation of some key heavily unionized industries such as trucking and airlines. These industries have become much more competitive since the government removed entry barriers and rate regulation.[12] In this more competitive environment, firms may resist unionization more strenuously than in the past because their market position is no longer protected by the government.

Some recent evidence on the relationship between the decline in union organizing activity and product market competition in U.S. manufacturing is mixed. Abowd and Farber (1987) argue that changes in product market competition, as reflected in changes in the total quantity of quasi rents available to be divided between the union and the employer, are an important determinant of the quantity of union organizing activity. They find that union organizing activity is positively related to the change in the total quantity of quasi rents but that there is still a substantial negative time trend to organizing activity that is not explained by changes in product market competition. They are unable to find any direct relationship between changes in import penetration and union organizing activity.

VII. The Decline in Demand for Unionization among Nonunion Workers

The decline in demand for unionization among nonunion workers is an important contributor to the overall decline in the demand for unionization.

[12] See Rose (1985, 1987) for analyses of the relationships among regulation, market power, and unions in the trucking industry. The problems of both the firms and the unions in the airline industry are common knowledge.

Evidence was presented in table 5 that the fraction of nonunion workers who desired union representation fell from 38.6% to 33.0% between 1977 and 1984. Further evidence of the decline in demand for union representation among nonunion workers is clear from the data on NLRB supervised representation elections discussed in the previous section and contained in table 8.

The possibility that the decline in the demand for unionization among nonunion workers can be fully accounted for by structural shifts in the labor force can be dismissed easily. Table 9 contains estimates of a probit model of the probability that a nonunion worker demands union representation. These estimates are derived using the full sample of 1,528 nonunion workers from the QES and AFL surveys.

The specification in the first column includes only a constant and the AFL dummy variable. The estimated coefficient on the AFL dummy is significantly less than zero, reflecting the fact that the demand for union representation fell significantly between 1977 and 1984. The model in the second column additionally includes the 19 labor-force structure dummy

Table 9
Probit Model of Demand for Unionization by Nonunion Workers, QES and AFL Data Selected Parameters

Variable	(1)	(2)	(3)	(4)
Constant	−.289	Not reported	.213	Not reported
	(.0494)		(.139)	
1984 (AFL)	−.168	−.126	.0056	.0353
	(.0664)	(.0714)	(.0709)	(.0760)
Satisfaction with job	...	...	−.637	−.671
			(.108)	(.113)
Satisfaction with pay	...	...	−.487	−.425
			(.0760)	(.0787)
Satisfaction with job security	...	...	−.278	−.270
			(.0868)	(.0914)
Perception that unions improve wages	...	...	.599	.612
			(.0943)	(.0990)
Labor-force structure	No	Yes	No	Yes
Log-likelihood (n = 1528)	−986.8	−916.2	−898.0	−838.5
Mean probability assuming all:				
1977 (QES)	.386	.375	.349	.344
	(.0189)	(.0181)	(.0175)	(.0169)
1984 (AFL)	.324	.332	.351	.355
	(.0159)	(.0158)	(.0156)	(.0156)
Difference	−.0624	−.0434	.00166	.0109
	(.0247)	(.0245)	(.0238)	(.0237)

NOTE.—Labor-force structure includes a set of 19 variables representing main effects for sex, race, age (five categories), education (four categories), industry (six categories), and occupation (five categories). The number in parentheses are asymptotic standard errors. The all-QES mean probability is computed using the actual values of the labor-force structure variables for the combined sample assuming $AFL_i = 0$ for all observations. The all-AFL mean probability is computed using the actual values of the labor-force structure variables for the combined sample assuming $AFL_i = 1$ for all observations.

variables. Based on a likelihood ratio test, these labor-force structure variables clearly are significantly related to the probability that a nonunion worker demands union representation (p value < .0001). After controlling for labor-force structure, the coefficient of the AFL dummy variable is significantly less than zero (p value = .077).

The second panel of table 9 contains average predicted probabilities of demand for union representation by nonunion workers assuming first that all workers were in 1977 (AFL = 0) and second that all workers were in 1984 (AFL = 1). The difference in these probabilities is also presented. The estimates in the first column verify the significant 6.2 percentage point (SE = 2.5) decline in the probability that a nonunion worker demands union representation. The estimates in the second column show that after controlling for changes in labor-force structure this probability fell to 4.3 percentage points (SE = 2.4, p value = .077). Thus, structural shifts can account for about 30% of the decline in demand for union representation among nonunion workers.

An important theme in an earlier literature on the demand for union representation is that workers join unions in order both to improve their wages and to protect themselves from what they feel is arbitrary treatment by their supervisors (Rees 1962). Seidman, London, and Karsh (1951), in their important study "Why Workers Join Unions" argue (p. 77) "that personal experiences in the plant play a large part in the thinking of workers, and that an unpleasant personal experience becomes a powerful motivation that turns workers toward a union." Following these arguments, Farber and Saks (1980) investigate the role in determining a worker's vote in a representation election played by (1) a worker's satisfaction with his or her job and (2) a worker's perception of the ability of unions to address problems on the job. They find strong support for the view that workers are more likely to vote for union representation when they are dissatisfied and feel that unions can improve conditions. These considerations suggest that at least part of the decline in demand for union representation among nonunion workers might be accounted for by an increase in job satisfaction and a deterioration in worker attitudes toward unions in general.

The QES and the AFL survey have comparable measures of job satisfaction in two key dimensions (pay and job security) as well as overall job satisfaction. The three measures of satisfaction were developed using a four-value response scale. These were recoded to two values (1 = satisfied, 0 = not satisfied).[13] The two surveys also have comparable measures of worker perceptions of the ability of unions in the abstract to improve wages and working conditions (union instrumentality). This was also re-

[13] The four possible responses to these questions (How satisfied are you with ———?) were (1) very satisfied, (2) somewhat satisfied, (3) somewhat dissatisfied, and (4) very dissatisfied.

Table 10
Job Satisfaction and Union Instrumentality, Sample Breakdowns by Union Status, QES and AFL Data

	Nonunion Workers		Union Workers	
	1977	1984	1977	1984
Fraction satisfied with:				
Overall	.867	.894	.879	.853
Pay	.587	.745	.748	.770
Job security	.729	.850	.762	.783
Fraction reporting unions improve wages and working conditions	.852	.757	.903	.917
n	663	865	298	217

coded from a four-value response scale to two values (1 = unions improve wages and working conditions, 0 = unions do not).[14] In both surveys, the questions referred to are worded virtually identically, and the allowed responses are scaled alike. While there may be problems due to the fact that the two surveys are different in overall structure, the properties of the samples are similar enough and the particular questions are similar enough to proceed with a comparison with some confidence.

The first part of table 10 contains information on the fraction of the workers in the QES and AFL samples that reported satisfaction in each of the three dimensions. It is clear from this table that nonunion workers reported very high levels of overall satisfaction with their jobs in 1977 and 1984, and that the fraction satisfied rose between these 2 years (p value of change = .097). The most striking result for nonunion workers in table 10 is that reported levels of satisfaction with pay and job security rose dramatically between 1977 and 1984. Both of these changes are statistically significant with p values < .001.

The analogous statistics for union members are included in table 10 in order to shed some light on the question of whether the increase in satisfaction among nonunion workers is likely to be an artifact of differences in survey design between the QES and the AFL survey.[15] In fact, the patterns for union workers are quite different than for nonunion workers. Union workers' overall satisfaction fell slightly between 1977 and 1984 while their satisfaction with the specific aspects of their jobs rose slightly. None of these changes are statistically significant at conventional levels (p value > .37 in each case). These findings suggest that that the results for the

[14] This question was, "Tell me if you (1) agree strongly, (2) agree somewhat, (3) disagree somewhat, or (4) disagree strongly that unions improve the wages and working conditions of workers."

[15] The 52 union workers who were deleted from the QES sample to correct for the oversampling of union workers are included in this analysis.

nonunion workers are unlikely to be an artifact of differences in survey design. If the higher levels of satisfaction among nonunion workers were due to some difference in the organization of the surveys or the precise wording of the questions, this sort of bias would surely show up among union workers as well.

The reasons for this increase in perceived job satisfaction of nonunion workers are not clear. Satisfaction with pay may reflect how workers evaluate their pay relative either to their best alternatives or to some norm that they consider equitable. Given the well-known stagnation in real earnings since the mid-1970s (Loveman and Tilly 1988; Kosters and Ross 1987), the general increase in worker satisfaction with pay then suggests that the standards against which workers judge their wages fell. The increase in satisfaction with job security presents a similar puzzle.

With regard to union instrumentality, the numbers in the second part of table 10 suggest that, while most nonunion workers still believe that unions improve the wages and working conditions of workers, the fraction of nonunion workers who believe that unions are effective in this dimension fell significantly from 1977 to 1984 (p value $< .001$). Thus, nonunion workers are less likely to believe that unions can help with a central area of concern on the job. Over the same period of time, the fraction of union workers who reported that they believe that unions improve wages and working conditions actually increased slightly.

It remains to demonstrate the links between worker preferences for union representation and these subjective measures of job satisfaction and union instrumentality. Table 11 contains, for each year, the fraction of nonunion workers who would vote for union representation broken down by satisfaction and perceptions of union instrumentality. It is clear that worker preferences for unionization are very strongly related to satisfaction and union instrumentality and that these relationships persist between 1977 and 1984. Each of the differences by satisfaction/instrumentality level

Table 11
Fraction of Nonunion Workers Who Would Vote for Union Representation Broken Down by Job Satisfaction and Union Instrumentality, QES and AFL Data, Showing Who Would Vote for Union Representation

	1977 (n = 663)		1984 (n = 865)	
	No	Yes	No	Yes
Satisfaction:				
Overall	.671	.342	.615	.289
Pay	.522	.291	.511	.259
Job security	.533	.331	.485	.295
Unions improve wages?	.204	.418	.181	.370

in the fraction who would vote for union representation is statistically significant with p value $< .001$.

The third column of table 9 contains estimates of a simple probit model of demand for union representation among nonunion workers that includes the four subjective variables. All of the satisfaction measures have a statistically significant effect (p values $< .001$) in the hypothesized direction on nonunion workers' preferences for union representation. Workers who are satisfied with their job are significantly less likely to demand union representation. The single measure of union instrumentality also performs as expected. Workers who feel that unions improve pay and working conditions are significantly more likely to desire union representation than workers who do not feel that unions are instrumental in this dimension.

Once satisfaction and instrumentality are controlled for, the hypothesis that the coefficient on the AFL dummy equals zero cannot be rejected at any reasonable level of significance (p value $= .937$). The predicted probabilities based on these estimates in the second panel of table 9 verify that there is virtually no difference in the predicted probabilities of demand for union representation by nonunion workers once these subjective factors are controlled for. Thus, it seems that all of the decline in nonunion workers' demand for union representation can be accounted for by an increase in measured job satisfaction and a deterioration in workers' perceptions of union instrumentality.

The estimates in column 4 of table 9 repeat the analysis of column 3 but include the additional 19 controls for labor-force structure. None of the findings are altered by this change.

In order to investigate the magnitude of the effects of job satisfaction on the demand for union representation, table 12 contains average predicted probabilities over the entire sample of 1,528 nonunion workers assuming particular configurations of the satisfaction variables for all workers. These are based on the estimates in column 4 of table 9, and the magnitudes are impressive. If it is assumed that all workers are not satisfied with their job overall (SAT = 0), the mean probability that a nonunion worker demands union representation is almost 23 percentage points higher (SE = 4.0) than in the case where it is assumed that all workers are satisfied with their job (SAT = 1). The difference is somewhat smaller, though still quite large for the other two measures of satisfaction. If it is assumed that all workers are not satisfied in any of the dimensions, the mean probability that a worker demands union representation is a dramatic 46.5 percentage points higher (SE = 4.1) than in the case where all workers are satisfied in all three dimensions.

Table 12 also contains mean-predicted probabilities assuming particular configurations of the union instrumentality variable. It is clear that worker perceptions of union instrumentality are quite important. If all workers are assumed to feel that unions improve wages and working conditions,

Table 12
Mean Predicted Probability of Demand for Union Representation, Nonunion Workers

Variable	Mean Probability Assuming Variable = 0	Mean Probability Assuming Variable = 1	Difference
Satisfied with job	.551	.323	−.227
	(.0373)	(.0121)	(.0398)
Satisfied with pay	.443	.303	−.140
	(.0217)	(.0141)	(.0268)
Satisfied with job security	.419	.332	−.0872
	(.0265)	(.0128)	(.0301)
Satisfied by all three measures	.724	.258	−.465
	(.0352)	(.0140)	(.0410)
Do not perceive that unions improve pay*	.384	.207	−.177
	(.0128)	(.0230)	(.0265)
Satisfied by all three measures and do not perceive that unions improve pay*	.766	.132	−.634
	(.0336)	(.0199)	(.0431)

NOTE.—All predicted probabilities were computed using the estimates in col. 4 of table 9 and the sample of 1,528 nonunion workers. The actual values of all of the variables except those manipulated in the table are used. The numbers in parentheses are asymptotic standard errors.
* The complement of the union instrumentality variable is used so that the measure has a relationship with the demand for union representation that is of the same sign as the relationships of the satisfaction measures.

the mean probability that a worker demands union representation is 17.7 percentage points higher (SE = 2.7) than in the case where no workers feel that unions improve wages and working conditions. This works quite powerfully in conjunction with the satisfaction measures. In the extreme case, where all workers feel that unions improve wages and working conditions and where all workers are not satisfied in any of the three dimensions, the mean probability that a worker demands union representation is 63.4 percentage points higher (SE = 4.3) than in the case where no workers feel that unions improve wages and working conditions and all workers are satisfied with their jobs in each of the three dimensions.

The conclusion from the analyses in tables 9–12 is that job satisfaction and perceptions of union instrumentality are *very* important factors in individual worker demand for union representation. The magnitude of the effects of these variables dwarfs the decline in demand between 1977 and 1984 summarized in table 8. This lends strong support to the views, expressed in the earlier literature and cited earlier in this section, that a central force motivating workers to demand union representation is "unpleasant personal experience" in Seidman, London, and Karsh's (1951) terms.

VIII. Concluding Remarks

The dramatic decline since the mid-1970s in the fraction of the labor force that is unionized is a phenomenon that is not yet fully understood.

Evidence was presented that showed that employer resistance to unionization has increased. Evidence was also presented that showed that the demand for union representation among nonunion workers has declined substantially since the mid-1970s. It was found that very little of the declines in these quantities can be accounted for by changes in the structure of the labor force.

One strong result that was found is that the decline in demand for union representation can be fully accounted for by an increase in the satisfaction of nonunion workers with their jobs and a decrease in their belief that unions are instrumental in improving wages and working conditions. However, the rationale for this is not obvious. Objectively, nonunion workers were no better off in 1984 than they were in 1977, but satisfaction levels increased. It may be that the economic dislocations of the 1970s and the increased competitiveness of the economy have reduced workers' expectations.

An important unresolved issue is exactly why employer resistance to unionization has increased so dramatically. One obvious answer is the increased competitiveness of U.S. markets, both in markets for traded goods and in previously regulated domestic markets. In this more competitive environment firms may feel that their survival is threatened by the higher costs associated with unionization in a way that it was not 20 years ago.

Another answer to the question of increased employer resistance is that the political and social climate may have changed so that the role that unions have played in American society and the economy is being called into serious question for the first time since that role was defined in the 1930s. Some recent work completed at MIT (Kochan, Katz, and Mckersie 1986) suggests the following scenario. Employers have never accepted unions as an integral part of their firms, but until the 1970s overt antiunion behavior was not socially or politically acceptable. The compact forged in the 1930s and codified as public policy in the National Labor Relations Act protected the union movement. In the 1960s employers began to implement effective strategies to remain nonunion when opening new plants. With the economic recessions of the 1970s and 1980s, more overt antiunion behavior became socially and politically acceptable, turning what had been a stagnation of the union movement into a virtual rout. Explicit antiunion strategies, including such tactics as development of innovative nonunion personnel systems, active resistance to organizing efforts, and locating plants in areas unsympathetic to unions have become the standard mode of operation in U.S. industry.

While the change in the strategy of employers could be thought to be the result of changes in social and political attitudes that arose independently of economic factors, it is reasonable to conclude that both employers' strategies and general attitudes toward unions have been affected by the dramatic changes in the U.S. economy over the past 2 decades.

What can the union movement do to recoup its losses? The results on the relationship between worker demand for union representation on the one hand and job satisfaction and union instrumentality on the other suggest that the task is to convince workers of the effective role that unions play in the workplace. However, it may be that, until workers feel less satisfaction with their jobs, this is a nearly impossible task. The recurring theme is that the competitiveness of the economy has increased dramatically. Unions need to convince workers that they offer real value in a competitive environment.

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