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TRADE UNIONS AND INCOME INEQUALITY

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Trade Unions and Income Inequality

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ABSTRACT

This paper examines the distributional impact of union wage gains by private-sector production workers in 1970. Quantitative estimates of the structure and impact of collective bargaining wage gains are developed along several dimensions of income: (a) the distribution of full-time year-round earnings, (b) the distribution of income among nonaged families and individuals, and (c) the incidence of poverty among nonaged families and individuals.

The first step of the analysis involves estimating the union impact on earnings. Using household data from the March 1971 Current Population Survey, I find that the proportionate effect of unions on earnings averages 18% for union workers, and varies from worker to worker depending on worker and job characteristics. Nonunion workers in industries with significant collective bargaining coverage are also found to reap significant "threat effect" wage gains.

Unions are found to compress the wage structure among union wage-earners and the income structure among union families. Summary measures of inequality increase by some 3-8% when union wage gains are deducted from earnings. When union and nonunion families are combined, and "threat effect" gains are introduced, the overall union impact becomes negligible.

Collective bargaining among production workers is not a major anti-poverty program, at least not in 1970. When union wage gains are deducted from earnings, the incidence of poverty rises by less than 3%.

These findings suggest that the distribution of union wage benefits among private-sector production workers does not have a significant impact on aggregate income inequality.

Trade Unions and Income Inequality

Do trade unions increase or decrease income inequality in the United States? This question has been debated for some time, in popular literature as well as in academic journals. Orthodox or neoclassical labor economists have generally argued that the economywide or macroeconomic effects of unions on income distribution are negligible. Thus, for instance, the ebb and flow of union membership in the twentieth century is not awarded a spot in orthodox debates over the sources of historical inequality trends (e.g., Budd, 1960; Williamson, 1976).

In particular labor markets, however, unions may have an impact on the level and structure of wages. Views differ, however, about what this impact is. Neoclassical economists tend to view trade unions as monopolies; as such, union success in extracting "rents" will hinge on the elasticity of the labor demand curve they face. Drawing on Marshall's Laws of Derived Demand, neoclassicals argue that the demand for skilled labor is generally less elastic than the demand for unskilled labor. It follows, then, that monopoly gains will be greatest for highly skilled workers, and not the less-skilled rank and file of the labor movement. The implication, then, is that unions widen the wage structure and increase earnings inequality.[1]

Thus, the view which has trickled down to textbooks and survey articles is that the overall effect of unions on earnings inequality is small, but probably disequalizing.[2]

The orthodox position, however, is not without its critics. Some researchers, for instance, see a union impact in distributive share trends (Kalecki, 1971, pp. 65-74; Levinson, 1954; Ozanne, 1959). Case studies in particular industries have also found significant union compression of wage structures (Reynolds and Taft, 1956; Reynolds, 1964). These case studies have not been contradicted by more recent econometric studies using household survey data. Nearly all such studies find that the relative wage gains of low-skilled "laborers" in unions are greater by far than those of highly skilled craft workers (Table 1). Finally, two very recent econometric studies find that unions exert a sizeable equalizing effect on measured wage inequality (Freeman, 1980; Hyclak, 1979).

Union compression of wage structures comes as no surprise to many institutional labor economists. In an effort to "take wages out of competition," unions usually attempt to reduce or eliminate interplant, interregional, and even international wage differentials.[3] The wage structure in the internal labor market may come under attack as well. Since unskilled workers are numerically dominant in most large industrial unions, wage benefits were delivered in equal-cents-per-hour across all skill classifications in many early contracts (Douty, 1953). Even when bargaining is along craft rather than industrial lines, coordinated craft bargaining has at times led to similar compression.[4] Finally, nearly all cost of living (COLA) clauses in collective bargaining agreements contain formulas which deliver benefits in equal-cents-per-hour

Table 1

Trade Union Impacts on the Occupational Wage Structure[a]

	Craft	Operative	Laborer	Sample
1. Weiss (1966)[b]				
CCR = 20	.14	.10	.13	1960 Census
CCR = 60	.09	.14	.15	
(four-firm concentration ratios)				
2. Stafford (1968)	.24	.26	.52	1966 Survey of Consumer Finances (Family heads who are not female, self-employed, farmers, retired, permanently disabled, students, or college graduates)
3. Ashenfelter (1972)	.107	.162	.252	1967 SEO (White urban males)
4. Boskin (1972)				1967 SEO (All persons 14+ working for wages and salaries)
All	.155	.152	.247	
White males	.155	.151	.236	
5. Ryscavage (1974)				1973 CPS (Full-time private wage and salary earners in nonfarm occupations
All	.292	.239	.359	
White males	.264	.211	.337	
6. Gay (1975)	.32	.40	.67	1967 SEO (Full-time males 14+ in mining, manufacturing, transport, communications, and public utilities)
7. Bishop (1976)	.19	.30	.38	1970 Census (Nonfarm, full-year workers)
8. Ashenfelter (1976)				
1967	.13	.14	.27	1967 SEO (White urban males)
1973	.23	.19	.34	1973 CPS (White urban males)
1975	.21	.20	.36	1975 CPS (White urban males)
9. Bloch and Kushkin (1978)[c]				1973 CPS (White, non-Spanish, prime-age males in manufacturing employment)
(a)	.255	.228	.616	
(b)	.156	.165	.334	
(c)	.195	.186	.384	

[a]The statistic reported in this table is the proportionate union wage gain:

$$M = \left(\frac{y^u - y^N}{y^N} \right).$$

In several studies, the reported values of M were disaggregated by industries, or occupations were disaggregated to greater detail. In these cases, 1970 employment weights were used to aggregate up to the statistics in this table.

[b]Percentage income advantage of a worker in industries with high collective bargaining coverage (U = 90) over those with low coverage (U = 50). CCR is an adjusted four firm concentration ratio.

[c]Bloch and Kushkin regress separate equations for union and nonunion workers. Rows (a) and (b) "standardize" workers across the sectors using union and nonunion coefficients, respectively. Row (c) reports coefficients from a pooled regression of union and nonunion workers.

to all workers; such clauses thus mechanically translate inflation into wage compression.[5]

Conflicting arguments and evidence thus make firm conclusions as to the direction and magnitude of the impact of unions on inequality problematic. This study makes a contribution to this debate by providing quantitative estimates of the distribution of union wage benefits to private-sector, blue-collar wage earners and their families. The union effect on family incomes is particularly significant, since studies to date have focused only on factor shares or wage inequality. This study, while providing an alternative estimate of the union impact on earnings inequality, also estimates the union impact on family income inequality, which is, perhaps, a more important dimension of inequality.

Before delving into statistical analysis, I think it will be useful to explain the organization of this paper. Section I provides estimates of the union impact on the annual earners of both union and nonunion production workers. In Section II, I combine data on the location of union and nonunion production workers in the earnings distribution with estimates of the distribution of union wage gains to yield quantitative estimates of the union impact on the distribution of full-time earnings. In Section III, I employ a similar methodology to estimate the union impact on the distribution of family and individual incomes. In Section IV, I estimate the importance of union wage gains in maintaining family incomes above the poverty line. Finally, Section V provides a summary of findings, some conclusions, and some caveats.

I. UNIONS AND EARNINGS

In gauging the union impact on income inequality, the first step is to estimate their impact on the level of annual earnings. This is done by estimating the following earnings equations:

$$\log (Y_i^u) = B_1^u X_i^u + B_2^u CB_i^u + e_i^u \qquad (1)$$

$$\log (Y_i^n) = B_1^n X_i^n + B_2^n CB_i^n + e_i^n \qquad (2)$$

where the "u" and "n" superscripts refer to union and nonunion workers respectively. The dependent variable in each case is the log of annual earnings. X is a vector of worker and job characteristics, and CB is the share of production workers in this worker's industry who are covered by collective bargaining agreements. The variables in X (which are listed at the bottom of Table 2) are meant to capture variation in worker productivity, nonpecuniary aspects of jobs, and regional living costs. Proxies being what they are, there is surely significant uncaptured variation; we assume that this unmeasured variation is uncorrelated with the proxies in X or with CB and is impounded in a well-behaved residual (e). Race and sex dummy variables were included in these regression equations in such a manner as to allow the experience-earnings profiles of the different race and sex groups to vary.

The separation of union and nonunion workers is justified by our belief that the wage-determination process in the structured internal labor markets which characterize the union sector is not a mirror-image of the wage determination process in the less-structured nonunion sector.

Table 2

Selected Regression Coefficients[a]

	OLS Regression Coefficients (t-ratio)		
	Pooled	Union	Nonunion
Union membership	.1185[b] (14.21)	-	-
Collective bargaining coverage	-		
21-40%	.0681[b] (4.57)	.0102 (.439)	.0530[b] (2.66)
41-60%	.0957[b] (6.59)	.0783[b] (3.78)	.0592[b] (2.85)
61-80%	.1526[b] (9.99)	.1335[b] (6.43)	.1082[b] (4.80)
81-100%	.1741[b] (11.62)	.1441[b] (7.28)	.1737[b] (7.22)

[a]The dependent variable is the log of annual earnings. The other independent variables are Sex, Race, Experience; $(Exp)^2$; $(Exp)^2$ x Sex; $(Exp)^2$ x Race; Head; Viet. Vet.; Other Vet.; Cent. City; Other City; Rural; 5 Industry Dummies; 3 Occupational Dummies.

[b]Significant at .005 level of confidence.

A single equation specification combining both union and nonunion workers imposes such an implausible restriction.

As noted above, CB is the percentage of production workers in the industry who are covered by collective bargaining agreements. The ability of unions to secure enduring wage gains from an employer will be jeopardized if there is a substantial competing nonunion sector in an industry. If the industry is well organized, however, competitive wage pressures are reduced and unions' bargaining leverage--and hence wages--will increase. Thus higher levels of coverage will lead to higher wages for union members. Our expectation, then, is that B_2^u is positive.

But CB also affects the wages of the unorganized. Evidence suggests that nonunion employers in highly organized industries often raise wages in tandem with the unionized segment of the industry as a defensive strategy to prevent unionization.[6] This leads us to expect a positive sign for B_2^N as well. This effect of unions on nonunion wages is hereafter referred to as the "threat effect."

For ease of exposition, CB is assumed to have a linear effect in equations (1) and (2). Some have suggested, however, that industry coverage effects, if they exist at all, are not likely to be linear (Rees, 1979, p. 143). For this reason, in the actual estimation of equations (1) and (2) below, CB is a set of four dummy variables representing quintiles of collective bargaining coverage.[7]

A large data file which provides demographic information on workers, annual earnings, and union membership is the March 1971 Current Population Survey.[8] Since we are concerned with the distributive effects of blue-collar unions, a hefty sample of 13,413 private-sector full-time, year-round production workers was drawn from this survey. Of this group 5,636 (42 percent) belonged to unions. Measures of collective bargaining coverage by industry were obtained from Freeman and Medoff (1979).

Ordinary least squares estimation of the coefficients in equations (1) and (2) yielded statistically significant coefficients for 7 of the 8 CB dummy variables (see Table 2). As hypothesized, the coefficients are positive and increasing in both equations.[9]

Before examining the size and structure of the proportionate union wage effects implied by the estimated coefficients, one must first determine whether the wage determination processes in the union and nonunion sector do, in fact, differ. Specifically, this means testing whether the coefficients in equations (1) and (2) are equal. If equality is rejected, the assumption of differing wage determination processes is affirmed. Two null hypotheses were tested: (a) pair-wise equality of all of the coefficients; and (b) pair-wise equality of all coefficients except intercepts. Both null hypotheses were rejected.[10]

The total effect of unionism on a worker's earnings will include the effect of both industry coverage and union membership. Consider two identical workers—one a union member and the other nonunion--in an industry with a given level of collective bargaining coverage (CB). From equations (1) and (2) the total proportionate effect of unionism on the worker's wages is

$$e^{(B_1^u - B_1^n)X + B_2^u CB} - 1,$$

for the union worker and

$$e^{B_2^n CB} - 1$$

for the nonunion worker, where X is a vector of the workers' (identical) characteristics. The proportionate union wage effect will thus depend upon worker and job characteristics: different Xs yield different benefits.

The effect of unions on earnings can be determined by choosing a value of X and allowing CB to range from 0 to 100% (see Table 3). The following discussion assumes X characteristics of an average union member. In highly organized industry (81-100% coverage), such a union member earns some 43% more than an identical nonunion worker in unorganized industries. This is a sizeable premium. But the earnings premium for nonunion workers in well-organized industries--the "threat effect"--is also substantial. A nonunion worker in an industry which is 85% organized earns 21% more than an identical worker in an industry which is largely nonunion. Clearly, there are many nonunion workers who reap the benefits of collective bargaining.[11]

II. UNIONS AND EARNINGS INEQUALITY

The impact of blue-collar trade unions on earnings inequality involves an assessment of their effect on the share of total earnings accruing to workers in various earnings classes. This impact, in turn, hinges on two variables: (1) the proportion of workers in an earnings class who are union members or who reap "threat effect" gains; and

Table 3

Proportionate Union Wage Effects

(1) Industry Collective Bargaining Coverage (%)	Earnings Index (2) Nonunion	(3) Union	(4) Ratio Col. (3)/Col. (2)
0-20	100	106.6	1.066
21-40	105.5	109.1	1.034
41-60	106.3	118.2	1.141
61-80	117.4	121.1	1.032
81-100	121.1	143.2	1.182

(2) the average size of the wage gains. The first variable requires knowledge of the distribution of union and nonunion earners, and the second, the distribution of relative wage gains by earnings classes. We begin with the distribution of union and nonunion workers.

The most thorough profile of the relative position of the two groups of workers in the earnings distribution is contained in Appendix Table A, which presents the joint relative frequency distribution of full-time, year-round earnings in 1970. This table provides a detailed breakdown of the location of both union and nonunion production workers in the earnings distribution.

Rather more illuminating, however, is the grouped distribution in Table 4. If the three earnings groups--union production, nonunion production, and nonproduction workers--were similarly distributed, 20s would appear in all the cells of the table. In fact, union workers are underrepresented in the lower and upper quintiles and overrepresented in the upper-middle quintiles. Nonunion production workers, however, are concentrated in the lowest three quintiles. Since the "nonproduction" group is such a mixed lot--including low-wage clerical and sales workers as well as professional, technical, and managerial workers--the uniformity of their distribution is hardly surprising; nor is their "bulge" in the highest quintile.

In the previous section we found that the relative wage gains conferred by union membership varied from worker to worker. Other studies reach a similar conclusion and have compared the union relative wage effect for various demographic groups (Boskin, 1972; Ashenfelter, 1972).

Table 4

Grouped Distribution of Full-Time, Year-Round
Wage and Salary Workers

Population Earnings Quintile	Production Workers: Union (%)	Production Workers: Nonunion (%)	Nonproduction Workers (%)
Lowest fifth	10.3	33.6	16.6
Second	17.0	23.7	19.2
Third	25.4	19.0	18.2
Fourth	30.4	14.8	18.7
Fifth	16.9	8.9	27.3
	100.0	100.0	100.0

No one to date, however, has explicitly calculated the structure of relative wage gains by income class. Such a calculation allows comparison of wage gains to workers in the lower earning deciles to those in upper deciles. If the average proportionate wage gain falls as earnings rise, unions exert an equalizing effect on the distribution of earnings; on the other hand, if proportionate wage gains rise with earnings, we infer a disequalizing union impact.

The methodology employed in constructing such a distribution of relative wage gains involves a straightforward transformation of earnings equations (1) and (2) in Section I. The proportionate union wage effect is

$$M = \frac{\bar{y} - \bar{y}^*}{\bar{y}} ,$$

where $\bar{y}$ is the actual average earnings of workers in a given earnings class and $\bar{y}^*$ is the earnings of these workers in the absence of collective bargaining. Thus defined, M is the proportion by which workers' wages would change, were the gains brought about by collective bargaining to disappear.

M is readily calculated for any income class by using equations (1) and (2). These computed values are displayed in Table 5 and Figure 1. The horizontal axis in Figure 1 ranks workers by earnings quintiles. Among unionists, the largest relative wage gains are received by workers earning the lowest and highest wages. The average proportionate wage gain falls as income rises throughout the lowest 80% of the distribution,

Table 5

Relative Wage Effects by Earnings Class

Earnings Quintile[a]	Production Workers		
	Union (%)	Nonunion (%)	All Workers (%)
Lowest fifth	20.5	3.7	4.0
Second	19.0	5.3	5.2
Third	17.9	6.2	6.6
Fourth	17.2	7.2	7.1
Fifth	19.3	8.2	4.6
Average	18.2	5.5	5.5

[a]This denotes quintiles of the distribution of all wage and salary earners. Thus, the average relative wage effect for union production workers who fall in the lowest fifth of <u>all</u> wage and salary earners is 20.5%.

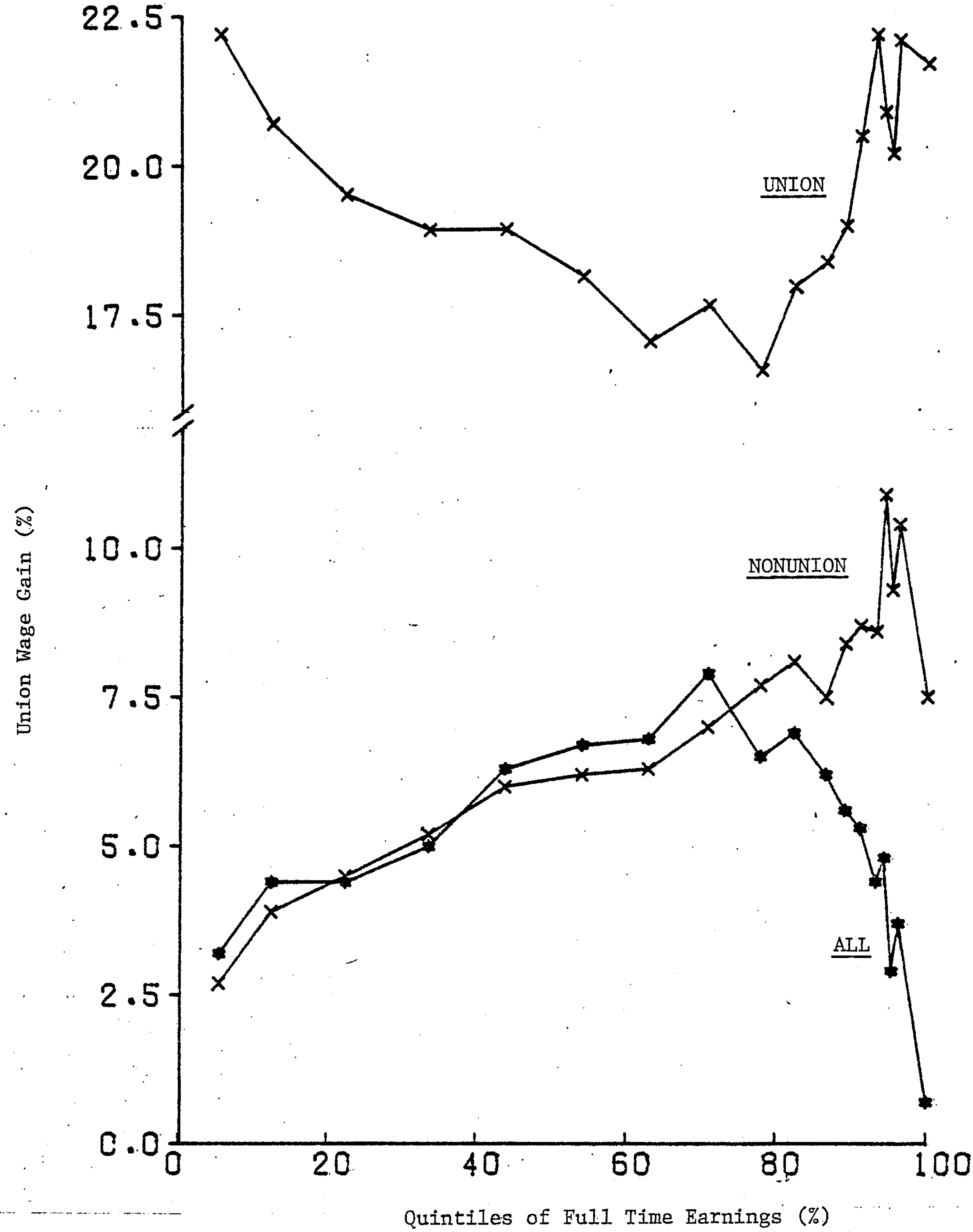

Figure 1. The Distribution of Union Wage Gains

rising sharply thereafter. The average gain for unionists is approximately 18%. Figure 1 reveals that approximately the lowest 50% and the top 10% of the unionists are above this mean; all others are below it.

In Section I it was found that unions exerted a significant "threat effect" on the wages of nonunion production workers in their industry. The structure of these nonunion gains is also displayed in Figure 1. Their structure would seem to impart a small disequalizing impact on nonunion production earnings. Proportionate wage gains rise more or less continuously from under 4% for the lowest quintile to a bit over 8% in the top earnings quintile.

The final set of statistics in Figure 1 combines production workers with nonproduction workers to yield a set of average gains among all full-time workers. This graph is not readily deduced from the preceding graphs. The differences in average proportionate wage gains across earnings classes derive from two sources: (1) different structures of relative wage gains among union and nonunion production workers; and (2) shifting weights of union production workers, nonunion production workers, and nonproduction workers. The net result of all of these changes is the "inverted-U" in Figure 1.

What do these wage effects tell us about earnings inequality? An answer, or rather a set of answers, is provided in Table 6, which presents various summary measures of earnings inequality in the presence and absence of union relative wage gains. The rows labeled "Actual" were constructed by combining the joint relative frequency distribution in Appendix Table A with the mean value of earnings in each cell of the

Table 6

Union Impacts on Summary Measures of Earnings Inequality[a]

	Gini	Log. var.[b]	Atkinson[c] e = 1.5	Atkinson[c] e = 2.5	Atkinson[c] e = 4.0
A. All workers					
Actual	.306	.388	.232	.425	.708
Minus union wage gains	.312	.392	.235	.423	.703
Union impact (%)	-1.9	-1.0	-1.3	.5	.7
B. Union workers					
Actual	.207	.181	.118	.245	.560
Minus union wage gains	.205	.187	.120	.256	.581
Union impact (%)	-1.0	-3.2	-1.7	-4.3	-3.6

[a]Full-time, year-round earnings.

[b]The variance of the log of income, i.e., $I \sum_i \sum_j \log^2 (y_{ij}/\bar{y})\, f(y_{ij})$. $i = 1, \ldots, 19$ earnings classes and j = union, nonunion, nonproduction.

[c]The formula for the Atkinson Index is

$$I = 1 - \left(\sum_{ij}\sum (y_{ij}/\bar{y})^{1-e} f(y_{ij})\right)^{1/1-e}$$

where e represents distributional value judgments. Higher values of e reflect greater concern for redistribution at the lower end of the distribution. Thus a 5% relative wage gain to workers in the lowest earnings decile will have a greater impact on measured inequality for e = 4 than for e = 1.5. It should be noted, however, that the Log variance and Atkinson measures all weight transfers in the lower tail of the distribution more heavily than those in the upper tail (higher values of e simply increase the bias).

For further discussion of the properties of these inequality indexes see Atkinson (1970) or Williamson (1976).

same table. The counterfactual values, that is, inequality in the absence of the union wage changes, were constructed by deducting the proportionate impacts shown in Figure 1 from the group means and then recomputing the inequality index. The impact of unionism is thus gauged by comparing the "Actual" with the second row of each column. If the number in the second row is bigger than the first, the union relative wage gains have reduced income inequality.

Looking first at Panel A, it can be seen that the union relative wage gains exert a small equalizing impact on most measures of overall earnings inequality. For instance, earnings inequality as measured by the Gini coefficient is reduced approximately 2% by blue-collar union wage gains. The dissenting evaluations associated with the larger values of the Atkinson parameter (e) reflect a greater sensitivity to transfers in the lower tail of the distribution. Recall from Figure I that the average effect of unionism among all workers (i.e., both production and nonproduction) is disequalizing in the lower 60-70% of the earnings distribution, and equalizing thereafter. The Atkinson measures in the right-hand column of Table 6 are progressively more closely attuned to the disequalizing effect, and less sensitive to the equalizing effect.

Panel B measures the impact of union wage gains on inequality among union production workers only. As Figure I suggests, their impact is small, but somewhat more equalitarian than for the entire population.

In sum, the largest average benefits from collective bargaining accrue to workers in the middle range of the earnings distribution. This is not because the gains from unionism are greater for workers in this middle group; on the contrary, the relative wage gains of the low-wage union members are well above average. The problem, however, is that the likelihood of union membership is greatest for production workers in the middle range of the earnings distribution. These offsetting effects make for a very small, equalizing union effect.

III. UNIONS AND THE SIZE DISTRIBUTION OF INCOME AMONG FAMILIES AND INDIVIDUALS

As with the earnings distribution, the first step in determining the union impact on family income inequality involves locating union families and individuals in the distribution of income among all families and individuals. A "union family" is defined as any family with a blue-collar union member. Next, the distribution of proportionate family income gains brought about by collective bargaining is estimated. As was the case with earnings, if the average proportionate union income effect falls as income rises, we infer an <u>equalizing</u> effect on income distribution; similarly, a positive relationship implies a <u>disequalizing</u> impact.

A detailed joint relative frequency distribution of family incomes is presented in Appendix Table C. More useful for our purposes, however, is the grouped distribution in Table 7. Union families are very much underrepresented in the first quintile; in the remaining quintiles, however, they are spread fairly uniformly. Nonunion families are spread fairly evenly among all the quintiles.

Table 7

Grouped Distribution of Money Income: Families and Individuals[a]

Income Quintile	Union Families (%)	Nonunion Families (%)
Lowest fifth	6.7	23.4
Second	26.0	19.8
Third	20.5	18.4
Fourth	25.9	18.1
Fifth	20.9	20.3
	100.0	100.0

[a]Civilian, nonaged families and individuals.

Do the relative income gains brought about by collective bargaining display similar uniformity? Using family income data in the CPS data file, we can translate the proportionate wage gains estimated in Section I into proportionate income gains for workers' families. Let Δy denote the dollar earnings gain brought about by collective bargaining, y and YFAM the observed level of earnings and family income, respectively. It follows that the proportionate income gain brought about by collective bargaining (Δy/YFAM) may be written as follows:

$$\frac{\Delta y}{YFAM} = \frac{\Delta y}{y}\left(\frac{y}{YFAM}\right) = MS,$$

where M is the proportionate wage gain brought about by unions and S is the share of family income resulting from production earnings. Relative income effects computed in this manner are presented in Table 8 and Figure 2. (The income shares used in these calculations and a disaggregated distribution of relative income effects are presented in Appendix Tables D and E, respectively.)

Families are ranked by income on the horizontal axis in Figure 2. The size of the union relative income effect falls continuously from a peak of approximately 18% for the lowest 30% of families and individuals down to approximately 12% for the top income decile. Thus, unlike the structure of union earnings gains, the structure of income gains among union families is consistently equalizing.

In Section II we found that "threat effects" of unions had a disequalizing impact on the distribution of nonunion production earnings. Do they have a similar impact on the distribution of income among

Table 8

Relative Income Effects by Income Quintiles

Quintile[a]	Union (%)	Nonunion (%)	All (%)
Lowest fifth	18.1	1.0	2.2
Second	16.7	1.7	4.9
Third	16.2	1.6	5.5
Fourth	15.3	1.3	5.3
Fifth	12.1	.6	2.7
Average	15.4	1.2	4.1

[a]This denotes quintiles of the distribution of all families and individuals. Thus, the average income effect for union families who are in the lowest fifth of all families is 18.1%.

families and individuals? Figure 2 suggests not. While the average effect on nonunion family incomes is quite small (only 1.2% of family income), the gains are structured, for the most part, in an equalizing manner. Nonunion families in the middle three quartiles are the largest beneficiaries of these gains.

When the gains of union and nonunion families are averaged, the result is the "inverted-U" in Figure 2. Relative income gains rise to a peak by the fourth income decile, then fall more or less continuously thereafter. The largest effects, that is those in excess of 4%, are received by families and individuals in the middle 60% of the distribution, with the smallest gains going to families and individuals in the upper and lower tails of the distribution.

The implication of these relative income effects for summary measures of income inequality are displayed in Table 9. The first row of the table presents inequality measures for the observed distribution of income. The second row gives similar statistics for a distribution which deducts the union relative wage gains.

In passing from row 1 to row 2 in Panel A, the Gini measure registers a small increase in inequality, implying that unions reduce measured inequality by 1.6%. On the other hand, the log variance and Atkinson measures register slight increases in inequality. What explains the different verdicts? Note in Figure 1 that while the structure of union relative wage gains is sharply equalitarian over the top 30-40% of incomes, it is mildly disequalizing in the lowest income deciles. As

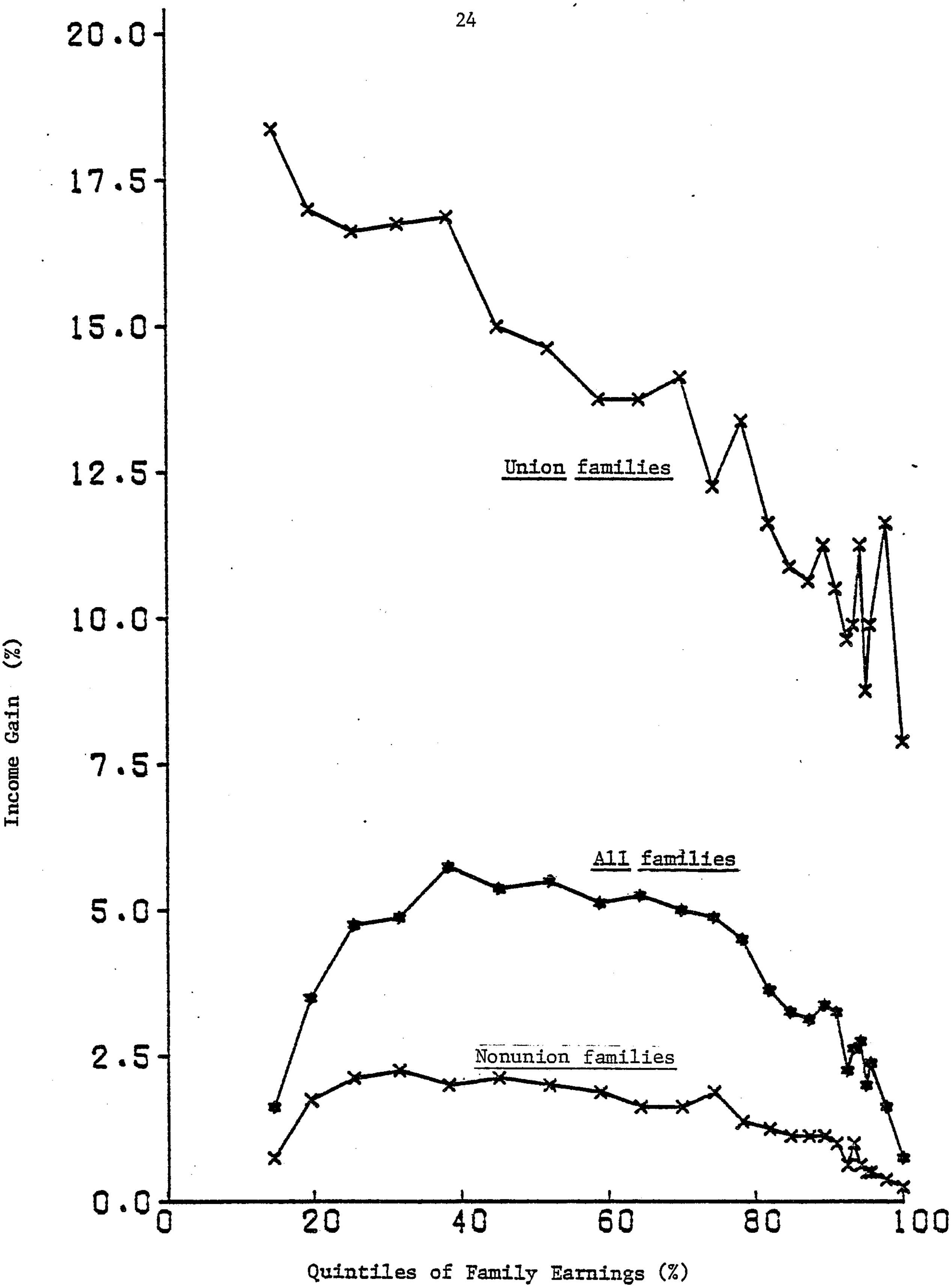

Figure 2: THE DISTRIBUTION OF UNION WAGE GAINS

Table 9

Union Impacts on Summary Measures of Income Inequality Among Nonaged Families and Individuals

	Gini	Log Var.	Atkinson e = 1.5	Atkinson e = 2.5	Atkinson e = 4.0
A. All families					
Actual	.367	.573	.315	.484	.629
Minus union wage gains	.373	.571	.316	.479	.620
Union impact (%)	-1.6	.4	-.3	1.0	1.5
B. Union families					
Actual	.248	.228	.146	.249	.395
Minus union wage gains	.258	.246	.157	.264	.414
Union impact (%)	-3.9	-7.3	-7.0	-5.7	-4.6

noted earlier, the Atkinson measures are more sharply attuned to the latter effect than is the Gini coefficient. Thus, the overall impact of union gains on family income inequality is ambiguous. Inequality measures that weight transfers in the upper and lower tails of the distribution more equally, however, register a slight decrease in inequality.

The impact of union wage gains on the distribution of income among union families, however, is unambiguously equalitarian: all measures of inequality are reduced.

IV. UNIONS AND POVERTY

Poverty, as measured in government statistics, depends not only on the level of family income, but also on family need. In 1977, for instance, the poverty income line for a nonfarm husband-wife family with three children was $6143 dollars. The poverty line for a single nonaged male (with no children) was $3267, or approximately 53% of the former (U.S. Bureau of the Census, 1979, pp. 203-209). It is possible to determine the poverty status of a family, and in doing so construct a rough income equivalence measure, by simply dividing a family's actual income by the poverty line appropriate for that family. This constructed ratio will be termed "POVRATIO." For instance, if our husband-wife family had an income of $9215 and the single male had an income of $4822, both would have a POVRATIO index of 1.5 and both would be equally "well-off" in terms of this rough welfare index.

Since the CPS file provides the poverty-line values for all families and individuals, it is possible to construct POVRATIO values for families in our sample. The results of such an exercise are displayed in Table 10. By construction, any family with a value of POVRATIO less than one is poor. Thus, the first entry in the "Union" column tells us that less than one percent of all families are "union" and poor. Similarly, 5.5% of poor families (.6/10.8 x 100) include a blue-collar trade unionist, and 2.4% (.6/20.3 x 100) of all union families are poor. We conclude then that in 1970 poverty was not common among union families and unionism was not common among the poor.

Table II displays the computed relative income effects by POVRATIO class. (For supporting statistics see Appendix Table F.) As was the case with family income, the structure of union gains is equalizing. The structure among the nonunion group and among all families displays the familiar "inverted-U" shape.

Do the income gains brought about by collective bargaining play an important role in reducing the incidence of poverty? The evidence in Tables 10 and 11 suggests that they do not. Suppose that the gains of union and nonunion production workers are fully at the expense of families in the highest income classes and that families are uniformly distributed within any POVRATIO class. If we further assume that any effects on factor supply brought about by collective bargaining are negligible, we can readily calculate poverty rates in the absence of collective bargaining by simply subtracting out our computed relative

Table 10

Joint Relative Frequency Distribution by POVRATIO Classes

	Family Type		
POVRATIO[a]	Union (%)	Nonunion (%)	All (%)
0	.6	10.3	10.8
1	1.3	7.1	8.4
1.5	2.4	8.6	11.0
2	5.9	17.4	23.3
3	4.8	14.9	18.7
4	5.3	23.4	27.7
All	20.3	80.7	100

[a]Lower bound.

Table 11

Relative Income Effects by POVRATIO Class

	Family Type		
POVRATIO[a]	Union (%)	Nonunion (%)	All Families (%)
0	17.3	1.0	1.7
1	16.4	1.7	3.9
1.5	15.6	1.8	4.8
2	15.2	1.7	5.1
3	14.7	1.2	4.7
4	12.6	.7	3.0

[a]Lower bound.

income effect. The results of such an exercise are displayed in Table 12. The numbers in the first and second columns of the table are the joint probability estimates from Table 10 (e.g., .6% of all families are union and poor). The second row of Table 12 presents the counterfactual joint probability value. In the absence of union wage gains, the number of poor union families increases by one-third. The effect of collective bargaining on all poor families, however, is rather small: the number of poor families and individuals increases by less than 3% if union wage gains are deducted from family income.

V. CONCLUSION

The wage gains of private-sector blue-collar unionists do not, it seems, have a large impact on income inequality. This study finds that the largest average gains from collective bargaining accrue to workers in the middle of the distribution of annual earnings. The impact of this redistribution is slightly equalitarian, although this interpretation depends upon the inequality measure used. In any event, the effect is quite small.

An interesting finding is that the impact of unions on family income inequality is larger than their impact on earnings inequality. The structure of gains among union families and individuals is equalitarian: absent the union gains, measured inequality is increased by from 4 to 7%. Among all families, the largest income gains accrue to workers in the middle 60% of the income distribution. This results in an ambiguous

Table 12

Percentage of all Families and Individuals
below Poverty Line

	Union (%)	Nonunion (%)	Total (%)
Actual	.6	10.3	10.9
Minus union wage gains	.8	10.4	11.2
Percent change	33.0	1.0	2.8

effect on summary measures of income inequality. Regardless of the inequality metric chosen, however, the union impact on family income inequality is less than 2% in either direction.

Union wage gains do not play a major role in reducing poverty. This is not because low-wage workers do not receive substantial benefits from collective bargaining: a consistent finding in this study is that the proportionate wage and income gains for low-wage workers are significantly above the average for all production workers. The problem is that relatively few union members are in families that are below or near the poverty line. My finding is that if the wage gains from collective bargaining are deducted from family incomes, the proportion of poor families and individuals in the population increases by less than 3%.

A vexing knot encountered in unraveling union distributional impacts is the incidence of union wage gains. In this study I have only examined the distribution of benefits of collective bargaining. If one views union distributional impacts in the framework of a zero-sum game (i.e., what unionists gain, others must lose), then a full account of the distributional impacts must also examine the pattern of costs which they impose. This rather formidable task is left for future studies.[12]

In closing, I should like to point out that this study has focused only on direct labor market impacts. The U.S. labor movement, in addition to being an important labor market institution, is an important political institution as well, and the redistribution brought about by political lobbying may well dwarf the direct labor market effects of unions. Since discarding the doctrine of "voluntarism" in the 1930s,[13] the American labor movement has been an important political lobby for social insurance and labor standards legislation, economic regulation and public higher education, all of which have had important distributional consequences. The estimates in this paper thus represent but one tile in a larger distributional mosaic.

APPENDIX

Table A

Joint Relative Frequency Distribution of Annual Full-time, Year-round Earnings

Earnings Class[a]	Production Workers: Union	Production Workers: Nonunion	Nonproduction Workers	All Workers
$0	.004	.032	.019	.054
$3,000	.008	.037	.027	.071
$4,000	.013	.042	.046	.100
$5,000	.018	.039	.054	.111
$6,000	.023	.033	.047	.103
$7,000	.028	.029	.046	.102
$8,000	.027	.022	.039	.088
$9,000	.028	.018	.033	.078
$10,000	.022	.014	.036	.072
$11,000	.013	.009	.022	.044
$12,000	.011	.008	.023	.042
$13,000	.006	.004	.015	.026
$14,000	.004	.003	.013	.020
$15,000	.003	.003	.015	.021
$16,000	.002	.001	.008	.011
$17,000	.001	.001	.008	.010
$18,000	.001	.001	.008	.009
$19,000	.001	.001	.037	.039
Total	.211	.296	.493	1.000

[a]Lower bound.

Table B

Relative Wage Effects by Earnings Class

Earnings Class[a]	Production Workers Union	Production Workers Nonunion	All Workers[b]
$0	.222	.027	.032
$3,000	.207	.039	.044
$4,000	.195	.045	.044
$5,000	.189	.052	.050
$6,000	.189	.060	.063
$7,000	.181	.062	.067
$8,000	.170	.063	.068
$9,000	.176	.070	.079
$10,000	.165	.077	.065
$11,000	.179	.081	.069
$12,000	.183	.075	.062
$13,000	.189	.084	.056
$14,000	.204	.087	.053
$15,000	.221	.086	.044
$16,000	.208	.109	.048
$17,000	.201	.093	.029
$18,000	.230	.104	.037
$19,000	.216	.075	.007
Mean	.182	.056	.055

[a]Lower bound.

[b]Includes nonproduction workers.

Table C

Joint Relative Frequency Distribution of Income:
Nonaged Families and Individuals

Income Class[a]	Union	Nonunion	Total
$0	.007	.139	.146
$4,000	.006	.044	.050
$5,000	.010	.047	.058
$6,000	.011	.050	.061
$7,000	.017	.050	.067
$8,000	.017	.052	.069
$9,000	.019	.049	.068
$10,000	.019	.050	.069
$11,000	.017	.040	.056
$12,000	.016	.041	.057
$13,000	.012	.031	.044
$14,000	.010	.028	.038
$15,000	.008	.027	.036
$16,000	.006	.021	.028
$17,000	.005	.020	.025
$18,000	.005	.016	.021
$19,000	.004	.012	.016
$20,000	.003	.012	.015
$21,000	.002	.008	.009
$22,000	.002	.007	.009
$23,000	.001	.007	.008
$24,000	.001	.004	.006
$25,000	.002	.018	.021
$30,000	.001	.022	.023
Total	.203	.797	1.000
Mean income	$11,367	$10,696	$10,833

[a]Lower bound.

Table D

The Share of Private-Sector Production Earnings in Income

Family Income[a]	Union Families: Production Earnings	Union Families: Union Production Earnings	Nonunion Families: Production Earnings
$0	.847	.805	.259
$4,000	.894	.847	.354
$5,000	.893	.849	.371
$6,000	.895	.857	.328
$7,000	.932	.883	.325
$8,000	.909	.863	.314
$9,000	.885	.828	.293
$10,000	.886	.821	.249
$11,000	.868	.790	.244
$12,000	.834	.756	.226
$13,000	.808	.726	.216
$14,000	.811	.718	.186
$15,000	.794	.721	.159
$16,000	.720	.636	.155
$17,000	.703	.615	.158
$18,000	.695	.602	.112
$19,000	.659	.591	.116
$20,000	.688	.593	.198
$21,000	.688	.572	.113
$22,000	.635	.506	.072
$23,000	.630	.536	.074
$24,000	.579	.492	.087
$25,000	.578	.507	.048
$30,000	.432	.335	.037

[a]Lower Bound.

Table E

Relative Income Effects

Income Class[a]	Union	Nonunion	All
$0	.186	.007	.016
$4,000	.177	.021	.040
$5,000	.174	.017	.044
$6,000	.177	.017	.046
$7,000	.165	.016	.042
$8,000	.165	.018	.054
$9,000	.165	.017	.058
$10,000	.157	.015	.054
$11,000	.158	.014	.058
$12,000	.159	.014	.055
$13,000	.145	.015	.050
$14,000	.153	.011	.048
$15,000	.140	.010	.038
$16,000	.134	.009	.035
$17,000	.128	.010	.034
$18,000	.133	.008	.038
$19,000	.126	.008	.038
$20,000	.123	.005	.029
$21,000	.129	.008	.036
$22,000	.136	.004	.033
$23,000	.116	.004	.018
$24,000	.111	.004	.021
$25,000	.130	.003	.015
$30,000	.090	.002	.006
Average	.155	.012	.041

[a]Lower bound.

Table F

The Share of Private-Sector Production Earnings in Income

POVRATIO[a]	Union Families: Production Earnings	Union Families: Union Production Earnings	Nonunion Families: Nonunion Production Earnings
0	.768	.838	.244
1.0	.816	.879	.359
1.5	.846	.896	.377
2.0	.819	.882	.310
3.0	.779	.844	.215
4.0	.675	.749	.122

[a]Lower bound.

NOTES

[1]See Rees (1977, p. 92) and Friedman (1954, p. 208). Sherwin Rosen (1970) presents a rigorous demonstration of these propositions in the framework of a model in which unions maximize their members' collective earnings.

[2]Two examples are Rees (1977, pp. 92-93) and Johnson (1975, p. 26). Empirical support for this view may be found in Lewis (1963).

[3]For a concise discussion of this behavior see Reynolds (1956).

[4]For instance, coordinated craft bargaining in the railroads led to gradual wage compression extending over several decades (see Burgoon, 1970).

[5]For instance, see "Wage Structure in Steel Mills Narrows during the 70s," Monthly Labor Review 103(12) (Dec. 1980). Analyzing time-series data on contractual wage changes, Mitchell (1980, pp. 149-151) finds evidence of wage compression in the union sector; the greatest compression, however, has occurred among workers with COLAs in their contracts.

[6]Numerous case studies and journalistic accounts suggest such a strategy (e.g., Conant, 1959; Sease, 1981). An econometric study by Gay (1975) also finds evidence of such behavior.

[7]The above discussion thus leads us to expect that coefficients of the dummies will be positive and increasing in value, although the rate of increase may vary from level to level.

[8]A unique feature of the March 1971 CPS file is that it allows us to match information about union members with information on family

incomes. The information on family income is only available in the March survey. Since 1973, information on union affiliation was only obtained in the May survey. The March 1973 CPS is unique in that information on both family incomes and union affiliation is available.

[9] Many studies of union wage impacts include only union membership dummy variables in their earnings equations and omit any measures of industry collective bargaining coverage (e.g., 6 of the 9 studies in Table 1). The evidence in Table 2 suggests that this may be a serious omission, since the wage advantage conferred by union membership within a coverage class is usually fairly small when compared to the effect of coverage.

[10] Podgursky (1980, pp. 48-50). The purpose of the second test was to assure that a large significant difference in intercepts was not dominating small insignificant differences in slopes.

[11] Several recent studies of union wage effects find that ordinary least squares (OLS) estimates are biased by worker "self-selection" in that workers with unmeasured but remunerative attributes may systematically join or avoid unions (e.g., Ashenfelter and Johnson, 1970; Schmidt and Strauss, 1976; Lee, 1978). Although a priori there is no way to be sure of the direction of bias, these studies usually find that OLS estimates overestimate the union wage effect.

In order to assess the effect of this "self-selection" bias on these estimates, a random subsample of 950 observations was drawn from the CPS file. Consistent econometric corrections procedures were employed on 18 different, but a priori plausible, specifications of the wage equations. While the OLS estimates were quite robust over all of the model specifications, their "corrected" counterparts were found to be highly sensitive to minor specification changes. In some cases unions reduced wages by over 60%, while in others they raised wages by over 40% (Podgursky, 1980, pp. 53-64). Others find similar instability (Mitchell, 1980, pp. 110-111).

In the face of such variation, the choice of a single set of "selection-corrected" estimates is clearly problematic. For this reason OLS estimates are used in the study.

[12]The issue of the incidence of collective bargaining wage gains has been examined from neoclassical (Johnson and Mieszkowski, 1970) and Cambridge perspectives (Kalecki, 1971, pp. 156-164). In these simple general equilibrium models, the pattern of incidence, not surprisingly, depends on a variety of parameters about which we know little.

Several simple experiments were performed in this study under the assumption that the full burden of union wage gains fell upon: (1) production workers; or (2) nonproduction workers. In case (1) the net union effect was slightly disequalizing and case (2) slightly equalizing. In neither case did measured earnings inequality change by more than 5% (Podgursky, 1980, pp. 98-100).

[13]"Voluntarism," as articulated by AFL president Sam Gompers, was the labor movement equivalent of "laissez-faire" (see Rogin, 1971). With the rise of the CIO in the 1930s, the labor movement overcame its aversion to government intervention in the labor market.

REFERENCES

Ashenfelter, Orley. 1972. Racial discrimination and trade unionism. Journal of Political Economy, 80, 435-464.

_____. 1976. Union relative wage effects: New evidence and a survey of their implications for wage inflation. Princeton University, Industrial Relations Section.

_____ and Johnson, George. 1972. Unionism, relative wages and labor quality in U.S. manufacturing industries. International Economic Review, 13, 488-508.

Atkinson, Anthony B. 1970. On the measurement of economic inequality. Journal of Economic Theory, 3, 244-263.

Bishop, John H. 1976. Queuing for union jobs and the social return to schooling. Institute for Research on Poverty, Discussion Paper 360-76, Madison, Wisconsin.

Bloch, Farrell E., and Kushkin, Mark S. 1978. Wage determination in union and nonunion sectors. Industrial and Labor Relations Review, 31, 183-192.

Boskin, Michael. 1972. Unions and relative real wages. American Economic Review, 62, 466-472.

Budd, Edward C. 1967. Introduction. In Edward C. Budd (ed.), Inequality and Poverty. New York: W.W. Norton and Company.

Burgoon, Beatrice. 1970. Effects of the structure of collective bargaining in selected industries: The railroad industry. Labor Law Journal (Aug.), 491-498.

Conant, Eaton H. 1959. Defenses of nonunion employers: A study from company sources. Labor Law Journal, 10, 100-109.

Douty, Harry M. 1953. Union impact on wage structures. In Industrial Relations Research Association, Proceedings of the Sixth Annual Meeting. Madison, Wisconsin.

Freeman, Richard B. 1980. Unionism and the dispersion of wages. Industrial and Labor Relations Review, 34, 3-23.

______ and Medoff, James L. 1979. Industrial and Labor Relations Review, 32, 143-174.

Friedman, Milton. 1954. Some comments on the significance of labor unions for economic policy. In David McCord Wright (ed.), The impact of the union. New York: Kelly and Millman.

Gay, Robert S. 1975. The impact of unions on relative real wages: New evidence on effects within industries and threat effects. Ph.D. Dissertation, Department of Economics, University of Wisconsin-Madison.

Hyclak, Thomas. 1979. The effect of unions on earnings inequality in local labor markets. Industrial and Labor Relations Review, 33, 77-84.

Johnson, George E. 1975. Economic analysis of trade unionism. American Economic Review, 65, 23-28.

Johnson, Harry G., and Mieszkowski, Peter. 1970. The effects of unionization on the distribution of income. Quarterly Journal of Economics, 84, 539-561.

Kalecki, Michal. 1971. Selected essays on the dynamics of the capitalist economy. London: Cambridge University Press.

Levinson, Harold. 1954. Collective bargaining and income distribution. American Economic Review, 44, 308-316.

Lewis, H. Gregg. 1963. Unionism and relative wages in the United States. Chicago: University of Chicago Press.

Mitchell, Daniel J. B. 1980. Unions, wages and inflation. Washington, D.C.: Brookings, 1980.

Ozanne, Robert. 1959. The impact of unions on wage levels and income distribution. Quarterly Journal of Economics, 73, 177-196.

Podgursky, Michael. 1980. Trade unions and income inequality. Ph.D. dissertation, Department of Economics, University of Wisconsin-Madison.

Rees, Albert. 1977. The economics of trade unions. 2nd Ed. Chicago: University of Chicago Press.

______. 1979. The economics of work and pay. 2nd Ed. New York: Harper and Row.

Reynolds, Floyd G. 1964. The impact of collective bargaining on the wage structure in the U.S. In John Dunlop (ed.), The theory of wage determination. London: Macmillan and Co.

______ and Taft, Cynthia H. 1956. The evolution of the wage structure. New Haven: Yale University Press.

Rogin, Michael. 1971. Voluntarism: The Political Functions of an Anti-Political Doctrine. In David Brody (ed.), The American labor movement. New York: Harper and Row.

Rosen, Sherwin. 1970. Unionism and the occupational wage structure in the U.S. International Economic Review, 11, 269-286.

Ryscavage, Paul M. 1974. Measuring union-nonunion earnings differences. Monthly Labor Review, 97, December, 3-10.

Sease, Douglas R. 1981. Little giants: Mini-mill steelmakers no longer very small. Wall Street Journal (Jan. 12), p. 1.

Stafford, Frank P. 1968. Concentration and labor earnings: Comment. American Economic Review, 58, 174-181.

U.S. Bureau of the Census. 1979. Current population reports. Series P-60. No. 119, Characteristics of the population below the poverty level: 1977. Washington, D.C.: U.S. Government Printing Office.

Weiss, Leonard. 1966. Concentration and labor earnings. American Economic Review, 56, 96-117.

Williamson, Jeffrey G. 1976. The sources of American income inequality, 1896-1948. Review of Economics and Statistics, 58, 387-397.