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ECONOMIC EFFECTS OF THE FIREFIGHTERS' UNION

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ABSTRACT

Economic Effects of the Firefighters' Union

This is a study of the effects of unionism in the public sector occupation of firefighting. A large and detailed set of data permits the examination of submarkets of this occupation. A before/after methodology is introduced to obtain more precise estimates of union wage differentials. The study's findings are: (1) that there is a greater union effect on fringes than on salaries which indicates a significant alteration in the composition of the compensation package; (2) that the estimates from the before/after methodology confirm the cross-section results which show modest union wage differentials; and, most significantly, (3) that the union effect varies along different dimensions -- most notably the length of the contractual arrangement between municipality and union.

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The growth of public sector unionism in the last two decades raises important questions about the economic effects of unions of governmental workers. Do these unions have large or small impacts on wages or costs? Most recent research, based on cross-sectional comparisons of wages and union status across cities or selected decision units (see Lewin for a useful summary), have found only small effects. In part because of data availability, little attention has been given to the impact of public unions on fringe benefits and the structure of wages, and little to differences in the impact of unions under different market conditions. Although public sector unionism is a recent development, no study has estimated union effects on a before-after basis.

This study seeks to remedy these gaps in our knowledge. It examines the economic effects of the International Association of Fire Fighters (IAFF) using a large pooled cross-section sample of cities. It compares cities before and after unionism as well as on a cross-section basis; it estimates the impact of unionism on fringes and the structure of wages as well as total compensation, and on cities with different forms of municipal governments. The large sample permits estimation of separate equations for union and nonunion cities which cast light on the way unionism alters the wage-setting process.

Previous work on the IAFF has dealt with relatively small samples (Ashenfelter, 130 cities; Ehrenbrugh, 256 cities) and been limited to cross-section comparisons of wage effects. Ashenfelter found a modest union impact on wages (the average cross-section result for 1961 to 1966 was .036) but controlled only for cost of living differences across cities. Ehrenburgh made the important innovation of using the presence of an IAFF contract as the union variable and found a larger effect (.098). Neither study corrected for 'omitted' city factors that might be associated with both unionism and the level of wages.

This study finds moderate union compensation effects using cross-section data on over 1,000 cities and comparisons of pay before and after unionism in 300 to 500 cities, and finds that most of the effect occurs via fringes rather than straight-time pay. The similarity between before-after and cross-section comparisons suggests no serious selectivity bias in the standard cross-section union equations. The study also finds quite different union effects depending on the length of time organized and on the type of cities, suggesting that market and organizational factors influence the economic impact of unionism.

The paper is divided into four sections. Section I describes the institutional setting in which the IAFF and cities bargain over compensation. Section II describes the data used in the analysis. Section III sets out the econometric model. Section IV presents the basic empirical results and considers differential effects of unionism among types of cities. The paper concludes with a brief summary of findings.

I. The Firefighter's Market

Firefighting is the most extensively organized public employee occupation. The IAFF has been organizing fire fighters since 1918, so many large fire departments were IAFF locals prior to the surge of public employee unionism in the sixties. In 1960, approximately 74 percent of the nation's 137,884 firefighters were organized. Through the sixties, the IAFF's ranks grew (from 1962 to 1964, 85 new locals and 2,044 members; from 1964 to 1966, 122 new locals and 7,000 members); however, this growth did little more than keep pace with the growth of the occupation itself. By 1972, there were 194,785 firefighters in the United States, and approximately 77 percent organized.

As municipal employees, firefighters are paid out of a publicly financed budget, not out of a firm's revenue like private employees. Revenue from property taxes is usually the largest part of a municipal budget, supplemented by income and sales taxes. The property tax-based revenue is often divided between more than one governing unit or district, all depending on the same base. Because municipal government decisions depend on a politically formulated budget, the collective bargaining process in firefighting is political in nature. Distinct from actual negotiations, IAFF lobbying on issues of budget formulation and budget allocations is a key process affecting firefighters' pay. Often the IAFF finds itself competing with other municipal unions for a larger share of the budget. (Although this inter-union competition is similar to craft union competition in the construction industry, the occupational peculiarities of fire fighting limit the usefulness of such an analogy.)

The second "tier" of the process, the actual negotiations, also has distinctive political aspects. Government authority is generally not as centralized as that of a private institution. A bargaining representative may not even have the authority to allocate funds for a settlement he may have reached. A legislature may repudiate the settlement, or certain state funds may not be forthcoming to ensure a settlement at the local level. Despite the problems of repudiation, and the potentially lengthy delays involved in litigations, rarely will a contract dispute result in a strike. Depending on state laws, public employees and union officials may face stiff fines, jailing or even dismissal for participating in a public-employee strike. Although the IAFF eliminated a "no-strike" clause from its constitution at its 1968 convention, harsh consequences do deter firefighter strikes. In 1972, there were only 11 work stoppages by firefighters in the United States. The average work stoppage involved a department of 52 employees and lasted for only four days. To remedy the costly problems of delays due to repudiation and litigation, 17 states have adopted arbitration systems for public sector bargaining impasses. Five of these states have adopted a "last best offer" arbitration system in which the arbitrator chooses either the last position of union or of management on an issue-by-issue basis, or on a package basis. However, the results of the arbitration process are often longer delays, and occasionally more court action.

IAFF contract negotiations in the town of Ipswich, Massachusetts illustrate the problems involved in this second tier of public bargaining. In Ipswich contract negotiation has involved an initial settlement, repudiation of that settlement eight months later, failure of mediation and fact finding by appropriate state agencies, a last best offer arbitration award in favor of the IAFF local, court action which vacated the arbitration panel's award, and an appeal by the local of the court decision. Finally, the town and the firefighters agreed to a retroactive agreement in April 1978, but through the first months of 1978 the Ipswich firefighters continued to work under their 1973 contract.

Overall, contrary to the claim of Wellington and Winter, industrial relations in this public sector do not appear to create great union power because of the essential nature of the services provided. The threat of repudiation and court delays makes it more difficult for the heads of the locals to deliver a wage agreement to their ranks. The penalties accompanying strikes severely limit the economic power of the union. Moreover, seniority compensation (called longevity pay in the protective services) and pension benefits are often non-portable from department to department, depending on state rulings. This non portability of benefits is sometimes coupled with residency requirements. In such a case, a firefighter cannot change departments without first changing his residence and losing his longevity salary increments and his position in his employer's pension system. This immobility within the firefighter occupation works against the bargaining power of the IAFF, because firefighters are not free to change jobs to look for better conditions. It would appear unreasonable to expect enormous union wage effects.

II. Data

This study has obtained data from a variety of sources on wages, fringes, unionism, and diverse control variables to estimate the economics of the IAFF. Wage and fringe data are obtained from statistics on wage and salary expenditures (S), number of full-time employees (E), and duty hours per week (H), and city contribution (C) to employees' retirement benefits and to insurance programs, in the Municipal Yearbook. The average hourly wage is measured by $W = (S/E)/52H$ on the assumption that workers are employed year round. Fringes paid per hour are defined as $F = (C/E)/52H$ while total compensation (TC) is the sum $F + W$. The data on wages are available for two years: 1966, when relatively few collective contracts were signed for firefighters, and for 1976. Fringe data relate to 1976 only and thus cannot be analyzed in the before-after framework.[1]

In addition to the average hourly wage or salary and fringes, the analysis uses total salary (52HW) and entrance and maximum salaries (defined as "the annual base salary of a firefighter during his first twelve months on the force,... and the maximum annual base salary paid full-time firefighters not holding any promotional rank"), as dependent variables likely to be affected by unionism. If the IAFF influences hourly wages by changing hours worked rather than yearly salary, the coefficient on the hourly and yearly wage equations will differ. If the IAFF changes the wage structure or accelerates promotions, it is likely to have different effects on entrance, maximum, and average salaries.

Unionism is measured by the presence of an IAFF contract and in some calculations by the presence of an IAFF local. As Ehrenberg argues, the contract variable provides a better measure of the potential effectiveness of the union in altering wages and work conditions. Because of the lack of data on unionization for some cities in 1966, the study employs three samples: (1) a sample of 1,015 cities for the 1976 cross-section when figures exist for all of the relevant cities; (2) a sample of 597 cities for the 1966-1976 analysis based on the smaller number of cities in the Municipal Yearbook 1966, with cities having "not reported or not applicable" union figures categorized as nonunion (a reasonable surmise as those cities were not included in the 1964 IAFF Convention Reports as sending delegates to the union convention); and (3) a smaller sample of 307 cities in which the "not reported or not applicable" were deleted from consideration.

The other "control" variables fall into two categories: those which differ over time[2] and those which do not change over time for each city (or in which the change is assumed to be insignificant in the determination of wages).

Those variables which change over time include: population; opportunity wage; alarms per 1,000 people; per capita income; median value of single family housing; and per capita general revenue from city's own sources. Population is expected to be positively related to the demand for fire services, while opportunity wage[3] should be negatively related to the supply of firefighters. The fire insurance rating could signal a compensating differential for work in more hazardous communities; number of alarms, as a proxy for number of runs by a department, is expected to be positively related to wage. Per capita income, median housing value, and per capita revenue should yield positive coefficients in wage equations as well.

Those variables in which the change is assumed to be insignificant in the determination of wages are: percent of population that is non-white; four region controls; government type (council-manager, commission plan, mayor-council, town meeting); and land area. With population controlled land area may be positively related to wages (fire department responsible for more territory) or negatively related to wages (fire department responsible for a more densely populated area, in which firefighting may be more hazardous).[4]

III. Econometric Models

Two types of econometric models will be used to estimate the effect of the IAFF on pay. As in previous studies of union wage effects, a cross-section model of the following form will be estimated first:

$$(1) \qquad \ln W_1 = a_1 + \alpha_1 U_1 + \beta_1 \vec{X}_1 + \varepsilon_1$$

where W_1 = wage or compensation

U_1 = 0-1 variable for whether or not city has a collective contract for period 1

$\vec{X}_1$ = list of controls for period 1

ε_1 = residual, assumed $N(0, \sigma^2)$

To the extent that the X's control for factors which are correlated with $\ln W_1$ and unionism, so that $E(\varepsilon_1 U) = 0$, equation (1) will yield an unbiased estimate of the union effect.

It is reasonable to expect that the union effect will differ along various dimensions. Information on organization in a prior (base year) period permits the model (1) to be extended to allow the length of organization to influence the union effect. Let U_{01} = 0-1 dummy variable for having a contract in the base year and in the current period; U_0 = 0-1 dummy variable for having a contract in the base year only; and U_1 = 0-1 dummy variable for a contract in the current period. Then we can expand (1) as follows:

$$(2) \qquad \ln W_1 = a_1' + \alpha_1 U_1 + \alpha_2 (U_{01}) + \alpha_3 (U_0) + \beta_1' \vec{X}_1 + \varepsilon_1'$$

If being organized over two periods has a greater impact on wages than organization over a shorter span, the coefficient on $U_{01}(\alpha_2)$ will exceed that on $U_1(\alpha_1)$. If a city that signs a contract in an early period but not in a later period maintains union wage scales, the coefficient on $U_0(\alpha_3)$ will be positive. If neither of these factors operate, (2) will collapse in the model (1).

If the $\vec{X}$'s do not completely control for wage-determining factors that are correlated with unionism, (1) and (2) will not yield the desired impact parameters. Assume, for example, that there is an omitted city effect (D) that is positively correlated with unionism and with wages: in this case the OLS estimate of (1) would yield an upwardly biased estimate of the true α_1. Alternatively, if the omitted effect were positively correlated with U_1 and negatively with wages, the OLS estimate would be downward biased.

To handle this problem, we make further use of data on an earlier cross-section, in which the omitted factor is also expected to operate. Formally, assume that in (1) $\varepsilon_1 = D + \ell_1$ where $E(DU_1) \neq 0$ and $E(\ell_1 U_1) = 0$. Assume also that D also affects wages in the other period:

$$\text{(3)} \qquad \ln W_0 = a_0 + \alpha_0 U_0 + \beta_0 \vec{X}_0 + \frac{1}{\lambda} D + \ell_0$$

where $\frac{1}{\lambda}$ is a scaling factor that permits city effects to vary over time and where $E(\ell_0 \ell_1) = E(\ell_0 \vec{X}) = E(\ell_0 U) = 0$. In this base year equation, there is no specification of detailed union variables, as the U_0 period is chosen early enough so that interaction variables with even earlier periods have little meaning. Substituting for D in (1) we obtain:

$$\text{(4)} \qquad \ln W_1 = \lambda \ln W_0 + (a_1 - \lambda a_0) + \beta_1 \vec{X}_1 - \lambda \beta_0 \vec{X}_0 + \alpha_1 U_1 - \lambda \alpha_0 U_0 + \ell_1 - \lambda \ell_0 .$$

If $\lambda=1$ and $U_0 = 0$, (4) becomes a straightforward before-after comparison:

$$\text{(5)} \qquad \ln W_1 - \ln W_0 = \hat{a} + \beta_1 \vec{X}_1 - \beta_0 \vec{X}_0 + \alpha_1 U_1 - \alpha_0 U_0 + \ell_1 - \ell_0$$

where $E((\ell_1 - \ell_0) U_1) = 0$.

When $\lambda \neq 1$, however, equation (4) must be estimated conditional on λ. An estimate of λ can be obtained by regressing $\ln W_1$ on $\ln W_0$ and the other right-hand side variables of (4). However, because $E(W_0 \ell_0) \neq 0$, the estimate of λ will be biased downward, causing an upward bias in the union coefficient if, conditional on all other variables, W_1 is positively correlated with W_0. To estimate the magnitude of the bias, let $b_{W_0U_1 \cdot \vec{X}}$ be the regression coefficient linking the wage in the base period to unionism in period 1, conditional on all other variables; let $r_{W_0U_1 \cdot \vec{X}}$ be the corresponding partial correlation coefficient and P ($0 \leq P \leq 1$) be the ratio of the variance of ℓ_0 to the variance of $\ln W_0$. Then the bias on the union coefficient α due to the failure to obtain the appropriate value of λ is determined by (see Griliches & Ringstad, p. 197):

$$(6) \qquad \underset{n \to \infty}{\text{plim}}\ \hat{\alpha} = \frac{b_{W_0U_1 \cdot \vec{X}}}{1-r^2_{W_0U_1 \cdot \vec{X}}} P\alpha + \alpha$$

while the bias in estimating λ is

$$\underset{n \to \infty}{\text{plim}}\ \hat{\lambda} = 1-P/(1-r^2_{W_0U_1 \cdot \vec{X}})$$

Empirically, the values of the relevant regression and correlation coefficients suggest that the bias on the union term is slight.

More complex equations, which include union interaction terms as in (6) can also be developed, with comparable results. The key point is that by addition of the base year wages to the equation, one obtains at least some control on the charges that results are due to unobserved city factors.

The empirical work focuses on two dependent variables, wages paid firemen and fringe benefits. For several reasons the union effect is likely to be greater on fringes than on straight-time pay. First, in setting their bargaining goals unions are likely to give greater weight to the preferences of the more senior employees who favor fringes and less to young marginal employees than would occur in a competitive market (see Freeman for detailed discussion). Second, because of the timing of elections, and the short-term horizon of politicians, local governments are likely to be more willing to pay fringes whose costs will accrue more greatly in the future than in the present.

When straight-time pay is the dependent variable, (1) or (4) or (5) can be estimated directly. When fringes are the dependent variable, it is also necessary to control for the overall level of pay, due to the likely positive income elasticity of fringes. Simply adding total compensation to the list of independent variables would not be correct, for fringes are included in total compensation. To deal with this problem, let us write the model with variables in linear form:

$$(7) \qquad F = \beta_0 + \beta_1\vec{X} + \beta_2(TC) + \gamma U + \varepsilon$$

$$= \beta_0 + \beta_1\vec{X} + \beta_2(F{+}S) + \gamma U + \varepsilon$$

where F = dollars of fringes

TC = dollars of total compensation

S = dollars of straight-time pay

Solving for F by separating F out of TC,

$$(8) \qquad F = \beta_0/1-\beta_2 + \beta_1\vec{X}/1-\beta_2 + (\beta_2/1-\beta_2)S + (\gamma/1-\beta_2)U + \varepsilon/1-\beta_2$$

The true union effect, γ, is isolated by first solving for β_2. Let $\hat{\beta}_2$ represent the coefficient on S obtained in (8). The product of the coefficient on unionism in (8) and $1-\beta_2$ (which equals $1/1+\hat{\beta}_2$) then yields the desired parameter γ of equation (7).

Besides the union effect on straight-time pay and fringes, unionism may also affect the structure of wages in the city, as has been found to be the case of the teachers' union (Gustman and Segal). We will examine the effects of the IAFF on maximum pay, entrance salary, and hours worked.

Finally, to evaluate the possibility that unionism influences the entire process of wage-setting and/or has different effects on different types of governmental bodies, separate union and nonunion wage equations will be estimated. With separate equations, union effects will show up in different weights placed on different factors. Average effects can be calculated by using the regression weights to standardize the wages. For example, if $\hat{\alpha}_i$ is the coefficient on X_i in the nonunion equation, and $\bar{X}_u$ is the mean of the variable for unionists, the overall union wage effect would be the difference between what unionists got (W_u) and what they would have obtained if they were nonunion:

$$W_u - \Sigma\hat{\alpha}_n \bar{X}_u , \tag{9}$$

Conversely, using the union equation to provide weights, the union effect could be estimated as:

$$W_n - \Sigma\hat{\alpha}_u \bar{X}_n \tag{10}$$

where W_n = wage of nonunionists.

IV. Basic Empirical Results

Table 1 presents estimates of the effect of an IAFF local and, more importantly, a signed contract on several measures of the wages and work conditions of firemen for 1976 and 1966. The table presents calculations for several samples, defined by the selection criteria described earlier. In column 1 unionism is measured by the presence of an IAFF local; in columns 2-6 by a signed contract. Columns 1 and 2 show, consistent with the Ehrenberg argument, that the presence of a contract is the key to a trade union effect, with the measures of pay in the largest sample unaffected by the IAFF local *per se* but positively influenced by the presence of the contract. The remaining columns focus on the contract variable. For 1976, the table reveals a modest, but significant, union effect on total compensation, a smaller effect on hourly and annual salary, little impact on hours worked, and a small effect on the maximum and entrance scales. In 1966, by contrast, the union effects are much less

Table 1

Cross-Section Estimates of IAFF Earnings Differentials, 1966 and 1976*

Year	1976				1966	
	1015		597	307	597	307
Dependent Variable	local	contract				
1. Total Compensation	.010 (.016)	.042 (.017)	.041 (.017)	.056 (.021)		
2. Wage (salary/hour)	-.007 (.018)	.032 (.019)	.031 (.017)	.043 (.019)	.012 (.025)	-.002 (.023)
3. Salary	-.007 (.015)	.031 (.016)	.023 (.015)	.034 (.017)	-.016 (.018)	-.014 (.015)
4. Hours	.006 (.007)	-.001 (.008)	-.008 (.008)	-.009 (.015)	-.028 (.015)	.016 (.014)
5. Maximum Salary	.008 (.008)	.018 (.009)	.026 (.010)	.015 (.014)	-.020 (.013)	-.010 (.013)
6. Entrance Salary	.006 (.008)	.023 (.009)	.029 (.011)	.029 (.014)	-.003 (.013)	.002 (.014)

1976 controls: land area, percent nonwhite, per capita city revenue, population, alternate salary, fire rating, alarms per 1000 people, per capita income, median value of housing, government type, region.

1966 controls: land area, percent nonwhite, per capita city revenue, population, alternate salary, fire rating, alarms per 1000 people, median value of housing, government type, region.

*Standard errors are in parentheses.

and do not differ significantly from zero. In both years the results differ only slightly by sample size.

The markedly greater effect of unionism on total compensation than on salaries in the 1976 cross-section imply a sizeable effect on fringes, since fringes make up 14.6 percent of total compensation. Table 2 shows the importance of introducing a more detailed union break, as in equation (2). Using three contract variables (contract in both periods, $U_{76,\ 66}$; contract in 1976 only, U_{76}; and contract in 1966 only, U_{66}[5]) reveals a clear and striking pattern. With the exceptions of the hours variable (in which there are no significant differences among the three coefficients) and of the maximum salary variable with n=597 (in which the contract-in-1976-only coefficient is insignificantly greater than the contract in both periods coefficient), it is always the case that the $U_{66,\ 76}$ coefficient is largest and the U_{66} coefficient smallest. Furthermore, the $U_{66,\ 76}$ coefficients in the non-salary scale equations are nearly twice the coefficients on simple contract variables in the corresponding "equation 1" - type cross-sections, indicating that the union effect varies significantly with length of time organized.

Fringes

Table 3 examines the union effect on fringes by regressing \$1/year in fringes on unionism and the control variables used in Table 1 and by estimating logistic probability equations that relate the probability of having city-funded fringes to unionism. The results are clear. Total fringe spending is raised by \$170.38 or 9.4 percent of fringes by presence of an IAFF contract (column 1, line 5).

More specifically, the effect of unions on fringe spending can be of two types: unions can raise the likelihood that a city offers a given fringe or it can raise spending

Table 2: 1976 Cross-Section with Detailed Contract Variables

	n = 307			n = 597		
	Contract in 1976 and in 1966 ($U_{76,66}$)	Contract in 1976 only U_{76})	Contract in 1966 only (U_{66})	Contract in 1976 and in 1966 ($U_{76,66}$)	Contract in 1976 only (U_{76})	Contract in 1966 only (U_{66})
1. Total Compensation	.082 (.034)	.026 (.023)	-.061 (.031)	.073 (.033)	.022 (.018)	-.071 (.030)
2. Wage (sal/hr)	.079 (.031)	.036 (.021)	.009 (.028)	.046 (.033)	.026 (.018)	-.009 (.030)
3. Salary	.052 (.028)	.021 (.019)	-.020 (.025)	.029 (.029)	.016 (.016)	-.036 (.026)
4. Hours	-.027 (.015)	-.014 (.010)	-.029 (.013)	-.018 (.016)	-.011 (.009)	-.027 (.014)
5. Maximum Salary	.027 (.022)	.022 (.016)	-.005 (.020)	.017 (.021)	.026 (.011)	-.015 (.018)
6. Entrance Salary	.035 (.024)	.025 (.016)	-.007 (.021)	.031 (.021)	.028 (.011)	-.005 (.019)

Other controls: Same as 1976 controls, Table 1.

Table 3: IAFF Effect on Fringe Benefits,

by Type of Fringe, for 1976

(n = 1098)

	All Fringes	Retirement	Insurance
1. Mean (in $/year)	1812.50 (1471.00)	1415.50 (1330.20)	397.0 (350.0)
2. Union Effect (in $)	152.25 (94.25)	80.68 (87.76)	72.65 (23.42)
3. Union Effect (in $) with salary controlled	148.63 (88.81)	76.44 (82.10)	72.25 (23.42)
4. Union Effect, salary controlled as % of mean (line 3/line 1)	.082	.054	.182
5. Union-Fringe elasticity (corrected according to (8), Sec III)	.094	.061	.184
6. Proportion of cities with a city-funded program	.934	.879	.872
7. Logit coefficient for likelihood of program	.917 (.309)	.461 (.222)	.703 (.225)
8. Union Effect on likelihood of program	.057	.049	.079

*Standard deviations/standard errors in parentheses.

Note: Other controls are according to "1976 controls" from Table 1.

on that fringe. To examine the former possibility a logit curve [$P = 1/1+\exp - \Sigma_i X_i$] was estimated linking the probability of having specified fringes to a contract and other variables. As can be seen in lines 7 and 8, according to the logistic equations, unionism raises the probability of having those benefits of .05 and .08 points respectively at the mean levels.

The union effect on the spending on fringes is examined in lines (1) through (5). Line (3) shows the union coefficient taken from the equation (8) model of Section III. Then correcting these coefficients according to the analysis of that section yields the line (5) coefficients. The union effect on both fringes taken together is 9.4 percent; on retirement fringes (which comprise 78.1 percent of this fringe package), it is 6.1 percent; and on the insurance benefits (with a mean of $397 for all cities), 18.4 percent. As the results in Table 1 imply, the union effect on fringes is significantly greater than on salaries and wages.

IAFF representatives offer interesting explanations for the greater union impact on fringes. A Los Angeles representative cites the importance of the IAFF's central research department in "informing the locals' representatives of the successes of various experiments in the arrangement of fringe packages." He also feels that city negotiators are sympathetic to the argument that the government has a "social responsibility" to keep fringe benefits up to levels offered to employees in the private sector. Finally, he says that the costs of fringes "do not impact immediately." A bargaining representative and trustee for the Yonkers fire fighters also feels that city negotiators do not realize the full impact of fringe costs. This is especially true if an administration is going out of office. He has found that bargaining with a lame-duck administration, when there is going to be a change in party control, allows the union local to make significant gains in fringes. He says the administration of one party will sometimes make it more difficult on the rival party's administration by putting a greater strain on following years' budgets, while winning support of a special interest group -- the firefighters.

Before/after analysis

To what extent might the preceding results be in error because of the possible correlation between unionism (other independent variables) and omitted city factors? Does before/after contract analysis yield similar or different estimated union effects?

As a first step toward getting a handle on this question, it is useful to find out which cities are being unionized. From 1966 to 1976, approximately 19 percent of the sample of 597 became organized. Logistic curve estimates of the impact of wages on the probability of organization yielded only a slight tendency for organization to be affected by initial wages (a logit coefficient of .55 with asymptotic standard error of .80). Since the correlation between unionization and salary or wages is not clear cut, the implication is that the omitted city factor is not an important determinant of wages and salaries that is correlated with unionism.

Part I of Table 4 presents explicit estimates of the contract effect on average yearly and hourly salaries and on the maximum and entrance salaries given a λ of 1. It regresses the change in salaries on the presence or absence of contracts in 1976 and 1966 broken into the three discrete categories used earlier: contracts in both years ($U_{76,66}$); no contract in the initial year, contract in 1976 (U_{76}); and contract in 1966, none in 1976 (U_{66}); with a deleted group of no contracts in both periods.

If unionism had the same effect on pay in both years there would be <u>no</u> effect on the $U_{66,\ 76}$ variable while those on U_{76} and U_{66} would have equal opposite signs. If unionism had larger effects in 1976 than 1966, as indicated in the cross-section regressions, the $U_{66,76}$ dummy would obtain a positive coefficient which would be smaller than that on U_{76} while the U_{66} group would have negative coefficients of equal magnitude. As is evident in the tables, neither of these possibilities turns out to be the case. While the U_{66}

Table 4: Estimates of the Impact of Unionism

Obtained from "Before/After" Methodology*

	Dependent Variable:	Hourly Wage	Average Salary	Entrance Salary Scale	Maximum Salary Scale
I	n = 307				
	$U_{76,66}$	.037 (.034)	.053 (.027)	.035 (.022)	.032 (.021)
	U_{76}	.013 (.024)	.017 (.018)	.026 (.015)	.017 (.014)
	U_{66}	.021 (.030)	.008 (.023)	.007 (.020)	.005 (.018)
	λ	1	1	1	1
	n = 597				
	$U_{76,66}$	-.005 (.037)	.033 (.031)	.039 (.020)	.036 (.020)
	U_{76}	-.020 (.021)	.006 (.017)	.026 (.011)	.022 (.010)
	U_{66}	-.019 (.033)	-.011 (.027)	-.003 (.018)	.005 (.016)
	λ	1	1	1	1
II	n = 307				
	$U_{76,66}$	.061 (.029)	.053 (.025)	.030 (.020)	.036 (.020)
	U_{76}	.026 (.020)	.020 (.017)	.017 (.014)	.025 (.015)
	U_{66}	.016 (.025)	-.002 (.022)	.011 (.018)	-.008 (.009)
	λ	.436 (.050)	.628 (.067)	.684 (.065)	.630 (.066)
	n = 597				
	$U_{76,66}$	.026 (.030)	.032 (.027)	.030 (.018)	.038 (.018)
	U_{76}	.007 (.016)	.011 (.015)	.022 (.010)	.026 (.011)
	U_{66}	-.012 (.027)	-.022 (.024)	-.003 (.016)	-.004 (.009)
	λ	.399 (.034)	.465 (.043)	.604 (.042)	.574 (.041)

*Standard errors are in parentheses.

variable obtains negative coefficients in many cases not dissimilar in magnitude to those for the U_{76} group, the biggest effects are on cities with contracts in both years. The striking pattern revealed in Table 2 remains, even with base year wages or salaries controlled.

Part II of Table 4 presents the contract coefficients for the case of a variable λ. Although λ's range from .40 to .68 in the OLS estimators for the four dependent variables examined, the union coefficients are nearly identical in three cases with those from the $\lambda=1$ estimators. In the fourth case, with hourly wage as the dependent variable, the $U_{76,66}$ group shows an even greater advantage over the other groups than in the $\lambda=1$ case. As presented in Section III, however, there is a bias in the variable λ which needs to be estimated.

Regressing $\ln W_{66}$ on the three contract dummies and the other control variables yields $r_{W_{66}\, U\cdot\vec{X}}$'s which range from .0003 to .089. Thus for all of the three contract coefficients for any wage or salary dependent variable, $R^2_{W_{66}\, U\cdot\vec{X}}$ bounds at .008, or effectively zero. Allowing P from section III to be .50 as an upper bound, we see that $\hat{\lambda}$ (Table 4, part II) will underestimate λ by a factor of 2 at most. Twice the calculated λ's will yield estimates of the union effect close to those obtained with λ fixed at 1, the case already examined. Since λ's will range between the calculated λ's and $\lambda=1$, and since the contract coefficients vary only slightly between when these upper and lower bounds for λ are used, we conclude that the biases in the contract coefficients as presented in Table 4, part II are also slight. A comparison of the coefficients in Table 4 with those in Table 2 shows that the estimates of the coefficients on unionism are only slightly affected by different allowances for the omitted factor. Although in the hourly wage regressions there is some decrease in the union coefficients with the addition of a control for base period wages, union coefficients in the salary regressions are virtually unaffected by the addition of such a control.

The most novel result in the calculations thus far is the much greater impact of unionism when contracts are signed in 1966 and 1976 than in the other cases. What explains this result? The most likely explanation is that the trade union effect differs notably by period of time and/or by type of city, with either those organized earlier being more prone to union influence or for the union effect to increase with time organized. More broadly, the results suggest that there is no single union compensation effect, but rather that effects differ depending on the environment in which the union and city find themselves. In the remainder of this study we examine the interrelation between union wages and 'environmental' factors in greater detail.

The nature of the union effect

One way of examining the interrelation between unionism and environmental factors is to estimate separate wage functions for contract and noncontract cities and to compare the resulting coefficients on variables. When the effect of a variable is larger in the contract sample, this means that it is conducive to a greater union wage effect and conversely when it is lower in the union sector.

Table 5 compares coefficients on selected variables in 1966 and 1976. It shows clearly that the earnings functions of the union and nonunion cities are quite different. However, from one time period to the next, patterns of coefficients of the contract and noncontract are not consistent. The most consistent pattern from sample to sample, from one wage statistic to another, is that the per capita city revenue coefficients for the contract subsamples are greater than those in the noncontrac subsamples. According to the estimates in column 3 for a contract and a noncontract city starting from a given wage level, a unionized city will tend to increase its fire fighters' wages more for a given increment in per capita city revenue, possibly because the increased revenue results from political pressures by the firefighters. Other noticeable differences in the coefficients include those in median value of housing variable, which tend to be significantly greater for the noncontract sample, and in per capita income, which also are generally greater in the noncontract equations.

Table 5: Comparison of Selected Coefficients from

Contract and NonContract Earnings Functions for 1966 and 1976

(Standard Errors in Parentheses)

Independent Variables:	1966 City Revenue K	1966 City Revenue NK	1966 Median Value of Housing K	1966 Median Value of Housing NK	1976 City Revenue K	1976 City Revenue NK	1976 Median Value of Housing K	1976 Median Value of Housing NK	1976 Per Capita Income K	1976 Per Capita Income NK
Dependent Variables:										
Total Compensation	--	--	--	--	.094 (.027)	.026 (.022)	.151 (.067)	.125 (.077)	-.033 (.097)	.113 (.093)
Salary	.053 (.029)	.056 (.016)	.119 (.053)	.161 (.030)	.075 (.024)	.020 (.021)	.160 (.060)	.166 (.075)	-.023 (.086)	.082 (.089)
Hourly Wage	.044 (.048)	.087 (.022)	.137 (.086)	.201 (.041)	.100 (.026)	.034 (.025)	.134 (.067)	.175 (.086)	.034 (.094)	.075 (.104)
Maximum Salary	.079 (.024)	.063 (.011)	.121 (.044)	.172 (.021)	.044 (.018)	.014 (.010)	.098 (.045)	.119 (.037)	.061 (.063)	.113 (.045)
Entrance Salary	.073 (.028)	.035 (.012)	.016 (.050)	.123 (.022)	.053 (.020)	-.0004 (.010)	.106 (.051)	.097 (.035)	-.069 (.071)	.076 (.042)
Retirement Experience (in $)	--	--	--	--	445.80 (202.34)	83.87 (88.65)	375.85 (511.44)	-358.15 (310.85)	3.02 (720.53)	578.53 (373.79
Insurance Experience (in $)	--	--	--	--	34.54 (44.71)	11.94 (29.97)	-126.60 (113.01)	41.79 (105.09)	-74.90 (159.21)	29.61 (126.36

	Contract Samples	Non-Contract Samples
1966:	n = 76	n = 523
1976:	n = 287	n = 728
Mean Retirement Expenditures:	$1745.70	$1302.70
Mean Insurance Expenditures:	$489.23	$370.75

Next, the separate earnings equations can be used to estimate the contract effect, by examining what contract (noncontract) wages would be if members were paid according to the other equations, as indicated in (9) and (10) of section III. These estimates presented in Table 6 reinforce the earlier results revealing moderate union effects on total compensation and straight-time pay and large effects on fringes in 1976 (5.2 percent to 8.4 percent on retirement benefits; 14.7 percent to 16.1 percent on insurance benefits). The 1966 contract effects are near zero and usually negative.

Finally, are there any differences in union effects depending on the governmental structure of cities? Are mayor-council cities more influenced by unionism, because of less continuity of the political decision makers?

To examine this possibility, separate regressions were run for the two major government types, mayor-council (MC, n=356) and council-manager (CM, n=657) selected from the entire population of cities for 1976. No significant differences exist by government type: contract coefficients in salary regressions were 3.9 percent and 4.1 percent for MC and CM cities respectively, as compared to 3.1 percent for the entire n=1015 sample. Also consistent with results for the entire sample are the MC and CM contract coefficients in the total compensation equations -- both about 1 percent larger than their corresponding salary coefficients. These slightly larger compensation coefficients again signal greater impact in the area of fringes. MC and CM contract coefficients (with the level of compensation controlled) for all fringes are 8.5 percent and 7.4 percent, consistent with the 9.4 percent for the total 1976 sample. However, contract coefficients for these two government subsamples are slightly less than the coefficient for the entire sample in retirement benefit regressions (MC - 4.6%; CM - 4.2%; all - 6.1%), and somewhat greater in insurance benefit regressions (MC - 26.9%; CM - 19.1%; all - 18.4%).

Table 6:

Estimates of Union Differentials Using Contract and Non-Contract Earnings Functions

ln(wage statistic)	1 Average for Contract Cities	2 Average for Non-Contract Cities	3 Union Paid by Nonunion	4 Nonunion Paid by Union	5 (1-3)	6 (4-2)
1966	n = 76	n = 521				
Salary	8.583	8.682	8.701	8.678	-.018	-.004
Hourly Wage	.670	.583	.661	.630	.009	.047
Entrance Salary	8.542	8.534	8.545	8.541	-.003	.007
Maximum Salary	8.674	8.687	8.693	8.657	-.019	-.03
1976	n = 287	n = 728				
Total Compensation	9.557	9.397	9.521	9.480	.036	.083
Salary	9.396	9.250	9.372	9.307	.024	.057
Hourly Wage	1.489	1.324	1.464	1.382	.025	.058
Entrance Salary	9.186	9.086	9.168	9.124	.018	.038
Maximum Salary	9.358	9.257	9.348	9.299	.010	.042
Retirement (in $)	1745.70	1302.70	1599.06	1370.73	.084[a]	.052[a]
Insurance (in $)	489.23	370.75	417.13	430.31	.147[b]	.161[b]

[a]Calculated by $\frac{\text{column 1 - column 3}}{\text{column 1}}$

[b]Calculated by $\frac{\text{column 4 - column 2}}{\text{column 2}}$

Conclusion

This study investigated the International Association of Fire Fighters as a case study of public sector unions. With a larger number of observations, and a larger set of variables, cross-section results show a small effect (local or contract) on salaries in both 1966 and 1976: from -1.6 percent to 3.4 percent and larger effects in fringes and total compensation. The IAFF's effect on fringes is as much as four times its impact on salaries. The IAFF has an especially significant impact on insurance benefits, with as much as an 18.3 percent contract effect. Furthermore, with compensation controlled, the study illustrates how the IAFF significantly alters the composition of the total compensation package itself.

This paper also introduced a superior before/after methodology to the study of union differentials in wages and salaries. For the firefighters, the before/after model confirms the cross-section results.

Finally and perhaps most importantly, we have found significant differences in union effects depending on economic environments, notably the length of the contractual arrangement, and other factors, which suggest the need for more detailed study of the diversity of union effects in the public sector.

FOOTNOTES

[1]Since the various wage statistics are derived from city expenditures, the figures may not be accurate measures of firefighters' compensation in unionized cities. If negotiations (and possible mediation, fact finding, repudiation, and court action) are not completed by the end of a fiscal year, these statistics will understate the eventual city expenditures for that year. Also, depending on methods of reporting, retroactive settlements will cause such expenditure statistics to overstate the true compensations in union cities for a given fiscal year.

[2]There are missing values for some independent variables, especially for smaller cities. So as not to lose cities from the samples, missing values are estimated in a step-wise fashion based on the method described by Griliches, Hall and Hauseman in "Missing Data and Self-Selection in Large Panels." While not perfect, these estimates are reasonable approximations for the natural logarithm of the given characteristic with the available information. Those characteristics missing for only a few observations, $\vec{Y}_{m1}$, are regressed on those independent variables which are present for the entire population, $\vec{Y}_{nm}$.

$$\vec{Y}_{m1} = a + B_1 (\vec{Y}_{nm})$$

Once the few missing values are estimated, a second set of characteristics (which are missing for a slightly larger group of cities, $\vec{Y}_{m2}$, are regressed on the now larger set of non-missing information.

$$\vec{Y}_{m2} = b + B_2 (\vec{Y}_{nm}) + B_3(\vec{Y}_{m1})$$

In a four-step process, all missing values are estimated with a least squares equation.

[3]The occupation classification "craftsmen and kindred workers," is used for the opportunity wage variable because, out of those general classifications available in the Census of Population for urban places of 10,000 or more, the "craftsmen" class had, on a national level, a composition by race and sec closest to that of the firefighter occupation. The composition by race and sex was the only check of the classification readily available. In dollar terms, in 1976, the average salary of a firefighter was approximately $10,845; the average of the median earnings of craftsmen was $8,113. In 1966, the figures were $5,896 for firefighters, and $5,238 for craftsmen.

[4]Control Variable Data are taken from: The Municipal Yearbook 1966/1976; Census of Population, vol. 1 1960/1970; County and City Data Book, 1967/1972; Fire Record of Cities, 1960/1970, as published by National Fire Protection Association.

[5]There are 84 cities in the "union 1966 - nonunion 1976" category in the n=307 and n=597 samples. The size of this group raises doubts about the consistency of the union data for 1966 and 1976. However, the size of the group is not improbable. IAFF convention reports show that 35 U.S. locals and 36 U.S. locals lost charters in the 1962-1964 and 1964-1966 periods, respectively.

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