The Effect of Unionization on Wages in the Public Sector:  
The Case of Firemen

by

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The recent growth in both the size and the militance of rank-and-file union membership in the public sector has been the cause for increasing public concern over the effects that these unions may have on the level of municipal services and the allocation of resources in urban areas. The purpose of this paper is to examine the effects of one public sector union, The International Association of Fire Fighters, on the wages of its constituent membership and give some indication of the likely consequences of this union's activities for municipal fiscal problems.

Although unions of government employees grew rapidly throughout the 1960's, only about 15 percent of all government employees at the federal, state, and local levels belonged to unions as of 1966. This was significantly below the level of unionization in the aggregate economy (about 30 percent of all nonagricultural employment) and very far below the level of unionization in manufacturing (about 50 percent of employment). Furthermore, even though it is widely agreed that the rank-and-file militance of unionized workers in the public sector increased throughout the 1960's, by 1967 government employees involved in work stoppages still constituted less than 5 percent of all workers so involved, although they made up some 9 percent of union membership. One implication of these data is that there is still

2Ibid., Table 144.
a significant proportion of public employees remaining unorganized. It
would be helpful therefore to know the consequences of unionization for the
wages of workers in the public sector who are likely to become unionized in
the future. A second implication of these data is that despite the increase
in conventional forms of collective bargaining in the public sector most of
the unions in this sector, unlike unions in the private sector, do not derive
their economic strength solely from the threat of a strike. Unions in the
public sector still depend on legislative action, lobbying, and other activi-
ties far more heavily than do unions in the private sector. Hence, it
seems unlikely that conclusions about the effects of unions on the wages of
workers in the private sector should be carried over directly to the public
sector without further investigation.3/

For a discussion of other areas where union operations differ as between
the public and private sectors see George H. Eldebrand, "The Public Sector," pp. 125-154 in Dunlop and Chamberlain (eds.), Frontiers of Collective Bar-
gaining, New York 1967. Apparently the only study of union effects on
wages which deals with an area including government employees is Helvin
a study of municipal transit, but it is difficult to tell from the data he
presents how many of the cities included in his data had transit operations
run governmentally. He concludes that unionism raised the average wage of
unionized transit workers relative to nonunion transit workers by "...15 to
20 percent for the 1920's, 20 to 25 percent for the early 1930's, 5 to
10 percent for the late 1930's, and something less than 13 percent for
the 1940's," ibid., p. 572. Using "rough and ready" methods H.G. Lewis
estimates that the unionization effect in this industry was not less, and
perhaps more, in the late 1950's than in the late 1940's, which would imply
an estimate in the range 12 to 22 percent for 1958. See H.G. Lewis, Unionism
and Relative Wages in the United States, Chicago 1963. The estimate for
1958 is slightly higher than Lewis' estimate for the whole economy in the
late 1950's, which implies that the transit workers were a very effective
union.
The plan of the paper is as follows: Section I contains information on the International Association of Fire Fighters (IAFF), dominant union organizing firemen in the U.S., including evidence on the methods which have been used by locals of the IAFF for the purpose of increasing their wages. Section II contains a discussion of the conceptual framework we use for measurement of the effect of the IAFF on the relative wages of firemen. Section III is a discussion of the data we use and our empirical results, while Section IV contains some analysis of the economic implications of the results and a few concluding remarks.

I. The IAFF

In many ways the International Association of Fire Fighters provides the most fruitful example for the study of union wage effects of those unions that organize primarily within the public sector.\footnote{In 1966 the IAFF claimed 115,000 members, making it the 42nd largest union in the U.S., but 6th in size among those unions that organize primarily governmental workers. The other major unions of government employees (and their reported membership in 1966) are: The American Federation of State, County, and Municipal Employees (AFSCME - 281, 777), the American Federation of Government Employees (AFGE - 199, 323), The National Association of Letter Carriers (NALC - 199, 628), The United Federation of Postal Clerks (UFPC - 143, 146), and the American Federation of Teachers (AFT - 125,000). Of these unions AFSCME and the AFGE have very heterogeneous memberships, making union-nonunion comparisons difficult; the NALC and UFPC organize workers in the post office department and thus are forced because of the uniformity of civil service pay schedules to lobby (bargain) for employees who do not belong to the union, making union-nonunion comparisons very difficult; and the AFT has organized only a relatively small fraction of all teachers.}

First, firemen are a relatively homogeneous group of workers. For example, in 1960 97.5 percent of firemen in the U.S. were white, 43 percent were high school graduates and of those who
were not high school graduates some 50 percent had one to three years of high school.\(^5\) In addition, the basic training of firemen across the U.S. has been standardized for many years.\(^6\) Finally, the IAFF organizes only uniformed employees of fire departments, so that the job duties of union members differ only slightly from city to city.\(^7\) Since the IAFF had not organized the firemen in many U.S. cities during the 1960's it is possible to make cross-sectional comparisons between union and nonunion cities, and these comparisons are greatly facilitated by the homogeneity of this occupational class.

Second, the demand for the services of firemen within a given city is probably less elastic than the demand for comparably skilled labor in the private sector. This characteristic of the demand for the labor of firemen is presumably similar in nature to the case of other services provided governmentally, e.g., police protection, sanitation services, and others. There are several reasons for expecting demand inelasticity: (a) The nature of the demand for the output of fire departments, call it fire protection services, is likely to make the demand for these services price inelastic. On the one hand, the majority of fire protection services are

\(^5\)U.S. Census of Population 1960, U.S. Summary, Final Report PC(1) - ID, Table 204.


provided locally and by the municipal government. Hence there is little fear that the services provided by union labor, once the fire department is unionized, will face competition from services provided by nonunion labor. On the other hand, reductions in the employment of fire department personnel generally affect the fire department insurance ratings in a city, and thus the overall fire insurance rating in a city. The overall fire insurance rating in a city is the key determinant of the basis rate for fire insurance, so that personnel reductions within fire departments have an important and sometimes dramatic effect on fire insurance premiums.2 (b) The extent to which other factors may be substituted for labor in the production of fire services is severely limited.2 (c) Although personnel expenditures make up some 90 to 95 percent of total fire department expenses, the latter make up only about 5 percent of total municipal expenditures. Hence, increases in firemen's wages are likely to have only a small effect on tax rates.

2 One extreme example is recorded in Municipal Fire Administration, p. 25. Fire-fighting forces in one city were cut in half in order to effect a tax saving which turned out to be only one-fourth the size of the city-wide increase in fire insurance premiums that resulted from the cut. The fire department's rating apparently counts as approximately one-third of the total rating of a city's fire defenses. It follows in importance only the water supply. Extensive discussion of the effect of fire defenses on insurance ratings is contained in ibid., chapter 2.

2 There is some evidence that fire insurance requirements also tend to reduce (or eliminate) whatever capital-labor substitution there may be in fire service production. In particular, the National Board of Fire Underwriters' standard grading schedule provides a minimum number of men on duty for each type of fire equipment. See Municipal Fire Administration, Appendix A.
Finally, the locals of the IAFF have always depended heavily upon political pressure and lobbying activities in their efforts to improve wages and working conditions. The extent to which the leadership has advocated or tolerated strikes or strike threats has varied over the years with the sentiments of rank-and-file union members. The IAFF was formed in 1918 amid the general militance and growth of trade unionism that resulted from the phenomenal price increases associated with the beginning of World War I.\(^{10}\) From the very beginning, however, IAFF officials apparently believed that more was to be gained by political activity than by open strike threats. In the face of strong demands from some locals, however, the original IAFF constitution only deemed it "inadvisable to strike or take active part in strikes..."\(^11\) By 1930 this provision of the IAFF constitution was strengthened to read: "We shall not strike or take part in any sympathetic strikes..."\(^12\) There had, in fact, been no recorded strikes by firemen since 1975, and there was apparently little sentiment for this method of approach.

\(^{10}\) See Orley Ashenfelter and John P.encavel, "American Trade Union Growth: 1900-1960," The Quarterly Journal of Economics, August 1969, pp. 434-448. Over the years 1917-20 the average annual rate of change of prices was on the order of 14 percent.

\(^{11}\) From the IAFF Constitution of 1917, Article XVI, Section 1.

\(^{12}\) IAFF Constitution, Edition of 1931, Article III, Section 2.
By 1966, in response to rising pressure from rank-and-file firemen, the IAFF formed a commission to study the advisability of eliminating the no-strike clause from its constitution. After considerable debate the no-strike clause was removed from the IAFF constitution at its 1968 convention.

In view of the indirect methods by which the IAFF has sought to influence wages one is led to question the existence of even prima facie evidence that it has done so. For the IAFF is a very de-centralized union, having only 14 members in its national staff and monthly dues of some $.65.\textsuperscript{13} In fact, there is considerable casual evidence that some of the IAFF's locals have been very effective economic organizations. Leaders of the IAFF locals have generally been well aware of the widespread antipathy on the part of the general public to strikes by uniformed organizations. Rather than provoke this anger and fear the leadership of the IAFF has generally enlisted the aid of the public, at the polls and elsewhere, to impress upon recalcitrant municipal leaders how just their demands are.\textsuperscript{14} There is one

\textsuperscript{13}See J. Joseph Loevenberg, \textit{op. cit.}

\textsuperscript{14}As early as 1927 IAFF locals in Montana turned directly to referendums at the polls to affect their salaries and hours and were eventually successful at obtaining a state-wide statute guaranteeing the 8-hour day for firemen. See Arthur L. Quin, "Montana First State in Country to Establish 8-Hour Day for Fire Fighter," \textit{International Fire Fighter}, April 1935, p. 3. More recently Jersey City local No. 1066 won the forty-two hour week by referendum at the polls in 1959. See Edward Riebenell, "Jersey City Gains 42-Hour Week in History-Making Referendum," \textit{International Fire Fighter}, June 1959, p.2. Finally, the most recently used tactic is to ask the public for pay parity with the police, even though the relative wage of firemen implied by parity generally results in an excess supply of applicants for positions as firemen and acute shortages of applicants for positions as policemen. Firemen have been able to win parity by political measures in recent years in Los Angeles, San Francisco, and Berkeley. See \textit{Police Collective Bargaining}, Public Personnel Association, Chicago 1969.
unique characteristic of the fireman's occupation that is likely to make this a more successful tactic for firemen than for workers in other industries or occupations. In particular, the fireman does not spend most of his time answering fire calls, but rather on duty waiting to answer such calls. This has led to much longer "duty" hours for firemen than in virtually any other occupation or industry. Although a 56 hour week was nearer the "norm" for firemen in 1969, 65 and even 72 hour work weeks were not unheard of. The general public appears very sympathetic, therefore, to pleas for reductions in the weekly hours of firemen. In view of this, it is not surprising that a large portion of the successful public appeals made by IAFF locals have involved reductions in hours. Weekly hours reductions without compensating declines in annual salaries, of course, are tantamount to increases in hourly wages. In the sequel, therefore, we will test for the effect of unionism on both the annual salary and the weekly hours of firemen. Our expectation is that the latter are more likely to be affected by unionism than is the former.

II. Conceptual Framework

One fruitful procedure for analyzing the effect of unionism on wages is based on the distinction in the $i$th city among the wage of unionized workers ($w^U_i$), the wage of nonunion workers ($w^N_i$), and the wage that would prevail for all workers if there were no unionism ($w^C_i$). The proportionate effect of unionism on the wage of unionized workers in the $i$th city is $R^U_i = (w^U_i - w^C_i)/w^C_i$. $R^U_i$ measures the proportion by which unionism raises (or lowers) the union wage above (below) what it would
be in the absence of unionism. The proportionate effect of unionism on the wage of nonunion workers in the $i$th city is $R_i^n \frac{(u_i^R - u_i^C)}{u_i^C}$. $R^n$ measures the proportion by which unionism changes (raises or lowers) the nonunion wage from what it would be in the absence of unionism. Finally, the proportionate effect of unionism on the union wage relative to the nonunion wage in the $i$th city is

$$M_i = \frac{(u_i^U - u_i^C)}{u_i^C} = \frac{(1 + R_i^U) - (1 + R_i^C)}{(1 + R_i^C)^n},$$

so that

$$\ln(1 + M_i) = \ln(1 + R_i^U) - \ln(1 + R_i^C).$$

Now the observed average wage in the $i$th city in the presence of unionism ($u_i$) may be taken as a geometric weighted average of the prevailing union and nonunion wages in the city so that

$$1 = \ln u_i = U_i \ln u_i^U + (1 - U_i) \ln u_i^C,$$

where $U_i$ is the fraction of workers unionized. Adding the identity

$$\ln u_i^C - U_i \ln u_i^C - (1 - U_i) \ln u_i^C = 0$$

to the right side of (1) gives

$$1 = \ln u_i = \ln u_i^C + U_i \ln(1 + R_i^U) + (1 - U_i) \ln(1 + R_i^C)$$

$$= \ln u_i^C + \ln(1 + R_i^U) + \ln(1 + M_i) U_i.$$

$R^n$ might be positive, for example, if the threat of unionism induced employers to buy off the threat with higher wages. $R^n$ might be negative, for example, if wages were raised in the union sector and this forced employees out of that sector and into the nonunion sector, thereby bidding wages below their levels in the absence of unionism. All of this is discussed in great detail in Lewis, op. cit., pp. 27-40.
Although $\tilde{U}_1$ and $U_1$ are in general observable, equation (2) clearly is not susceptible of estimation as it stands. Instead, it is necessary to estimate an equation of the form

$$\ln \tilde{U}_1 = \ln U_1^C + \gamma U_1 + Z_1$$

where $\ln U_1^C$ would be replaced by its observable determinants, $\gamma$ is a parameter, and $Z_1$ is a residual error. It is natural, therefore, to inquire as to how the least squares estimator of the parameter $\gamma$ in (3) is to be interpreted in terms of the development leading to (2). To see this, one need only write the identity

$$\ln(1 + H_1) \equiv \ln(1 + H) + [\ln(1 + H_1) - \ln(1 + H)]$$

$$\equiv \ln(1 + H) + [\ln(1 + R_1^U) - \ln(1 + R^U)]$$

$$- [\ln(1 + R_1^N) - \ln(1 + R^N)],$$

where $H \equiv (R^U - R^N)/R^U$, i.e.,

the proportion by which the average union wage differs from the average nonunion wage; substitute with (4) for $\ln(1 + H_1)$ in (2); and add the identity $\ln(1 + R_1^U) - \ln(1 + R^U) \equiv 0$ to the right hand side of the result to get:

$$\ln \tilde{U}_1 = \ln U_1^C + \ln(1 + R_1^U) + \ln(1 + \tilde{H}) U_1$$

$$+ U_1 [\ln(1 + R_1^U) - \ln(1 + R^U)]$$

$$+ (1 - U_1) [\ln(1 + R_1^N) - \ln(1 + R^N)].$$
Notice that (5) is identical to (3) if we write the residual, $Z_1$, as

$$Z_1 = U_1 \left[ \ln(1 + R_1^U) - \ln(1 + R_1^N) \right]$$

$$+ [1 - U_1] \left[ \ln(1 + R_1^N) - \ln(1 + R_1^U) \right].$$

hence, the least squares estimator of the parameter $\gamma$ in (3) may be interpreted as an unbiased estimator of $\ln(1 + \bar{W}) = \ln \bar{W}^U - \ln \bar{W}^N$, i.e., the logarithm of the average union wage relative to the average nonunion wage, if $Z_1$ is uncorrelated with $U_1$.\(^{16}\) If $Z_1$ is positively correlated with $U_1$, then $E(\gamma) > \ln(1 + \bar{W})$, and vice versa. The important point is that, in general, a positive correlation between the average wage in a category and the extent of unionism in a category does not necessarily imply that the average wage of union workers is greater than the average wage of nonunion workers.\(^{17}\) Even when $U_1$ and $Z_1$ are uncorrelated, however, the least squares estimator of $\gamma$ in (3) is not the most efficient

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\(^{16}\) This is the conventional requirement that an omitted variable be uncorrelated with the included variables if the coefficient estimators for the included variables are to be unbiased. See A.S. Goldberger, *Econometric Theory*, New York 1964, pp. 196-197.

\(^{17}\) This point is due, of course, to F.C. Lewis, op. cit., pp. 27-29. To make the point more clearly consider the following example. Suppose there were three equally sized and partially unionized cities and that $(\ln W_1^U - \ln W_1^N)$ and $U_1$ were as follows for the three: \(.10, .75; .0, .50; -.10, .25\). Suppose further that $W_1^U = 1.0$ in all of the cities so that $\ln W_1^N = 0.0$ in all cities. Since $\ln \bar{U}_1 = \ln W_1^N + U_1(\ln W_1^U - \ln W_1^N)$ it follows that $\ln \bar{U}_1$ in the three cities must be \(.075, .0, -.025\). Clearly $\ln W_1^U - \ln W_1^N = .75 (.10) + .5 (.0) - .75 (.10) = .05$, i.e., there is a 5 percent difference between the average wage of union workers and the average wage of nonunion workers. Fitting the equation $\ln \bar{U}_1 = \alpha + \gamma U_1$, however, gives $\gamma = .20$ which would be a badly biased estimate of $\ln(1 + \bar{W}) = .05$. In this case a positive correlation between the $U_1$ and the $U_1$ causes the upward bias.
because $\text{var}(Z_1) \neq$ a constant and instead varies systematically. Although we do not pursue this issue here a more appropriate scheme would be based on a random parameters model which explicitly took account of the heteroscedastic disturbance term in (3).

It is easy to see that there is one condition under which the least squares estimator of $\gamma$ cannot be a badly biased estimator of $\ln(1 + \bar{U})$. This is the case where $U_1$ takes on essentially only two values, 1.0 or 0.0, so that $Z_1$ and $U_1$ are orthogonal by construction.\footnote{\textit{18}} Fortunately for our purposes here this case is a close approximation of the situation for fire departments. In general, where a local of the IAFF exists it contains a very large fraction of all fire department employees. Table I shows the distribution in 1968 of cities by percent of organization in the two population size classes with which we will be concerned. As can be seen from the Table, only about 15% of the reporting cities had less than 90% or more than 0% of their fire protection employees unionized. In cities where there was a local of the IAFF, those employees who were not union members were probably non-uniformed employees. Unfortunately, it will not be possible to separate the wages of these workers from those of uniformed employees and we discuss the size of the bias which might result from this difficulty below.

\begin{equation}
\text{where } T \text{ is the number of observations (cities). Thus } \sum Z_1 U_1 = 0.
\end{equation}
Table I

Distribution of Cities by percentage of Fire Protection Employees Represented by Unions

<table>
<thead>
<tr>
<th>Percent of Cities with Representation of</th>
<th>City Population Size:</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>50,000 to 100,000</td>
</tr>
<tr>
<td>0%</td>
<td>18</td>
</tr>
<tr>
<td>1-24%</td>
<td>1</td>
</tr>
<tr>
<td>25-49%</td>
<td>0</td>
</tr>
<tr>
<td>50-74%</td>
<td>6</td>
</tr>
<tr>
<td>75-99%</td>
<td>9</td>
</tr>
<tr>
<td>90-99%</td>
<td>35</td>
</tr>
<tr>
<td>100%</td>
<td>31</td>
</tr>
</tbody>
</table>

Finally, in order to make equation (3) operational we must specify hypotheses about the way in which $\ln W^c_i$ is determined. If we assume that the supply of labor to a fire department within a city is perfectly elastic at a real wage of $\alpha e^{\epsilon_1}$, where $\epsilon_1$ is a disturbance term, then

$$\ln W^c_i = \alpha e^{\epsilon_1},$$

or

$$\ln W^c_i = \ln \alpha + \ln P_i + \epsilon_1,$$

where $P_i$ is a price index. Equation (6) implies that in the absence of unionism differences in money wages across cities would, apart from a disturbance term, be due primarily to differences in living costs.\textsuperscript{19} Substitution of (6) into (3) gives the estimating equation:

$$\ln \tilde{W}_i = \ln \alpha + \ln P_i + \ln(1 + \tilde{M}) U_i + (Z_i + \epsilon_i).$$

\textsuperscript{19}This is the so-called "competitive" hypothesis. The standard reference for a statement of this hypothesis is Melvin W. Reder, "Wage Differentials: Theory and Measurement," in *Aspects of Labor Economics*, NBER 1962.
We assume that $\epsilon_i$ and $U_i$ are uncorrelated.\footnote{This assumption is not so innocuous as it seems, and it has been questioned in Orley Ashenfelter and George E. Johnson, "Unionism, Relative Wages, and Labor Quality in U.S. Manufacturing Industries," Working Paper No. 9, May 1970, Industrial Relations Section, Princeton University. If, for example, workers who earn high wages tend to join unions because of their desire to purchase union services, then $\epsilon_i$ and $U_i$ would be positively correlated and the least squares estimator of $\ln(1 + \frac{U_i}{W_i})$ would tend to be biased upward. Intuitively, we would be giving some of the credit for the effect of wages on unionism to the effect of unionism on wages. This bias is likely to be largest where the non-pecuniary benefits of unionism tend to be high, e.g., in the industrial unionism of manufacturing industries. This problem is less likely to arise here, however, because most fire departments operate under a set of civil service rules which, even in the absence of unionism, tend to provide the same services that a union would. See ibid. for further discussion.}
III. Data and Empirical Results

For the years 1961-1966 the Municipal Year Book provides data on total fire department wage and salary expenditures ($), the number of full-time, paid employees ($\text{F}$), and the number of duty hours per week ($H$) for a large number of cities in several population size classes. An estimate of the average hourly wage of fire department employees is then

$$\bar{W}_i = \left[ S_i / (52 \cdot E_i) \right],$$

so that we may decompose $\ln\bar{W}_i$ as

$$\ln\bar{W}_i = \ln(S_i / E_i) - \ln(52 \cdot H_i).$$

$S_i / E_i$ is an estimate of the annual salary of fire department employees while $52 \cdot H_i$ is an estimate of the annual hours of work required to obtain that salary. It follows from (8) that our estimating equation, (7), may be decomposed as:

$$(7a) \quad \ln(S_i / E_i) = \ln\alpha_o^s + \alpha_1^s \ln F_i + \ln(1 + H_i)^s U_i + \epsilon_i^s$$

and

$$(7b) \quad \ln(52 \cdot H_i) = \ln\alpha_o^h + \alpha_1^h \ln F_i + \ln(1 + H_i)^h U_i + \epsilon_i^h,$$

where

$$\ln\alpha_o^s - \ln\alpha_o^h = \ln\alpha_o, \quad \alpha_1^s - \alpha_1^h = \alpha_1, \quad \ln(1 + H_i)^s - \ln(1 + H_i)^h$$

$$= \ln(1 + H_i), \quad \epsilon_i^s - \epsilon_i^h = \epsilon_i + z_i.$$

Since linear transformations of least squares estimators are also least squares estimators it follows that the estimators of the parameters in (7a) and (7b) will satisfy the restrictions imposed by (7), e.g., the estimates
of \( \ln(1 + \bar{N})^g \) and \(-\ln(1 + \bar{N})^h \) in (7a) and (7b) will add up to the estimate of \( \ln(1 + \bar{N}) \) in (7). From (7a) and (7b) we may thus obtain separate estimates of the union/nonunion effects on annual salaries \([ \ln(1 + \bar{N})^g \])\) and annual hours \([ \ln(1 + \bar{N})^h \]).

Estimates of the parameters \( \ln(1 + \bar{N})^g \), \( \ln(1 + \bar{N})^g \), and \( \ln(1 + \bar{N})^h \) are reported for a random sample of cities in two city population size classes in Tables II and III for each of the years from 1961 through 1966. It was necessary to narrow the data base down to what is represented in these tables for the following reasons: (1) By 1968 some 91 percent of all fire protection employees in cities of population 100,000 or larger were unionized.\(^{21}\) Thus it was necessary to use the smaller cities represented in these tables in order to obtain a reliable estimate of nonunion wages. It was also necessary to standardize by city size so as to control both for possible differences in the hazards of fire fighting as between different sizes of cities and for any differences in the non-pecuniary attributes of cities that may be associated with their size. (2) A random sample of cities had to be used rather than the set of all cities listed in the Municipal Year Book because of the difficulty of determining whether a given city was unionized in a given year. \( U_1 \) has been assigned the value 1.0 or 0.0 depending on whether the city was unionized or not as of the given date. Discerning unionization was the most difficult part of the data construction and the methods used, as well as the other parts of the data construction, are more fully described in the appendix.

\(^{21}\) Stieber, *op. cit.*, Table 4.
<table>
<thead>
<tr>
<th>Year</th>
<th>Average Hourly Earnings of Full-Time Year-Round Production and Nonsupervisory Workers</th>
<th>Effect on Annual Earnings</th>
<th>Effect on Nonsupervisory Annual Earnings</th>
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<td>1961</td>
<td>$520.00</td>
<td>$144.00</td>
<td>$250.00</td>
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<td>$225.00</td>
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Note: All figures are in dollars.
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<th>Year</th>
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<th>-1.02</th>
<th>-0.96</th>
<th>-0.35</th>
<th>0.30</th>
<th>0.60</th>
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<td>1966</td>
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Table II

An asterisk (*) is placed next to all coefficients that are not significantly different from zero at the .05 level on a two-tailed test.

Weekly Hours in Cities of Population 50,000 - 100,000: 1961 - 1966

Effect of Urbanization on Hourly Hours, Annual Series
Explained variance ($R^2$) for the typical equation underlying the results in Tables II and III was in the neighborhood of .5, with most $R^2$'s falling in the range .4 to .6. As can be seen from the fact that $U_{i}$ is not terribly significant, most of the explanatory power in these equations comes from the cost-of-living variable. The latter generally had a ratio of coefficient/standard error of around 10.0 to 12.0. Table II contains estimates of the unionism effects for cities in the population class 50,000 - 100,000. Results are reported for "all" cities in each yearly sample and for a subset of non-southern cities. The latter set of estimates may provide more control over labor quality variation between union and nonunion cities as well as a better control over cost-of-living differences. These are particularly important factors to control for since they are likely to be correlated with extent of unionism and the latter is highly correlated with geographical region. As can be seen from the table, for both samples $M^u$ tends to be negative, as expected. The estimated values imply that the average duty hours worked in union cities are about 4 percent smaller than in nonunion cities. On the other hand, $M^g$ also tends to be negative, which implies that firemen in unionized cities receive lower annual salaries as well. The average union/nonunion effect for hourly wages thus turns out to be very small, essentially zero in the non-southern sample and about

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22 By 1968 about 33 percent of all fire protection employees in the non-South were unionized, while the figure for the South was about 60 percent. See Stieber, Ibid.

23 There is no point in obtaining the anti-logarithms of the coefficients in these tables because they are so small. That is, $\ln(1 + \hat{M})$ is very close to $\hat{M}$ so long as $\hat{M} \approx .15$. 

4 percent in the all-cities sample. Virtually none of the estimates of \( \tilde{\pi}^a \), \( \tilde{\pi}^h \), or \( \tilde{\pi} \) are significantly different from zero at conventional test levels however.

Table III contains estimated union effects for cities in the size category of 25,000 to 50,000. Again, all of the estimates of \( \tilde{\pi}^h \) are negative, but for this city size most of the estimates are significantly different from zero. Although the estimates vary by year and between the two samples all of them fall between -.03 and -.11, implying that union duty hours averaged some 3 to 11 percent less than nonunion duty hours. Unlike the previous case the estimates of \( \tilde{\pi}^a \) are positive, although in most cases they would not be judged significantly different from zero by conventional standards. It follows, therefore, that in this city size class the effect of unionism is to lower duty hours, but without producing any compensating decline in annual salary. The resulting estimates of \( \tilde{\pi} \) average .095 for the non-southern sample and vary between .062 and .125. The comparable average of the \( \tilde{\pi} \) estimates for the all-cities sample is slightly higher at .136. Only two of the twelve estimates of \( \tilde{\pi} \) are not significantly different from zero at the .05 level.

Finally, there are strong indications of an increase in \( \tilde{\pi} \) in the last year of the sample, 1966. In particular, in Table II \( \tilde{\pi} \) takes on its largest values in 1966 and is significantly different from zero for the all-cities sample. Further, \( \tilde{\pi}^h \) takes on its greatest values in 1966 and is significantly different from zero in both samples. Indeed, 1966 is the only year in which an estimate of \( \tilde{\pi} \) or \( \tilde{\pi}^h \) in Table II is significantly different from zero. In Table III the largest estimates for \( \tilde{\pi} \) are also
for 1966. Perhaps more important, in 1966 the estimates of $\hat{\pi}^e$ are also significantly different from zero. Indeed, 1966 is the first year for which the most important component of the estimated $\hat{\mu}$ is the large and positive $\hat{\mu}^e$.

IV. Concluding Remarks

Our empirical results suggest that in 1966 the unionization of firemen may have raised the average hourly wage of unionized firemen by somewhere between 6 and 16 percent above what it would have been in the absence of the unionization of firemen.\footnote{The implicit assumption in this statement is that $\hat{\mu}$ may be taken to equal $\hat{\mu}^U$, which implies that $\hat{\mu}^p q^o$, i.e., that the effect of the presence of the unionization of firemen had a negligible effect on the wages of nonunion firemen. Since firemen are organized on a strictly local basis this does not seem implausible for the city size categories that we use to estimate $\hat{\mu}$.} The wage effect of unionization appears to have been distinctly higher in 1966 than in any of the years 1961 through 1965. Interestingly enough, the rank-and-file initiative in the IAFF which led to the abandonment of the union's no-strike clause in 1966 began in 1966. This implies that the abandonment of the no-strike clause should probably be taken as the consequence and not the cause of increased rank-and-file militance. It would be interesting to know whether the decision by other unions in the public sector to use the strike as a bargaining tactic has the same origins, for this would have important implications for public policies regarding strikes by public employees.

Two important questions remain. First, to what extent are the union wage effects we observe for firemen typical of the wage effects of other
unions in the public sector? Only further evidence can answer this question satisfactorily, but it seems as plausible as not that the wage effects of the IAFF may be larger than those of the "typical" union in the public sector. On the one hand, firemen are a homogeneous, craft-like occupation whose members have always tended to be closely knit and well organized. This is not the case of most government workers at either the federal, state, or local level, with the possible exception of the police. In addition, expenditures for firemen's wages and salaries make up a relatively small fraction of municipal expenditures. Finally, because of the long periods of inaction which firemen experience they have traditionally had duty hours which averaged significantly longer than the conventional 40-hour week. Reductions in hours are thus a favorite bargaining target for local unions because of the public sympathy engendered. Such a bargaining ploy is usually unavailable to the typical union.

Second, given our estimates of the effect of unionization on wages it is interesting to inquire as to the effect that this implies for municipal revenues. The size of the latter is presumably an important factor in the resistance that municipal governments are likely to put up to union wage demands. In 1967-68 cities with a population less than one-half million expended an average of about 9 percent of total general expenditures on fire protection. If we assume that ninety percent of fire protection expenditures are for wages and salaries, then total municipal revenues (expenditures) would have to be between .5 [ = (.06) (.09) (.90)] and .75.

1.3 [= (.16)(.09)(.90)] percent higher in unionized cities than in nonunion cities in order to produce the same amount of fire protection services.26/ Interestingly enough, in cities with a population of 1/2 million or more, fire protection expenditures averaged only about 5 percent of total expenditures. This implies much smaller union effects on municipal revenues, ranging from .3 to .7 percent.

Data Appendix*

The data construction underlying the results in Table II and III required enough approximations that it is necessary to explain the required computations in more detail. As we noted in the text, data were collected for a random sample of cities, and the list of cities used may be obtained from the author.

The Hourly Wage. Given the available data in the Municipal Year Books, two procedures were available for computing an hourly wage. The first, and the one described in the text, involved multiplying the number of full-time, paid employees times the number of duty hours per man per week times the number of weeks per year and dividing that result into the fire department's expenditures on wages and salaries. Data were not available on paid vacations, holidays, or sick leaves so that these are not incorporated into the final figure. Not all of a fire department's wage and salary expenditures go to uniformed personnel, although the latter tend to be the employees who are unionized, if any are. Hence $U_1$ is assigned the

*Mr. Joseph Vybeek bore the prime responsibility for data collection and computations, and I am heavily indebted to him for his efforts.

26/ This assumes that no capital - labor substitution is possible.
value 1.0 in some cases where it should take on a slightly lower value, generally in the range .9 to 1.0. This would tend to bias the estimator of M downward by perhaps 10 percent or less.

Since the Fire Chief is sometimes included and sometimes excluded from the local union we have estimated the Fire Chief's salary in each year and removed it from the computations. The allowance was made by subtracting the chief's salary from the wages-and-salaries expenditure figure and by subtracting one from the number of full-time employees. For 1963 and 1966 the figure for the Fire Chief's salary was unavailable, so the figure for 1962 was used for 1963 and the figure for 1965 was used for 1966. Deleting the Fire Chief's salary generally changed the estimated average hourly wage by less than one percent.

We used the second procedure available for estimating the average hourly wage to delete the computed data for any cities where reporting errors were likely. This procedure involves taking either the entrance salary or the maximum non-promotional salary and dividing it by the number of hours on duty per week multiplied by the number of weeks per year. Whenever the computed average hourly wage obtained by method 1 fell outside the interval computed via method 2 the observation was discarded as a reporting error. This resulted in the elimination of about 3 percent of all observations.

Finally, all cities reporting volunteer fire personnel were deleted from the sample.
Determining Unionization. The most difficult problem in data construction involved determining the unionization status of each city in each year. First, each city was compared with the list of local unions provided in the IAFF's 1969 pamphlet entitled *Fire Department Salaries in the United States and Canada as Reported by Locals*. If the city was listed in this pamphlet it was taken as unionized in 1969. Due to the incomplete listing in the IAFF pamphlet, failure to indicate unionization from this source was not taken as an indication that the city's fire department was not unionized. A second check involved use of the 1966 *Municipal Year Book*, which listed information concerning whether the fire department was unionized by indicating whether the fire department (a) had a contract with the city, (b) had bargaining rights with the city, or (c) had a dues check-off clause with the city. Again, a positive corroboration was taken to mean that the fire department was unionized in 1966; but if the city was not indicated as unionized, this was not taken to mean that a union did not exist in the city since many cities did not report this information. Any cities identified as unionized in 1966 and 1969 were assumed to be unionized in 1966.

A third check was based on a list of all recording secretaries of active IAFF locals given in the March 1958 issue of the IAFF's *International Fire Fighters*. Since this was reportedly a list of all locals, if the city was not listed, it was considered as not unionized in 1953. Any city that was identified from the above results to be unionized in 1958 and 1969 was taken as unionized throughout the sample years 1961 through 1966.
The final step in the identification process concerned those fire departments which became unionized between 1953 and 1966. This involved identifying local unions numbered 1295 through approximately 1600. Dates of unionization were approximated by the dates of the issues of the International Fire Fighter in which the unions were first listed. If a city was unionized in year $t$, it was considered as unionized in year $t+1$ and not unionized in previous years.

After all of the above checks, there remained some 15 cities that had been indicated as unionized by one source, but by no other source. These cities were dropped from the sample.

The Cost of Living ($P$). Estimates of $P$ for each city were made by using the comparative cost-of-living index published by the BLS in Handbook of Labor Statistics, 1969. The index is based on urban costs for a family of four.

Since the index is available for only some forty major metropolitan areas it was necessary to use these data to estimate the index for each city. When the city was close to one of the listed metropolitan areas the index value for the latter was used. In all other cases one of the regional indexes provided for non-metropolitan areas was used.