THE UNION-NONUNION WAGE DIFFERENTIAL: A REPLICATION AND EXTENSION

by

David E. Shulenburger
Associate Professor

Robert A. McLean
Assistant Professor

and

Sara B. Rasch
Doctoral Candidate

All of
The School of Business
University of Kansas

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Introduction

The growth of the nonunion sector of the building construction industry was one of the major developments in labor relations during the 1970's. Given the growth in residential construction volume during that decade, the growth of the nonunion sector was unsurprising. A union-nonunion wage ratio greater than unity would lead the marginal contractor (the last contractor to enter production) to employ nonunion labor, rather than union labor. Further, to the extent that existing contractors can become "double breasted," (that is, operate both union and nonunion operations), existing contractors have an incentive to shift their marginal output (the last contract received) from the union to the nonunion shop.

One must ask, then, why union construction wage growth remained high during the '70's. Kerr (1954) suggested that the construction labor market is sufficiently balkanized that union leaders prefer to set their members' wages on some equity basis and to let the disemployed members move to the (separate) nonunion sector. Thus, supply adjusts to the level of union wages, rather than union wages' adjusting to supply and demand.

If Kerr's balkanization hypothesis is an accurate description of union wage policy, and if union members do move to the nonunion sector, then the supply of nonunion construction labor increases when demand for construction activity is low and the nonunion sector is more sensitive to the business cycle than it would otherwise be. The result, then, would be to lower nonunion wages relative to union wages, inducing more construction activity to move to the nonunion sector. The latter response would serve to moderate the wage reduction in the nonunion sector. Given, then, some marginal rate of technical substitution of nonunion for union labor and a union wage policy predicated on balkanization, each set of labor market conditions should generate a unique equilibrium union-nonunion wage ratio.

Stephen Welch (1980) has generated one set of estimates of the determinants of that wage ratio (albeit without asserting its unique equilibrium characteristics). Our purpose here is to replicate and extend Welch's research, thus investigating further the determinants of that equilibrium. We were able to employ a more extensive data set than that available to Welch, thus extending his work to more recent years. We also extend Welch's research in that our specification incorporates market conditions directly, rather than using dichotomous variables to represent location in individual urban labor markets.

Like Welch, we specify a cross-sectional model of the union-nonunion wage ratio. Thus, we assume that the wage policy of construction craft unions is invariant across locations. Further, we assume that the marginal rate of technical substitution of nonunion for union labor is invariant in cross section. Several factors serve to make that MRTS differ from unity. Union craftsmen may be better trained (given the availability of apprenticeship training in the union sector), making union craftsmen more productive than their nonunion counterparts. Further, union craft work is more clearly standardized than is nonunion work. Thus, nonunion carpenters may perform laborer's work as well as carpentry. Thus, some of the difference in wages may be due to differences in work performed. A nonunitary MRTS will produce an equilibrium wage ratio different from unity, under most economic conditions.

Welch's Specification and Estimates

Welch (1980) estimated an equation of the form:

$$(1)\quad \ln W_{ij} = \alpha_0 + \alpha_1 U_{ij} + \alpha_2 U_{ij}^2 + \beta_2 T_2 + \ldots + \beta_6 T_6 + \gamma_2 C_2 U_{ij} + \ldots \gamma_{15} C_{15} U_{ij} + \varepsilon_{ij}$$

where $\ln W_{ij}$ is the natural log of the ratio of union to nonunion wages;

U is the ratio of unionized to total union and nonunion craftsmen employed;

T is a dummy variable for the trade;

and C is a dummy variable for the city.

He used only 90 of the observations available from the 1973 BLS survey. These observations included six crafts in fifteen cities. His study excludes four crafts for which data were reported. Table 1 below reproduces his equation as estimated.

His results with their highly significant coefficients on unionization, clearly show the proportion of workers organized to be directly related to the log of the union nonunion wage ratio. Taking the partial of his dependent variable with respect to unionization, setting it equal to zero, and solving for U shows that the maximum wage ratio is achieved when 65.68% of the craft is organized.

Welch's study suffers from two flaws, data inadequacy and model misspecification. The data inadequacy occurs because the BLS survey deals with wages of employed workers and not compensation. As Welch points out (page 156), prior research has shown that construction fringe benefits rise with negotiated wages (Gustman and Segal, 1972) and that nonunion wage supplements are smaller than union supplements (Northrup and Foster, 1975). Given this problem (which cannot be corrected with the available data) the peak union-nonunion wage ratio point cannot accurately be calculated. In addition, the BLS data report only employed craftsmen. Ideally, the unionization ratio would be calculated as total construction union members divided by total construction craftsmen. The use of "employed" rather than "total" carries with it the implication that equal proportions of union and nonunion craftsmen are unemployed. No tests of this implication are available.

The misspecification problem is serious. Several factors which might affect the wage ratio are not included in his model. These include institutional, demand, and supply factors which might have impacts on the wage ratio but which may be correlated with the level of unionization. Failure to include them runs the risk of attributing their impact to the unionization variable or other of his variables, which might be correlated with them.

Replication and Extension

We have extended Welch's estimating equation to take the form:

$$(2)\quad \ln W_{ij} = c_0 + c_1 U_{ij} + c_2 U_{ij}^2 + c_3 V_j + c_4 R_j + c_5 M_{ij} + c_6 B_{ij} + c_7 C + \ldots + c_n C + \varepsilon_{ij}$$

lnW--natural log of union nonunion craft wages

U--percent unionized

V--percent change in the volume of construction activity since the previous year

R--general unemployment rate

M--average manufacturing wages divided by craft wages

B--dummy variable for coordinated bargaining participation

C--craft union dummy variables.

Our model modifies Welch's by explicitly including demand, supply, and institutional forces which are hypothesized to affect the union-nonunion wage ratio. Welch's model used city dummy variables as a composite proxy for all of these three effects. This masks the effect of market conditions, which can be measured directly.

Research has consistently found union membership and bargaining power to be consistently positively related to demand. Beginning with Friedman (1951), however, economists have recognized that rapid inflation may be associated with a reduced union wage premium due to the lack of flexibility imposed on the wage process by multiyear contracts. More recently Ashenfelter's (1978) empirical work has found evidence consistent with this hypothesis while Hendrix (1981) concludes that the association between cyclical inflation and industrial concentration variation makes the assignment of causation to contract duration more treacherous. In the present case the advent of wage and price controls reduced the duration of negotiated construction wage contracts to an average of only eighteen months. Such brief duration contracts could hardly lead to much less flexibility of union than nonunion wages. Strong demand should thus increase union more than nonunion wages through increased bargaining power. We measure changes in demand here by the percentage change in the total volume of construction starts (as measured by data collected by F. W. Dodge). This variable is omitted in the 1972 regression as 1971 F. W. Dodge data are unavailable.

Supply is measured by the city's unemployment rate. This, of course, reflects the volume of labor available for construction. As such it better reflects the quality of labor available to the nonunion sector where only skill barriers to entry exist. Thus as the unemployment rate increases nonunion construction wages should be suppressed and the differential between union and nonunion wages wider.

Bowlby (1980) found that construction workers earn a substantial portion of their annual incomes from nonconstruction activity. Indeed 37 percent of construction workers' nonconstruction annual incomes is derived from work in the manufacturing sector alone. This sector is nearly twice as large a source of construction worker earnings than any other nonconstruction sector. When deciding whether to work in the manufacturing sector or construction sector the relative wages of the two sectors must be an important consideration. We include the ratio between a specific craft's negotiated wages and the manufacturing wage rate as an independent variable to capture this effect. Since most construction workers who come out of manufacturing probably go to the nonunion construction sector, that sector and the nonunion sector are part of the same labor market. Thus nonunion construction wages must be set just high enough to make construction a better alternative for workers than manufacturing.

This line of argument suggests that nonunion construction wages are established jointly with manufacturing wages. We thus expect this variable to be inversely related to our formulation of the dependent variable, i.e., the better the unionized craft performs relative to the manufacturing sector the greater the disparity between union and non-union construction wages.

The dummy variable for coordinated bargaining reflects this institutional arrangement. One presumes that both unions and contractors agree to this arrangement in the belief that it is in their self interest to do so. We include the variable to control for this arrangement but adopt no specific hypothesis about its impact on our dependent variable.

Finally craft dummy variables are included in the equation. These variables capture craft specific institutional factors. Some crafts bargain more frequently, have wider labor markets, are leaders, etc.; these factors would have an independent impact on the dependent variable.

Proportion unionized and proportion unionized squared are specified as in Welch (1980). Should a significant positive coefficient be estimated for the unionization variable support will exist for the microeconomic, competing factor argument. Should the coefficient on the unionization variable be insignificant, the balkanization hypothesis will be supported.

The most consistent set of crafts and cities were sampled in 1973 and 1977. The 1977 equation includes percent change in unionization since 1973 as an additional independent variable.

<u>Time Periods and Sample</u>

We estimated the equation for each of the four time periods, 1972, 1973, 1976, and 1977, which the B.L.S. has surveyed to date and include every craft surveyed. The set of cities differ across years and crafts differ across years and within time periods. Our four equations will therefore vary in the number of observations across periods and the number of craft dummy variables included. In each case, carpenters are omitted from the list of craft dummies (to avoid a singular data matrix).

The assortment of years in which surveys were made gives us an opportunity to analyze the effects of wage controls on the wage ratio; 1972 and 1973 were control years while 1976 and 1977 were not. We hypothesize that a smaller amount of the variations in the dependent variable will be accounted for during the controls years. This is consistent with the announced control period intention to reestablish traditional relationships.

The Estimates

Table 2 lists the estimates for Equation (2). The estimates clearly support the balkanization thesis. In no case is there a significant coefficient on the unionization squared variable or the percent change of unionized variable. The only significant coefficient on the unionization variable is in the 1977 regression and that coefficient is of the wrong sign. Simply stated, the evidence presented here is inconsistent with the thesis that the level, or changes in the level, of unionization have an impact on the ratio of union to nonunion wages in construction.

More complete specification of the model illustrates that Welch's finding of a significant relationship was an artifact of underspecification. This finding was foreshadowed by Welch's experimentation with his own model. The first model he reported (which included only U and U^2 had coefficients on U and U^2 which were significant at the .01 level.

When he added city and craft dummy variables the significance level on the union variable coefficients dropped to the .05 level.

It is unlikely that the lack of significance of unionization is due to a limited range problem. The mean percentage unionized in 1977 for the sample was 68.4%, with observations ranging from 4.7% to 99.6%. Similarly, the mean percent change in unionization from 1973 until 1977 was -5.2% with a range of from 92.9% decrease to a 116.4% increase. Unionization on average in 1973 was 63.2%.

The extremely high correlations between percent unionized and its square suggested the possibility of upward biased standard errors and, thus, of significant effects being found insignificant. To test the possibility that the dependent variable was, indeed, sensitive to percentage unionized, we applied the RESET test for specification error discussed by Ramsey (1974, pp. 34).

Equation (2) was estimated without percent unionized and its square. The residuals of that estimation ε_{ij}' were recorded and regressed on the two omitted variables:

$$(3) \quad \varepsilon_{ij}' = \Gamma_0 + \Gamma_1 U_{ij} + \Gamma_2 U_{ij}^2 + U_{ij} .$$

If the removal of U_{ij} and U_{ij}^2 from equation (2) introduces significant misspecification, then the F-test for the entire relationship expressed in Equation (3) will lead to the rejection of the null hypothesis, "no significant relationship."

The results of the RESET tests for 1972, 73, 76, and 77 were fully consistent with the results shown in Table 2 (a tabular summary of those tests is available from the authors on request). That is, for 1972, 1973, and 1976, the omission of percent unionized and percent unionized square would generate no significant underspecification of the model. For 1977, the only year for which percent unionized had a significant coefficient in Table 2, one must reject the hypothesis of "no relationship" for Equation (3).

The large statistically significant negative coefficient on unionization in 1977 is hard to explain. Taken literally it implies that the union-nonunion wage ratio was 50.1% smaller when 100% of the workforce was organized. Given the significance of only 10%, the small sample, and the lack of theory to support the sign, one is tempted to treat the coefficient as a statistical aberation. It does, however, fit into a pattern. For the years 1972, 1973, 1976 and 1977 the coefficients on percent unionized are .544, .339, .210, and -.501, respectively. Admittedly, the first three coefficients are not significant at traditional levels. The pattern of continual decrease might suggest that the best organized locals either moderated wage demands for competitive reasons or that nonunion contractors in their labor markets were willing to pay nearly union wages to get workers to work nonunion. We see this pattern as very weak evidence of a movement back toward an equilibrium.

In 1973 and again in 1976 the coefficient of the manufacturing wage ratio took on a statistically significant negative sign. In the other two years the coefficients were statistically insignificant. This finding is consistent with the argument that nonunion construction wages and manufacturing wages are set in the same labor market.

In all three years for which change in construction volume data was available a positive coefficient was estimated. In 1973 and 1977 the coefficient was statistically significant. Construction unions are apparently able to convert increased demand into a wider wage margin over nonunion workers.

The labor supply variable, unemployment, had the expected significant positive sign only in 1977. An insignificant coefficient was estimated for the other two periods. General labor supply looseness thus contributes to the union sector's differential, but the one out of four time periods for which this happens makes the empirical association quite tenuous.

In no case did coordinated bargaining have a significant impact on the wage ratio. Perhaps there is no practical significance to formal coordination. The communications network of craft unions may well accomplish the same end as does formal coordination.

The craft pattern is quite mixed. The bricklayer coefficient is significant three times, positive in 1972 and negative in 1973 and 1977. This may only reflect a repetitive two-year bargaining pattern.

During the wage control years 1972 and 1973, 18% and 31% of the variation in the dependent variables were accounted for. In the post-controls years, 1976 and 1977, 76% and 42% of the variation was explained. The controls were intended to replace market and institutional forces with "order." They apparently succeeded. No other major pattern of variation is apparent in the independent variables' coefficients across the controls-market era.

Discussion

The very limited set of data available support the balkanization notion. Craft unions do not appear to respond to loss of market by reducing their wage premium over unorganized workers. This finding is bolstered by the apparent interaction of manufacturing and nonunion craft wages. Craft unions can apparently tolerate even high general unemployment without lowering their wage demands.

Even though our results are consistent across time, we choose not to generalize to other time periods or to all crafts within the same periods. The largest set of observations represented in any of our estimates is 98. This is a small fraction of the universe of craft-city observations and was not randomly selected by the Bureau of Labor Statistics.

TABLE 1

Regression Results: Union-Nonunion Wage Ratio on Extent of Trade Union Organization and Selected Dummy Variables (N = 90)

Variables	Estimated Coefficients	t-Statistics
Constant	0.00962	0.069
U	1.15420	2.293**
U^2	-0.87755	-2.107**
T2	0.06647	1.141
T3	0.06550	1.131
T4	0.00495	0.084
T5 (sheet metal)	0.12463	2.185**
T6 (laborers)	0.11329	2.031**
C2U	0.02266	0.273
C3U	0.06495	0.760
C4U	0.01161	0.122
C5U	0.04150	0.506
C6U	0.13785	0.818
C7U	0.04104	0.344
C8U	0.03116	0.256
C9U (Miami)	-0.15626	-1.935*
C10U	-0.06185	-0.383
C11U	-0.05317	-0.501
C12U	0.05892	0.612
C13U	-0.12718	-1.246
C14U (Denver)	-0.15700	-1.957*
C15U	-0.10796	-1.066

$\bar{R}^2 = 0.3653$

S.E.E. = 0.15244

F = 1.92**

*Significant at the .10 level

**Significant at the .05 level

TABLE 2

Dependent Variable: Natural Log of Union/Nonunion Craft Wage

Variable	Year 1972	1973	1976	1977
Coordinated Bargaining	−.070 (1.47)	−.045 (.04)	.036 (.54)	−.027 (.58)
% Unionized	.544 (.73)	.339 (.65)	.210 (.55)	−.501 (1.67)*
$(\%\ \text{Unionized})^2$	−.654 (1.10)	−.357 (.84)	−.389 (1.12)	.092 (.39)
1973–1977 % Change in Unionized	–	–	–	.051 (.82)
Manufacturing Wage Ratio	.209 (.59)	−.568 (2.20)**	−1.055 (5.03)***	−.021 (.089)
Percent Change in Volume of Construction	–	.232 (2.53)**	.072 (1.23)	.133 (1.78)*
Unemployment	.015 (.94)	−.005 (.36)	−.016 (1.10)	.037 (2.014)**
Bricklayers	−.301 (2.042)***	−.242 (3.24)***	−.024 (.23)	−.157 (2.02)**
Cement Masons	−.35 (.46)	.078 (1.23)	.136 (1.56)	.018 (2.58)**
Electricians	.35 (.50)	.073 (1.27)	.172 (2.25)**	.105 (1.91)*
Iron Workers	–	–	.254 (2.14)**	–
Laborers	−.076 (.80)	.228 (3.16)***	.396 (6.18)***	.007 (1.91)*
Painters	–	–	.198 (2.87)**	–
Pipefitters	.052 (.57)	.064 (.67)	−.062 (.77)	.016 (.10)
Plumbers	−.050 (.64)	.032 (.54)	−.029 (.41)	.065 (.29)
Roofers	.054 (.36)	.038 (.64)	−.001 (.00)	−.012 (1.11)
Sheet Metal Workers	.043 (.47)	.167 (2.88)***	.066 (.07)	.119 (2.15)**
Constant	.216	.642	.909	.295
$\bar{R}^2$	.18	.31	.76	.42
F	1.83*	3.84***	6.96***	3.68***
N	48	98	30	64

t-statistics are in parentheses.
*Significant at .10 level.
**Significant at .05 level.
***Significant at .01 level.

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