A Reexamination of British Strike Activity

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1. Introduction

Recent research has verified a number of hypotheses regarding the evidence of strike activity [see e.g., Pencavel, 1970; Bean and Peel, 1974; and Sapsford, 1975]. Namely, it appears that the frequency of strikes is sensitive to the rate of inflation, previous wage increases, the ability of the firm to pay, and unemployment.

This paper examines some of the same issues considered by previous researchers for the period 1959 through 1976 for Great Britain. The results presented here are a refinement of the previously cited works in several respects. Firstly, the quarterly data employed herein cover two rather different periods of modern British economic history. Prior to about mid-1969, Great Britain was characterized by rather low rates of unemployment and inflation, and relative labor tranquility. After mid-1969, we find much higher rates of inflation and unemployment, and generally more labor unrest. Even simple scattergrams of the data used in this study reveal this structural shift. The Pencavel [1970] and Bean and Peel [1974] papers did not encompass these seemingly different periods.

Secondly, because of the time period chosen, this paper disaggregates to a greater degree than Sapsford [1975]; but not to the same extent as Bean and Peel. Because this paper relies on several data sources which have changed classification schemes several times, it includes metals manufacture, mining, and construction in the sample. Pencavel used, additionally, transport.

Lastly, this paper refines the econometric technique of the previous papers. The model parameters are estimated using Zellner's [1962] seemingly unrelated regression technique. Pencavel's approach to the estimation problem was to treat the incidence of strikes in the various industries as independent.

Bean and Peel were forced to pool the annual data on their four industries due to the short period involved in their study. A strictly enforced pooling of the data carries with it the implicit assumption of homogeneity of behavior across industries. The use of seemingly unrelated regression permits one to test not only the homogeneity of the regression coefficients across industries but also the statistical independence of the industries.

Furthermore, the regression results presented here have been corrected for autocorrelation and multicollinearity. The principal findings of previous researchers are qualitatively corroborated here. However, considerable quantitative instability of the estimated regression parameters across industries and over time is found.

2. The Model

Work stoppages are only one method by which labor can demonstrate its discontent with the workplace. Since work slowdowns and overtime bans are neither measurable nor reported, attention is confined to strikes. The model to be proposed below relies on economic variables to explain the total level of strikes. Not all strikes are reported as being the result of wage related disputes. For the period under consideration, the number of wage related disputes as a percent of the total rose from about 40 percent to 64 percent. While wages are not reported as the principal cause of the remaining disputes, one may invoke the compensating wage differential argument in asserting that wage related disputes are under-reported. In addition, the decision by the rank and file to go out on strike over any given issue will be tempered by conditions in the general economy as well as their own industry. For these reasons, the total number of strikes per quarter is chosen as the dependent variable in the economic model of strike behavior.

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Ashenfelter and Johnson [1969] have constructed an optimizing model of the firm that has been used successfully in previous studies. With slight modification the firm is modeled as a present value maximizing entity. It is assumed that they perceive the union's wage concession schedule as

$$Y_A = Y_* + (Y_o - Y_*)e^{-as} \qquad (2.1)$$

where a is the rate of decay of the union's wage demands, Y_o is the minimum acceptable wage increase after a strike of zero length, Y_* is the minimum acceptable wage increase after a strike of infinite length, and s is the duration of the strike.

In making its initial wage offer, the firm must maximize its net present value given by

$$V = \int_s^\infty [PQ - LW(1+Y_p) - rK]\, e^{-rt}\, dt - \int_o^\infty He^{-rt}dt \qquad (2.2)$$

after accepting a strike of duration s, where H is fixed costs, and W is the previous wage rate. The maximization gives the result that they will make their offer according to

$$Y_p = \frac{r}{a+r}\left[\frac{PQ - LW - rK + H}{LW}\right] + \left(\frac{a}{a+r}\right)Y_* \qquad (2.3)$$

and accept a strike of length

$$s = -\frac{1}{a}\, ln\left\{\frac{PQ - LW(1+Y_*) - rK}{W(1+\frac{a}{r})(Y_o - Y_*)}\right\}.$$

The pre-strike behavior of the union may be analyzed in a symmetric fashion. The union is presumed to believe that the firm has the following wage increase concession schedule

$$X_A = X_* - (X_* - X_o)e^{-bs} \qquad (2.4)$$

where b is the rate of increase of the firm's offer made in order to bring the strike to an end, X_* is the maximum acceptable wage increase after a strike of ∞ duration, and X_o is the maximum wage increase that the firm is believed willing to accept without a strike. The union wishes to maximize the net present gains for its membership as a result of a strike. A possible specification for the union's financial gain is

$$G = \int_s^\infty LX_p(1+\eta)e^{-r}\, dt - \int_o^s LWe^{-rt}dt \qquad (2.5)$$

where η (< 0) is the elasticity of demand for labor and X_p is the union's proposed wage increase.

The union's maximization scheme gives the result that they will make their wage demand according to

$$X_p = -\frac{r}{r+b}\left(\frac{w}{1+\eta}\right) + \left(\frac{b}{r+b}\right)X_* \qquad (2.6)$$

and be willing to accept a strike of length

$$s = -\frac{1}{b}\, ln\left\{\frac{W + X_*(1+\eta)}{(1+\frac{b}{r})(X_* - X_o)(1+\eta)}\right\}. \qquad (2.7)$$

From (2.3) and (2.6) one can state that a strike will occur whenever $X_p > Y_p$. If one argues that (2.3) and (2.6) are incompletely specified and, consequently, have an additive random error with zero mean, then the probability of a strike is given by

$$P(\text{strike}) = P(X_p > Y_p)\,. \qquad (2.8)$$

This probability depends on labor's share of profits in the previous period, the elasticity of demand for labor, and the way in which the firm and union formulate their prior assessment of each others minimum and maximum acceptable wage increase and the rate of decay of their demands. A strike becomes more likely the lower is the previous profit level relative to the previous wage bill. As the elasticity of demand approaches one the strike becomes more likely, and less likely thereafter. The probability of a strike increases with either $X*$ or $Y*$.

3. Empirical Analysis

The wage variables are constructed from quarterly indices of weekly wage rates. Real wage changes are defined as

$$\Delta R_t = [(W_{t+2} - W_{t-2})/2W_t] - [(P_{t+2} - P_{t-2})/2P_t)] \quad (3.1)$$

for both the aggregate variable and the industry specific variable.

In our specification of the model, the aggregate wage variable is contemporaneous while the industry variable is lagged one period. A number of alternate specifications were tried, but this one seemed to give the best fit in terms of Akaike's Information Criterion [1974]. In support of this specification one might argue that in making the decision to strike, labor formulates expectations on the basis of past own-industry wage gains and gains currently made by all other laborers. In any case, the own-industry wage coefficient is expected to be negative. If the union has done poorly in the past they are likely to be more militant and inclined to strike. The aggregate wage coefficient is expected to be positive. In order to remain competitive in the marketplace, union members will be more inclined to strike if all other laborers successfully negotiate big wage gains.

Like Pencavel, aggregate gross trading profits as a percent of wage and salary compensation is used as a measure of the ability of the firm to pay and as an indication of the robustness of the economy. Ideally, one would like to have industry specific data for this variable, but lack of availability forces one to use the aggregate data. The profits variable is lagged one period to reflect the time delay in the transmission of information about financial performance from the board room to the shop floor.

The effect of profits on the incidence of strikes has conflicting explanations depending on the chosen perspective. During periods of prosperity labor leaders will try to improve the financial position of the union and its members. Also, tight labor markets prevent the substitution of non-union labor for union labor. Both factors contribute to an increased strike frequency threat. This interpretation is consistent with the model presented above. On the other hand, if product markets are also tight the firm will not anticipate losing sales during a strike so may take a hardened stance on wage increases. However, if foregone profits are calculated to outweigh the cost of higher wages, acquiescence on wage increases is more likely. Which outlook predominates cannot be determined a priori.

Unemployment rates are available by industry and in the aggregate. These rates reflect employment opportunities. Consequently, the unemployment coefficients are expected to be negative. The use of the unemployment rate in this context is problematic. The first objection is that it completely ignores the internal labor market; but no readily available data can overcome this. A second objection is that for the second half of the period studied here, 1969-1974, the Phillips tradeoff is not applicable. However, the results of Table 1 suggest our interpretation of the unemployment variable is correct.

Unlike Pencavel, the use of seasonal dummies and a time trend is eschewed here.[1] Dummy variables indicating a Labour Party government

[1] The model specified below will include two dummy variables for the party in power and wage-price freeze. Pencavel argues that the party dummy variable is necessary to account for links between the unions and the Labour Party. The incomes policy variable is important in its effect on Y_o, the minimum acceptable wage increase after a strike of zero length. See Section 2.

The inclusion of three additional dummy variables for seasonality would result in high collinearity with the intercept. Additionally, dummy variables "sop up" so much of the variation in the dependent variable that their interpretation is suspect. The use of a time trend is subject to the same criticism.

TABLE 1

SUR Corrected for Multicollinearity and Autocorrelation

Industry	L_t	F_t	$U_{.t}$	$\pi_{.t}$	$R_{.t}$	U_{it}	R_{it}	Constant	R^2	RSS	Period/n	System R^2
COAL	−80.357	−22.083	−23.136	4.271	−38.718	−2.864	−6.305	−84.568	.52	3.6×10^5	1960.2	.675
	(−4.27)	(−.67)	(−1.32)	(4.11)	(−4.34)	(−.13)	(−1.86)				−1976.2	
MAN	25.353	−23.46	−37.411	1.84	12.688	−1987.5	8.255	12.502	.15	1.2×10^5	66	
	(1.64)	(−.85)	(−2.02)	(1.19)	(1.41)	(−.32)	(1.19)					
CON	1.259	.272	−6.978	−.021	−1.827	−7.37	−3.156	49.95	.10	1.2×10^4		
	(.36)	(.04)	(−2.2)	(−.10)	(−1.14)	(−2.17)	(−3.08)					
COAL	−224.38	−53.610	−84.051	2.429	−41.095	−47.903	35.56	165.48	.78	1.84×10^5	1960.2	.904
	(−9.43)	(−1.24)	(−3.48)	(1.98)	(−2.79)	(−.85)	(2.29)				−1969.2	
MAN	50.397	−58.254	18.871	−.814	−11.68	−3307.5	19.039	138.36	.43	5.9×10^4	38	
	(3.92)	(−2.09)	(.99)	(−.766)	(−.84)	(−.69)	(1.77)					
CON	1.668	−7.319	−7.208	−.277	−6.887	−1.657	−1.655	70.7	.09	6.5×10^3		
	(.38)	(−.86)	(−1.22)	(−.87)	(−.71)	(−.19)	(−.53)					
COAL	5.648	32.635	3.41	.614	−9.26	1.265	−1.580	11.95	.40	8.58×10^3	1969.3	.780
	(.789)	(1.92)	(.69)	(1.73)	(−.28)	(.42)	(−1.68)				−1976.2	
MAN	57.474	−12.127	−93.435	6.578	9.975	−1149.30	−4.786	−109.20	.46	1.07×10^6	28	
	(2.16)	(−.24)	(−3.93)	(3.20)	(.88)	(−.39)	(−.63)					
CON	2.202	9.532	−8.516	.128	.379	−8.046	−3.328	40.965	.15	5.4×10^3		
	(.38)	(.69)	(−1.79)	(.33)	(.145)	(−1.84)	(−3.06)					

and a wage-price freeze are included.[2] The presence of an incomes policy will act to reduce labor's acceptable reservation wage increase, thus reducing the likelihood of a strike. But at the same time business will become intransigent because it cannot pass on the wage increase. The expected sign on the incomes policy variable is thus uncertain. At least nominally, the unions would be expected to support a Labour government and the party dummy should be negative.

In summary, the linear equation used to model strike frequency is given by

$$S_{it} = \beta_0 + \beta_1 R_{\bullet t} + \beta_2 R_{it} + \beta_3 \pi_{\bullet t} + \beta_4 U_{\bullet t} + \beta_5 U_{it} + \beta_6 L_t + \beta_7 F_t + \epsilon_{it} \quad (3.2)$$

where S_{it} is the number of strikes per quarter in the i^{th} industry, $R_{\bullet t}$ is the change in weekly wage rates for all industry, R_{it} is the change in wage rates for the i^{th} industry $\pi_{\bullet t}$ is the gross trading profits as a percent of wages and salary compensation, $U_{\bullet t}$ and U_{it} are aggregate and industry unemployment rates, and L_t and F_t are the Labour government and incomes policy dummy variables.

4. Estimation of Model Parameters

First round estimates of the model parameters were made using Zellner's Seemingly Unrelated Regressions (SUR) technique, without correction for either multicollinearity or autocorrelation, for the coal mining (COAL), metals manufacture and engineering (MAN), and construction (CON) industries.

The SUR technique is an appropriate estimator because it would be incorrect to assume zero contemporaneous covariances. Random shocks to the economy will affect all three industries; bargaining and strike behavior in the industries do not take place in isolation and misspecification will manifest itself in nonzero covariances.

The first set of results was for the entire sample; first quarter, 1960 through fourth quarter, 1976. The results were quite encouraging in terms of expected signs and significance. The coefficient of determination for the system was .835.

As was stated previously, simple scattergrams of the data reveal that in mid-1969 the overall trend for each of the variables alters considerably.[3] Accordingly the sample was broken into two parts and the model parameters reestimated. The first subsample, for which the system R^2 was .935, showed considerable differences in the impact of the real wage change variables on strike frequency. In four out of six cases the coefficients had the wrong sign and were significant. The second part of the sample, for which the R2 was .892, yields equally disappointing results; although most of the coefficients have expected signs, few of them are significant at acceptable levels.[4]

The fact that the quality of the results differed so markedly from one part of the sample to another was indicative of several statistical problems. Attempts have been made to deal with both multicollinearity and autocorrelation.

Multicollinearity is a problem because in its presence one is unable to sort out the independent effects of the explanatory variables, i.e., on the basis of the t-tests too few null hypothesis are rejected. Contrarywise, autocorrelation causes one to reject too many null hypotheses.

The collinear variables were selected on the basis of zero order correlations. The first step was to reduce collinearity between the aggregate variables, $U_{\bullet t}$, $\pi_{\bullet t}$ and $R_{\bullet t}$, by regressing one on the other two. The aggregate unemployment rate was selected as the dependent variable, the others as independent variables. In each case, the F-statistic was significant. In all ensuing regressions, the actual aggregate un-

[2] The Labour Party dummy takes the value of one for the periods 1964.3-1967.2 and 1974.2-1976.2. The wage-price freeze dummy takes the value of one during the periods 1961-3-1961.4, 1966.3-1966.4 and 1972.4-1973.1. These periods are cited in "Faith, Five Hopes and Cassandra," *The Economist,* July 23, 1977. p. 75.

[3] A further symptom of economic woes of Great Britain was the currency devaluation of 1969.

[4] These regression results are available from the author.

employment rate was replaced by the difference between the actual unemployment rate and its estimate. The residual unemployment is the variation in the unemployment rate not explained by profits and real wage changes. For convenience, $U_{\bullet t}$ will be used to denote the unemployment rate variable.

The next step in the process was to regress each of the industry specific variables on $R_{\bullet t}$, $\pi_{\bullet t}$ and the new unemployment variable, for each of the three time periods. A new set of industry specific variables was constructed as the difference between the observed value and the predicted value. The corrections were not made for real wage changes in construction or mining unemployment for the period 1963.3-1976.4 because of the insignificant F-statistics for these equations. Again, the original notation is maintained for convenience.[5]

After correcting for multicollinearity, the model parameters were reestimated for each industry and each period using ordinary least squares (LS) in order to determine an appropriate set of autocorrelation coefficients. The values for ρ were determined using a Cochran-Orcutt iterative procedure. The model variables were then adjusted using the calculated values for ρ.

Once all the corrections had been made to the data, the model parameters were estimated using seemingly unrelated regression. The results are presented in Table 1. Note that if one were to use the estimates of Table 1 for purposes of forecasting it would be necessary to work through the reconstructed variables.

Considering the regression results by variable, the presence of a labor government seems to reduce the number of strikes in the coal mining industry, increase the number in construction and have no effect in manufacturing. This is probably due to the fact that the miners have historically been more loyal to the Labour Party than have other unions.

[5]The method used here is based on a Gram-Schmidt orthogonalization. By not completing the process of orthogonalization we preserve the economic interpretation of the variables. The statistical results for all transformational regressions are available from the author in an unpublished appendix.

The existence of an incomes policy does not make any difference in any of the three periods for any of the industries. An incomes policy in Great Britain has traditionally been more in the nature of guidelines.

The majority of the unemployment coefficients, though not always significant, are negative. This is particularly important in view of the objections to the use of unemployment mentioned earlier.

When significant, the profits variable is positive. This is consistent with the hypothesis proposed in Section 2. The fact that it is not uniformly positive is probably attributable to the alternative relationship between profits and strikes presented above.

The model presented in Section 2 suggested quite clearly that the own wage effect should be negative. The results of Table 1 are quite discouraging in this regard. The instability of the sign of this variable is indicative of the structural changes that have taken place in Great Britain. The same instability is present in the aggregate wage coefficient.

The unexpected result for the wage change variable may be attributed to the notion of expectations. If the union won large increases in previous disputes then they are likely to use similar tactics again. This interpretation has been suggested elsewhere [Bean and Peel] but requires a simultaneous equations specification for testing.

With the exception of the construction industry for the period 1960.2-1969.2, the overall analysis reveals that correcting for multicollinearity and autocorrelation does not reduce the number of significant coefficients as one might have expected, although some coefficients become significant and vice versa.

5. Conclusions

The empirical findings of previous researchers are here confirmed. Although the previous work in the area of strike activity may be criticized on several statistical shortcomings, the inferences from those essays are quite robust. In the work presented here, coal mining, construction, and manufacturing are treated as heterogeneous and interdependent industries.

A model of strike frequency was fitted to quarterly data for the period 1960.1 through 1976.4. The data on strike frequency, unemployment rates, and real wage changes were corrected for multicollinearity and autocorrelation. Additionally, the sample was divided into two subperiods. The dividing point was mid-1969.

When the sample is considered in its entirety, the significant coefficients appear with the expected sign. However, the three industries exhibit markedly different behavior before and after the 1969 demarcation point. Furtheremore, in considering either the entire sample or the subperiods, the three industries do not exhibit consistent responses to the same economic variables. These dissimilarities are probably best explored in future research on some of the institutional aspects of British labor relations.

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