A Longitudinal Analysis of Strike Activity in U.S. Manufacturing: 1957–1984

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This study examines the determinants of both strike incidence and duration using longitudinal data on 2,767 collective bargaining settlements reached between 1957 and 1984. Two major findings are that: strike incidence is procyclical and is positively related to uncompensated unexpected inflation over the previous contract. Strike duration is found to be countercyclical.

The purpose of this study is to examine the empirical determinants of strike activity using a large microeconomic data base. Previous microeconomic studies of U.S. strike behavior by Henry Farber (1978) and Martin Mauro (1982) used considerably smaller data sets (less than one-tenth as many contracts), while more recent longitudinal studies such as Joseph Tracy (1986) and Cynthia Gramm (1986) use data bases that cover a much shorter time period and are about one-half as large. Few recent longitudinal studies focus on the effect of macroeconomic variables on U.S. strike behavior, and none examine the role of inflation.

The data base used in this study, a longitudinal file of 2,767 collective bargaining settlements reached between 1957 and 1984, is sufficiently rich to allow the testing of a variety of hypotheses concerning the influence of inflation as well as the more standard hypothesis concerning the effect of unemployment on the likelihood of strikes. The main findings of this study are that *i*) strike incidence is procyclical, *ii*) strikes are more likely the greater is uncompensated unexpected inflation over the previous contract and *iii*) strike incidence is negatively related to relative wage growth over the previous contract. Strike duration appears to be countercyclical.

The first section of the paper develops the empirical model of industrial strike incidence and duration. The empirical results on strike incidence are given in Section II, while Section III describes the empirical results for strike duration. The final section contains concluding remarks. A description of the data base is presented in the Appendix.

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I. The Model Specification

This study examines the determinants of both strike incidence and strike duration. The major focus is on macroeconomic factors that affect the bargaining environment. In addition, several contract-specific factors are included. The data set used here is well suited to this purpose because it covers a longer time period than most other micro-level studies and thus allows for considerable variation in the macroeconomic environment.

The empirical model of strike incidence that is estimated in this paper includes the following explanatory variables: the inverse of the unemployment rate for prime-aged males; expected inflation at the time the contract is signed; unexpected inflation over the previous contract for which workers have not been compensated by a cost-of-living adjustment; the duration of the contract being negotiated; the change in relative wages over the previous contract; the change in real wages over the previous contract; and industry profits.

The inverse of the unemployment rate for prime-aged males (25–54) is used as an indicator of the state of the aggregate economy. It is measured in the quarter in which the strike began or, if there was no strike, in the quarter in which the contract was effective. The prime-aged male unemployment rate was used to avoid problems caused by demographic effects on the measurement of the aggregate unemployment rate. Earlier time-series studies found that strike frequency is procyclical.[1] The data base used here covers a longer time period than other micro-level studies and thus provides sufficient variation in the unemployment rate to test whether strike incidence is procyclical.

Several inflation variables are included. Expected inflation for the current contract is measured by the 12-month expected rate of price change based on the Livingston Index for the period in which the contract was effective. Since greater expected inflation may increase both the union's wage demands (as workers attempt to protect their real wages over the prospective contract), and the firm's expected profit (so that it is more willing to grant the workers' wage demands), the effect of expected inflation on strike incidence is hypothesized to be minimal.[2] A time-series study by Bruce Kaufman (1981) used a measure of expected inflation based on the Livingston index and did not find a significant effect.

Unexpected inflation is measured by the difference between the percent change in the CPI over the last contract and the expected inflation measured at the start of the previous contract.[3] The measure of uncompensated unexpected inflation is intended to capture the unexpected inflation for which workers have not been compensated through a COLA. For unescalated contracts, uncompensated inflation is equal to unexpected inflation, while for escalated contracts, it is equal to unexpected inflation times one minus the yield of the escalator clause. The escalator yield for each contract was calculated by dividing the percentage wage change due to the COLA in the previous contract by the percentage change in the CPI over the same period. Uncompensated unexpected inflation over the previous contract leads to demands for catch-up wage increases. Insofar as firms are less willing to accede to these demands, uncompensated unexpected inflation should be positively related to strike incidence.[4]

Higher relative wage growth over the previous contract implies that the union has improved its relative position in the wage distribution and therefore is unlikely to demand unusually high wage increases. A study by Robert Swidinsky and John Vanderkamp (1982) using Canadian data found a negative, but insignificant, effect for relative wage change over the previous contract. Real wage growth over the previous contract also implies that union demands for catch-up wage increases are likely to be lower. Orley Ashenfelter and George Johnson (1969) found a negative effect of real wage growth on strike frequency, but since they were using aggregate time-series data they were unable to test for the effect of relative wage change. Morley Gunderson, John Kervin, and Frank Reid (1986) also found a negative effect for real wage change (significant only

[1]See, for example, Orley Ashenfelter and George Johnson (1969).

[2]The hypotheses advanced in the text are aimed at identifying factors that influence the likelihood that rational agents will strike or accept a strike. This is consistent with theoretical work on strikes starting with the Ashenfelter and Johnson model (1969), which explains strikes as the result of rational behavior on the part of a firm facing a known union concession function. In their model, factors that raise the union's wage demands or lower the firm's reservation wage change increase the likelihood of a strike. Recent models, such as Drew Fudenberg, David Levine, and Paul Ruud (1983), Beth Hayes (1984), and Joseph Tracy (1987) view strikes as the result of the union's incomplete information. For a survey of theoretical strike models, see John Kennan (1986).

[3]Data on price and wage changes for the previous contract were reported as annualized percentage changes so they were multiplied by the number of years in the previous contract. In the case of expected inflation, this implicitly assumes that the expected inflation in each year of the previous contract is equal to expected inflation in the first year.

[4]Support for this argument can be found in work on wage behavior. Using the same data base as this study, Wayne Vroman and John Abowd (1988) find that union wages are far less sensitive to unexpected inflation over the previous contract than they are to expected inflation at the beginning of the contract.

at the 0.10 level) using longitudinal Canadian data.

Since relative wage growth and real wage growth are highly collinear variables in the data set used here, they are introduced separately as alternative explanatory variables. The relative wage change is measured as the percent change in wages over the previous contract minus the percent change in average hourly earnings in manufacturing, while the real wage change is measured as the percent change in wages over the previous contract minus the percent change in the CPI. (Note that the first settlement for each bargaining pair was excluded from the data set so that the data for the lagged contract are complete). Both variables are expected to be inversely related to strike incidence.

The duration of the contract being negotiated (measured in months) is included to capture the fact that the union and firm have more at stake when the contract is longer. Thus, a strike may be more likely since both sides will be reluctant to concede.

Industry profits, insofar as they reflect the firm's situation, have a direct negative effect on strike incidence since the cost of a strike is greater for the firm. On the other hand, they may be positively related to strike incidence through a positive effect on union wage demands since unions appear to view profits as an indication of ability to pay. In industries with industrywide or pattern bargaining, industry profits may in fact be more relevant than firm profits.

In addition to estimating the model described above for strike incidence, the determinants of strike duration are examined. Strike duration is expected to depend on the same factors as strike incidence since factors which raise the likelihood of a strike in a given negotiation are also likely to make it more difficult to settle and so lead to a longer duration. One exception to this is the business cycle effect. Recent empirical work on strike duration has found evidence that strike duration is countercyclical.[5] Thus the effect of unemployment on strike duration may be opposite to its effect on strike incidence.

Before discussing the empirical results, it is useful to briefly describe the data base being used. The data base is a longitudinal file of major collective bargaining agreements in 252 bargaining situations reached between 1957 and 1984. The settlements are exclusively in the manufacturing sector. Of these, 331 settlements involved strikes. Only strikes related to contract negotiations are considered. More information is provided in the Appendix.[6]

II. Empirical Results—Strike Incidence

Columns (1) through (4) of Table 1 present regression results from the OLS estimation of strike incidence equations. These equations were also estimated using maximum likelihood Probit estimation, a more appropriate technique given the binary nature of the dependent variable.[7] The Probit estimates by accounting for the bivariate nature of the dependent variable are more efficient. These results are presented in columns (5) through (8). Note that the coefficient estimates from the Probit analysis are not directly comparable to those from the OLS

[5]See Kennan (1985).

[6]For additional information on this data set see W. Vroman (1986). The original data set had 304 strikes, 27 strikes occurring after 1970 were added based on information provided by David Card.

[7]The Probit analysis is based on the following reformulation of the model:

$$Z_i = \beta' z_i + u_i$$

$$S_i = 0 \quad \text{if } Z_i \leq 0,$$

$$S_i = 1 \quad \text{if } Z_i > 0,$$

where $S_i = 1$ if a strike occurred and 0 otherwise and z_i is the vector of explanatory variables. If $u_i \sim IN(0,1)$, the likelihood function to be maximized is:

$$L = \prod_{S_i = 1} F(\beta' z_i) \prod_{S_i = 0} [1 - F(\beta' z_i)],$$

where $F(\ )$ represents the standard normal cdf. See G. S. Maddala (1983).

TABLE 1—STRIKE INCIDENCE[a]

	OLS Estimation				Probit Estimation-Marginal Effects			
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Constant	−0.050	−0.139	−0.137	−0.140	−0.417	−0.545	−0.543	−0.545
	(0.89)	(2.37)	(2.33)	(2.38)	(6.23)	(7.45)	(7.46)	(7.46)
Inverse of U Rate	0.328	0.284	0.295	0.293	0.334	0.298	0.311	0.301
(prime-aged males)	(6.96)	(5.62)	(5.83)	(5.73)	(6.79)	(5.51)	(5.79)	(5.50)
Contract Duration	–	0.003	0.003	0.003	–	0.004	0.004	0.004
		(4.71)	(4.67)	(4.42)		(4.72)	(4.60)	(4.63)
Expected Inflation	–	0.423	0.195	0.288	–	0.423	0.183	0.381
		(1.53)	(0.74)	(0.96)		(1.28)	(0.59)	(1.09)
Relative Wage Change	–	−0.378	–	−0.356	–	−0.311	–	−0.302
(Previous Contract)		(2.87)		(2.68)		(2.11)		(2.03)
Real Wage Change	–	–	−0.089	–	–	–	−0.012	–
(Previous Contract)			(0.70)				(0.02)	
Uncompensated Inflation	–	0.351	0.503	0.514	–	0.528	0.741	0.578
(Previous Contract)		(2.06)	(2.81)	(2.32)		(2.53)	(3.32)	(2.34)
Uncompensated Inflation	–	–	–	−0.041	–	–	–	0.328
(Negative)				(0.11)				(0.58)
Exclusion Tests[b]								
Industry Dummies	8.11	7.45	7.23	7.52	148.0	133.0	130.5	132.5
Month Dummies	1.57	1.61	1.58	1.59	18.9	20.1	19.6	20.1
$\bar{R}^2$	0.0666	0.0861	0.0835	0.0862				
SEE	0.314	0.310	0.311	0.310				
Log-Likelihood					−904.99	−872.71	−874.97	−872.64
χ^2					216.43	281.00	276.48	281.14
$N = 2767$								

[a]Absolute values of the t-statistics in parentheses.
[b]For the first four columns, these are F-statistics, while for the last four columns they are χ^2 statistics for the likelihood ratio tests.

estimation because they represent the effect of the independent variable on $F^{-1}(P_s)$, where F is the cumulative density function (cdf) for the normal distribution. The partial derivatives of the strike probability with respect to the independent variables depend on the level of the probability, that is, the steepness of the cdf. The marginal effects reported in the table give these partials evaluated at the average strike probability.[8] To test the null hypothesis: H_0: $\beta_2 = \beta_3 = \cdots = \beta_k = 0$, that is, that all the coefficients equal zero, likelihood ratio tests were performed. For all the equations, the hypothesis is rejected at the 0.01 level.

The results reported in Table 1 are for strike equations that include dummy variables for the two-digit manufacturing industries, SIC 20 to SIC 38 (excluding miscellaneous manufacturing)[9] as well as dummies representing the month in which the strike occurred or, if there was no strike, the month in which the contract was effective (excluding October). The results in columns (1) and (5) are intended to test whether strike incidence is procyclical. Columns (2) and (6) correspond to the model specified in Section I. Note that profits are omitted because they were consistently insignificant in all estima-

[8]These partials are $\partial P_{si}/\partial z_{ij} = f(\beta' z_i)\beta_j$, where $f(\ )$ is the normal pdf. The sample mean probability is 0.1196 and the value of $F^{-1}(\bar{P}_s)$ at the mean is -1.177. This figure was then used to find the value of $f(\ )$, which is 0.1996.

[9]The equations in columns (1)–(4) were reestimated using 251 dummy variables to test for bargaining pair fixed effects. These did not significantly alter the results for the other variables and they were reasonably consistent with the results for the industry dummies. The bargaining pair fixed effects were not used in the Probit estimation as this was infeasible.

tions. The remaining columns contain variations of this basic model.

The results for all specifications (using both OLS and Probit) support the hypothesis that strike incidence is positively related to the tightness of the labor market as measured by the inverse of the unemployment rate for prime-aged males. This variable enters positively and is highly significant in all regressions.

The duration of the contract being negotiated has the expected positive sign and is highly significant in all specifications. Thus, there is support for the hypothesis that longer contracts raise the likelihood of a strike.

The expected inflation variable is insignificant across all equations. Thus, there is no support for the hypothesis that higher expected inflation raises the likelihood of strikes. As noted above, it is possible that while expected inflation raises workers' wage demands, it also raises the firm's expected profits so that it has no influence on strike incidence.

Relative wage increases over the previous contract were hypothesized to lower the strike probability. The relative wage change variable has the correct sign and is significant at the 0.05 level or better in all specifications. In columns (3) and (7), relative wage change is replaced with the real wage change. The real wage change has the correct sign, but is insignificant. The *t*-statistic is close to zero in the Probit estimation. Note that when the real wage change is included, the effect of uncompensated unexpected inflation is increased. This latter variable could be interpreted as the uncompensated unexpected real wage loss over the previous contract and is negatively correlated with the real wage change variable. The results indicate that strikes (or disagreements over wages) are negatively related to relative wage growth and positively related to unexpected (uncompensated) real wage loss.

As noted above, the measure of uncompensated unexpected past inflation has a positive and significant coefficient, yielding evidence that demands for wage increases to catch-up with unexpected past inflation raise the probability of a strike. In columns (4) and (8), this variable is separated into a positive and negative component to test whether uncompensated inflation has a different effect when it is positive than when it is negative. The results indicate that positive uncompensated inflation has a significant positive coefficient while negative uncompensated inflation has an insignificant coefficient. This supports the hypothesis that uncompensated inflation increases the incidence of strikes because it leads to workers' demands for catch-up wage increases, which firms do not acknowledge. When workers are more than compensated for unexpected inflation, this issue does not seem to affect strike incidence.[10]

Nineteen industry dummy variables were used to capture industry fixed effects—one for each of the two-digit manufacturing industries except miscellaneous manufacturing. *F*-statistics for the test of the significance of the entire set of dummies are reported in columns (1) through (4), while the χ^2 statistics for likelihood ratio tests are given in columns (5) through (8). For all the equations, the hypothesis that the entire set of dummies makes no significant contribution to the explanation of strike incidence is rejected at the 0.01 level. Only five of the dummies, however, have coefficients that are significant at the 0.05 level (in a two-tailed *t*-test) across all equations. SIC 29, Petroleum and Coal Products, SIC 30, Rubber and Plastic Products, SIC 33, Primary Metals, SIC 34, Fabricated Metals, and SIC 35, Nonelectrical Machinery, have significant positive coefficients indicating that relative to miscellaneous manufacturing strikes are more likely in theseindustries. The results accord with the usual presumptions of behavior in these industries. For the equations reported in all columns but (6) and (7), the dummy for SIC 37, Transportation Equipment, was also positive and significant. The coefficients and *t*-statistics for the industry dummies for the specification in column (6) are given in Appendix Table A3. Table A3

[10] Uncompensated inflation is positive in 2,399 observations and negative in 368 observations.

TABLE 2—SEASONAL EFFECTS[a]

	OLS Estimation (2)	Probit Marginal Effects (6)
Month Dummies		
January	−0.051 (1.50)	−0.062 (1.64)
February	0.018 (0.57)	0.011 (0.33)
March	0.020 (0.62)	0.017 (0.52)
April	−0.032 (1.00)	−0.060 (1.55)
May	−0.033 (1.21)	−0.063 (2.00)
June	−0.019 (0.66)	−0.026 (0.84)
July	0.013 (0.47)	0.005 (0.16)
August	−0.034 (1.19)	−0.041 (1.31)
September	0.020 (0.64)	0.015 (0.45)
October	–	–
November	0.038 (1.28)	0.030 (1.01)
December	0.005 (0.14)	−0.004 (0.09)
Quarter Dummies		
$Q1$	−0.018 (0.92)	−0.019 (0.90)
$Q2$	−0.044 (2.49)	−0.059 (2.91)
$Q3$	−0.017 (0.95)	−0.019 (0.99)

[a]Absolute values of the t-statistics in parentheses.

provides the coefficient estimates for the industry fixed effects for two of the estimations reported in the text.

Monthly dummies were included in the analysis to determine whether there was a seasonal pattern to strikes. F-tests (columns (1) to (4)) and likelihood ratio tests (columns (5) to (8)) performed to test the significance of these dummies indicate that the null hypothesis of no effect cannot be rejected at the 0.05 level of significance except in columns (6) and (8). (It can, however, be rejected at the 0.10 level in all columns). Consistent with this, none of these dummies has a coefficient that is significant at the 0.05 level in a two-tailed test for the OLS estimation. The dummy for May has a significant negative coefficient in columns (6) through (8). Table 2 presents the results for the month dummies for columns (2) and (6). Since there did appear to be a seasonal pattern in the coefficients—those for April, May, and June were negative across all equations, the equations were reestimated using quarterly dummies (excluding the fourth quarter). The results for the quarterly dummies for the equation specification in columns (2) and (6) are also reported in Table 2. The dummy for the second quarter is negative and significant. (This is also true for the other equations). Likelihood ratio tests for the entire set of quarterly dummies reject the null hypothesis of no seasonal effect at the 0.05 level for the specifications in columns (6) through (8).[11] Thus, there does appear to be some seasonal pattern to strikes in this data set.

Recent studies using shorter time periods have also found significant seasonal effects. Gramm (1986) finds a significant negative coefficient for the third-quarter dummy and a positive (though insignificant) coefficient for the second-quarter dummy. The different results are likely due to the different data set and the shorter time period (1971–1980). David Card (1987) using data based on the same underlying data set as that used here finds a somewhat different pattern for the month dummies. He also finds that the whole set of month dummies is a significant factor in explaining strike incidence. Card's sample period, however, is quite different. His data go only through 1979 and he uses only the six most recent settlements for each bargaining pair, so that he has fewer settlements in the early part of the data period.

Comparison with Other Recent Studies. Most of the recent longitudinal studies of U.S. strike behavior do not focus on aggregate unemployment or inflation. One exception is the paper by Sheena McConnell (1987). Her data cover a shorter time period, 1970 to 1981, but she does find that strike incidence is procyclical. Tracy (1986) finds that above average local employment residuals are associated with a higher strike incidence, but that above average industry employment residuals are inversely related to strike incidence. Gramm (1986) finds no effect of local unemployment, but a positive effect of in-

[11]The χ^2 statistics for likelihood ratio tests for the exclusion of the quarterly dummies for columns (5) through (8) are 6.48, 9.28, 9.68, and 9.26, respectively. The critical value of χ^2 with three degrees of freedom at the 0.05 level is 7.82.

creases in product market demand. None of these studies examines the effect of unexpected inflation on strikes. An earlier time-series study by Kaufman (1981) examined several hypotheses similar to those tested here. He found that strike activity was negatively related to the unemployment rate for adult males, and that corporate profits had no effect on strikes. He also found that expected inflation as measured by the Livingston index did not have a significant effect on strike activity, while inflation over the previous contract period had a significant positive effect, which was reduced by the presence of escalator clauses.

Recent studies using Canadian contract data have shown more interest in the effect of macroeconomic variables on strike incidence. Swidinsky and Vanderkamp (1982), using a large microeconomic data base covering Canadian union settlements over the period 1967 to 1975, found that labor market tightness was a significant factor in explaining the propensity to strike. In their estimation, the relative wage change over the previous contract entered negatively (but was insignificant). A more recent study by Jean-Michel Cousineau and Robert Lacroix (1986) also uses longitudinal data on Canadian strikes (1967–1982) and estimates strike incidence equations using Probit analysis. (Swidinsky and Vanderkamp use OLS). Their model is quite different from the one estimated here, but there are several consistent findings—a significant positive effect of previous inflation and of lagged contract duration. (Note that Card (1987) using U.S. data also finds a significant effect of lagged contract duration). While lagged contract duration was not included here, both uncompensated unexpected inflation and the relative wage change are measured over the previous contract, and thus incorporate the length of the previous contract. Their results for lagged contract duration may reflect the fact that workers lose more in relative or real wages over longer contracts. A third Canadian study by Gunderson, Kervin, and Reid (1986) covering the period 1971 to 1983 also finds that strike incidence is procyclical and finds a negative effect of the change of real wages over the previous contract.

TABLE 3—STRIKE DURATION—OLS ESTIMATION[a]

Variable	(1)	(2)
Constant	3.458	2.539
	(5.09)	(3.42)
Inverse of U Rate	−1.211	−0.963
(Prime-Aged Males)	(2.69)	(1.96)
Expected Inflation	–	8.619
		(2.82)
Contract Duration	–	0.016
		(1.83)
Relative Wage Change	–	−2.112
(Previous Contract)		(1.45)
Uncompensated Inflation	–	−3.766
(Previous Contract)		(1.90)
$\bar{R}^2$	0.0952	0.1192
SEE	1.058	1.044
$N = 331$		

[a]Absolute values of the t-statistics in parentheses.

III. Empirical Results—Strike Duration

Table 3 presents the results from OLS estimation of strike duration equations. The dependent variable is the log of the strike duration (measured in days).[12] The first column presents results for an equation containing just the inverse of the unemployment rate and the industry and month dummies. The second column contains results for the estimation of the model described in Section I, also including the industry and month dummies. (As in Table 1, the equations estimated do not include industry profits since these are always insignificant).

In both columns, the inverse of the unemployment rate for prime-aged males has a negative and significant coefficient. This indicates that strike duration is shorter when

[12]These relationships were also estimated using hazard functions based on the Weibull distribution. The coefficient estimates are similar to those found using OLS. The Weibull distribution was used to allow for duration dependence. Evidence of positive duration dependence was found. This implies that the settlement rate increases with strike duration. Tracy (1986) uses a similar hazard function and also finds positive duration dependence. The hazard function estimates are included in a longer version of this paper and are available from the author. For more information on this hazard function, see John Kalbfleisch and Ross Prentice (1980, pp. 31 and 54–55).

the labor market is tighter (when $1/U$ is greater). The result that strike incidence is procyclical while strike duration is countercyclical may reflect the fact that tight labor markets increase the union's bargaining power and lead to greater wage demands and a slower decline in these demands over the course of a strike. In such periods, when product demand is high, the firm's strike costs are also high. The firm may be willing to take a short strike in the hope of reducing the union's demands, but is likely to settle more quickly due to its high strike costs.

Column (2) includes the other variables of the model. Among these variables, only the expected inflation variable is significant in a two-tailed test. Expected inflation is positively related to strike length. Both the duration of the contract being negotiated and the uncompensated inflation variables are significant in one-tailed tests. The contract duration has a positive coefficient as it did in the strike incidence equations, while uncompensated inflation has a negative coefficient in the strike duration estimation and a positive coefficient in the strike incidence estimation. The result for uncompensated inflation is similar to that for unemployment and may be due to the same factors. The relative wage change over the previous contract is insignificant.

None of the industry dummies or the month dummies are significant in two-tailed tests, but in both columns the dummy for the Primary Metals industry (SIC 33) has a positive coefficient that is significant at the 0.05 level in a one-tailed test. This suggests that strike durations are higher in this industry. The result reflects the long strikes in this industry in 1959 and 1960 and in 1968. The results for the industry dummies for column (2) are given in Appendix Table A3.

These results are consistent with recent studies by Kennan (1985) and Tracy (1986) that estimate hazard functions for strike duration. Using data on industrial production, Kennan finds evidence that strike duration is countercyclical. Tracy does not include a measure of aggregate economic activity, but does find that above average local employment residuals are associated with a reduction in strike duration.

IV. Concluding Remarks

This study examined the determinants of strike activity using a longitudinal contract data base. Several specific hypotheses about the factors affecting the propensity to strike were tested. The results support the view that labor market tightness (as measured by the inverse of the unemployment rate for prime-aged males) has a positive and significant effect on strike incidence. Further, strikes appear more likely the longer the contract currently being negotiated and the lower the relative wage change over the previous contract.

Special attention was addressed to the role of inflation. Expected future inflation appeared to have no effect on strike incidence, but uncompensated unexpected inflation over the previous contract had a significant positive effect. Further tests suggested that positive uncompensated unexpected inflation was responsible for this effect. It may be that both parties accept the notion that wages should be adjusted for expected inflation, but that catch-up for past inflation is a matter of disagreement between firms and unions. Further research into this issue is warranted.

Strike duration equations were also estimated. This analysis indicated that although strike incidence is greater in periods with tight labor markets, strike duration is shorter in such periods. In addition, strike duration was found to be positively related to expected inflation.

APPENDIX: Description of the Data

The data set used in this study contains 2,767 observations on major collective bargaining agreements in manufacturing reached during the period 1957 to 1984. Of these, 331 involved strikes. The data cover 252 bargaining situations. Table A1 gives the distribution of observations, situations, strikes, and strike incidence by two-digit industry. Table A2 provides information on the sample by month. The major advantages of this data set is that it is longitudinal and extends over a longer time period than other micro-level contract data bases.

The original data were collected at the Urban Institute based mainly on information reported in the U.S. Department of Labor publication, *Current Wage Developments* (CWD), and covered the period 1957 to 1979. The data base was later updated to include agreements reached between 1979 and 1984. Because of recent

TABLE A1—DISTRIBUTION OF OBSERVATIONS, SITUATIONS, STRIKES, AND STRIKE INCIDENCE BY TWO-DIGIT INDUSTRY

SIC Code	Industry	Observations	Bargaining Situations	Strikes	Strike Incidence (Percent)
Nondurable Manufacturing		1,381	114	114	8.3
20	Food	238	21	18	7.6
21	Tobacco	27	3	3	11.1
22	Textiles	206	13	9	4.4
23	Apparel	126	13	2	1.6
26	Paper	251	19	17	6.8
27	Printing and Publishing	70	8	4	5.7
28	Chemicals	266	21	28	10.5
29	Petroleum Products	62	4	11	17.7
30	Rubber and Plastic	80	7	21	26.3
31	Leather	55	5	1	1.8
Durable Manufacturing		1,386	138	217	15.7
24	Lumber	97	9	7	7.2
25	Furniture	44	5	6	13.6
32	Clay, Glass, Stone	224	22	15	6.7
33	Primary Metals	180	17	33	18.3
34	Fabricated Metals	152	15	33	21.7
35	Machinery	145	15	42	29.0
36	Electrical Equipment	226	24	34	15.0
37	Transportation Equipment	217	24	43	19.8
38	Instruments	61	4	1	1.6
39	Misc. Manufacturing	40	3	3	7.5
Total Manufacturing		2,767	252	331	12.0

declines in unionized employment, 41 of the situations in the original data were no longer major bargaining situations (1000 or more workers), and so were not reported in CWD. By contacting the parties involved in the situations, 23 of these situations were updated. These had been reduced in size, but were still bargaining. Eighteen situations could not be updated because they had either disbanded (12) or because data could not be obtained.

The situations included in the original sample were chosen to represent the larger situations in those three-digit industries that were heavily unionized. Thus, the sample represents about 70 percent of all employment covered by major manufacturing situations in 1978.

The final sample used for this study is smaller than the original sample because it excludes: a) the first settlement for each situation in order to have complete data on lagged contracts, b) contracts with a duration less than 7 months, and c) contracts, where a full set of data were unavailable. The data set used in this study is available from the author upon request. The original data base is on file at ICPSR, Institute for Social Research, University of Michigan.

TABLE A2—INFORMATION ON THE SAMPLE BY MONTH

Month	Observations	Strikes	Strike Incidence (Percent)
January	172	20	11.6
February	160	22	13.8
March	164	26	15.9
April	178	12	6.7
May	344	27	7.8
June	359	35	9.7
July	338	48	14.2
August	256	26	10.2
September	189	26	13.8
October	247	40	16.2
November	230	36	15.7
December	130	13	10.0

TABLE A3—INDUSTRY FIXED EFFECTS

Industry	Strike Incidence Probit-Marginal Effects (6)	Strike Duration OLS Estimates (2)
Food	0.033 (0.48)	0.317 (0.47)
Tobacco	0.057 (0.61)	0.382 (0.42)
Textiles	0.001 (0.01)	−0.130 (0.18)
Apparel	−0.154 (1.79)	−0.953 (0.98)
Paper	0.038 (0.56)	0.943 (1.38)
Printing	−0.002 (0.02)	0.396 (0.48)
Chemicals	0.079 (1.19)	0.813 (1.22)
Petroleum Products	0.199 (2.55)	0.361 (0.47)
Rubber	0.214 (3.01)	0.659 (0.96)
Leather	−0.138 (1.28)	−0.874 (0.70)
Lumber	0.036 (0.47)	1.205 (1.58)
Furniture	0.057 (0.71)	−0.361 (0.47)
Clay	0.005 (0.07)	0.555 (0.82)
Primary Metals	0.139 (2.06)	1.297 (1.95)
Fabricated Metals	0.155 (2.29)	0.347 (0.53)
Machinery	0.193 (2.87)	0.109 (0.17)
Electrical Equipment	0.092 (1.39)	0.472 (0.72)
Transportation Equipment	0.130 (1.95)	0.186 (0.29)
Instruments	−0.080 (0.78)	0.674 (0.53)

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