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Bargaining Power, Strike Durations, and
Wage Outcomes: An Analysis of
Strikes in the 1880s

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ABSTRACT

We study strike durations and outcomes for some 2000 disputes that occurred between 1881 and 1886. Most post-strike bargaining settlements in the 1880s fell into one of two categories: either a union "victory", characterized by a significant wage gain or hours cut, or a union "defeat", characterized by the resumption of work at the previous terms of employment. We find a strong negative relation between strike duration and the value of the settlement to workers, reflecting the declining probability of a union victory among longer strikes. For the subset of strikes over wage increases we estimate a structural model that includes equations for the capitulation times of the two parties and a specification of the wage increase conditional on a union victory. We find strong support for a relative bargaining power hypothesis: factors that enhance the workers' ability to withstand a strike tend to raise the wage increase in the event of a successful strike, while factors that enhance the employer's ability to withstand a strike tend to lower the wage increase in the event of a union victory.

Bargaining Power, Strike Duration and Wage Outcomes: An Analysis Of Strikes in the 1880s

Definitions of union bargaining power typically focus on the ability of a union to achieve higher wages, improved benefits, or better working conditions.¹ In an industrial relations system built around collective bargaining and business unionism, the fundamental source of this power is the ability to withhold labor services and inflict a costly strike on an employer. The legal framework created by the National Labor Relations Act (as amended) narrowly circumscribes the ability of a union to raise the costs of a strike, and the ability of an employer to avoid these costs. While such legislative restrictions on strike behavior may be socially desirable, they greatly limit the range of outcomes observable by a researcher, making it difficult to identify the "strike power" of a union and measure the effect of this power on negotiated outcomes.

In the labor market of the 1880s there were far fewer constraints on the actions of strikers or employers. Secondary boycotts, sympathy strikes and the use of strike replacements were all commonplace. This environment routinely produced a full range of strike outcomes, ranging from outright union victories to the permanent replacement of all strikers. Compromise settlements, which characterize most collective bargaining agreements today, were the exception rather than the rule.

This paper presents a detailed examination of the outcomes of over 2000 strikes from the period 1881-1886. We exploit the relatively clear distinction between successful and unsuccessful strikes in the labor market of the 1880s to identify the determinants of union bargaining power and study the connection between the parties' abilities to withstand a strike, on the one

¹See for example, Kochan and Katz (1988, p. 54).

hand, and the wage outcomes achieved by a successful strike, on the other.

Several features of this time period and the underlying strike data are important to our study.² The years 1881-1886 correspond to the period of growth of the Knights of Labor. For the first time, unskilled workers from a variety of industries joined the ranks of organized labor and participated in strikes. The wide range in skills and organizing histories of workers involved in strikes creates an ideal opportunity to study the sources of variation in strike outcomes. During the spring of 1886 a drive by radical and traditional labor organizations for an 8-hour working day lead to an unprecedented level of strike activity (see David (1936)). Workers in many cities participated in a "nationwide" strike on May 1 1886. In Chicago, the demonstrations culminated in a violent confrontation between police and strikers at Haymarket Square on May 3. Public reaction to this incident bolstered a new wave of employer opposition to union organizing. Comparisons of strike outcomes before, during, and after the Haymarket incident allow us to assess the effects of global events and attitudes on individual-level bargaining outcomes.

Most important, however, is the quantity and quality of data on strikes in the 1880s. As a result of the widespread "labor problems" of 1885 and 1886, the Bureau of Labor, under commissioner Carroll Wright, undertook an unprecedented data collection effort to enumerate every strike and lockout in the US between 1881 and 1886. The results of this inquiry, including detailed strike listings for over 5000 individual strikes, were published in the Third Annual Report of the Commissioner of Labor. These detailed listings provide

²Many comprehensive histories of the period are available, including Commons et al (1926) and Ware (1929).

the basis for our study. While by no means perfect, the data contained in the Third Report are in many respects better than the information available on contemporary strike activity.

Our analysis of individual strike outcomes reveals an important feature of disputes in the late 19th century. Post-strike wage settlements (and post-strike hours changes, in strikes over hours) were sharply bimodal. The Bureau of Labor's classification of strikes into union "wins" and "losses" was apparently justified by the nature of strike outcomes at the time. In strikes over a wage increase, for example, union losses were almost always associated with no change in wages. In cases of a union victory, on the other hand, the vast majority of strikes achieved a significant wage gain. From this observation we derive and test a simple "structural" model of strike durations and outcomes, based on the abilities of the union and the firm to endure a strike. We apply this model to the subset of strikes initiated over the demand for a wage increase. We specify equations for the capitulation times of the firm and the workers, and estimate the effects of various observable variables in raising or lowering these times. We then add a specification for the wage settlement in the event of a union victory. We find substantial support for a very simple "relative bargaining power" hypothesis: factors that shift the union's ability to withstand a strike relative to the firm's also raise the wage settlement in the event of a union victory.

II. Strikes in the 1880s: A Descriptive Overview

Information on strikes and lockouts for the period from January 1 1881 to December 31 1886 was assembled retrospectively by the Bureau of Labor in 1886 and 1887. Starting from a list of disputes described in newspapers and trade

magazines, the Bureau assigned field agents to track down the details of the known disputes and to gather information on other strikes or lockouts during the period. According to the Commissioner, ... "the parties instigating a strike were consulted...and the agent, after considering all the evidence to be gained on either side, reported what the facts seemed to be." (p. 10).³

By modern standards, the data collection effort underlying the preparation of the Third Report was extraordinary.⁴ Nevertheless, a recent study of the accuracy of the strike listings in the Third Report suggests that the Bureau of Labor did not achieve a complete census of disputes. Bailey (1991) compares the strikes reported in the Third Report with the set of disputes mentioned in local newspapers in Terre Haute Indiana between 1881 and 1886. His results suggest that only about one half of the disputes were actually recorded in the Third Report.⁵ This undercount poses no particular problem for our statistical analysis, provided that the Bureau enumerated a random sample of disputes. To the extent that uncounted strikes differ from the ones recorded by the Bureau, however, the available sample may present a biased picture of strike activity and outcomes in the 1880s.

The available information for each dispute includes the location and industry of the employer or employers, the number of employees affected by the dispute, the number of strikers directly involved, average wages and hours

³We have no direct evidence on how many field agents were employed by the Bureau, or in what fraction of cases the field agents managed to successfully interview the parties to a strike.

⁴An earlier study of strikes in 1880 was compiled as part of the 10th Census. The Tenth Report of the Commissioner of Labor reported similar information for strikes occurring between 1887 and 1894.

⁵Bailey actually assesses the combined accuracy of the Third Report and the later Tenth Report of the Commissioner of Labor, which reported strikes from 1887-1894.

before and after the dispute, the proximate cause of the dispute, and information on the use of strike replacements. Individual employer identities and the name of the union(s) involved in the dispute are unavailable, although an indicator is available for strikes that were ordered by a labor organization.⁶

Information is also provided on the beginning date and duration of the strike. As in modern data, the precise duration of a labor dispute is often ambiguous. The Bureau of Labor's procedures for defining the end of a dispute are not described in the Third Report. The procedures used in the Tenth Report of the Commissioner of Labor (also written under Wright's supervision) suggest that a strike was deemed to be over when the establishment was "open and operating as usual" (Tenth Report p. 15), either with the former striking employees or with strike replacements.

A second important issue in the interpretation of the data in the Third Report is the reporting of inter-related disputes. The general philosophy of the Bureau at the time of the Third Report was to treat the establishment as the unit of analysis.⁷ Multi-establishment strikes were recorded as a single dispute if they shared the same starting and ending dates, involved similar issues, and were resolved in the same manner, but were recorded as separate strikes if (for example) they ended on different dates or were resolved differently. The vast majority of the strike listings (83 percent) involve

⁶It is possible to identify many of the well-known strikes from this period: for example, the strike of telegraphers in 1883 (described in Ware (1929, p. 129-130) or the 1886 lockout of collar laundresses in Troy New York (in Filipelli (1990, pp. 542-546). The facts of the identifiable strikes presented in the Third Report are generally compatible with information from other sources.

⁷The Bureau's procedures are described in some detail on pages 11-14 of the report.

only a single establishment. On the other hand, there are many examples of consecutive listings in the Third Report that clearly refer to inter-related disputes.⁸

A measure of the degree of interdependence between the listings in the Third Report can be constructed from the strike totals presented in the Twenty-First Annual Report of the Commissioner of Labor. By the time of the Twenty-First Report the Bureau had moved to a broader definition of a single dispute: strikes that started at (roughly) the same time over similar issues were treated as a single dispute.⁹ To obtain comparable data for the 1881-1886 period, the Bureau re-analyzed the earlier strike reports. Whereas the Third Report lists 5809 strikes and lockouts from 1881-1886, the Twenty-First Report enumerates only 4211 independent disputes. A comparison of these totals suggests a 30 percent "overcount" in the Third Report. The interdependence of some of the strikes in our data set should be kept in mind in interpreting our empirical results.

From the available data in the Third Report we have chosen to analyze strikes from the states of Illinois, New York, and Massachusetts. These 3 states accounted for 41 percent of all disputes (strikes and lockouts) in the US in the early 1880s. During this time period lockouts represented only 6

⁸For example, the Third Report lists 16 strikes of bakers in Albany New York that all began on May 2 1885 (one involved 6 establishments, one involved 5 establishments, a third involved 2 establishments, and the remainder were single-establishment strikes). The strikes varied in length from 1 to 275 days in duration.

⁹See Griffen pp. 22-27 for a discussion of the changing views of the Bureau of Labor (later, the Department of Labor) on how to distinguish "independent" strikes.

percent of total disputes, and for convenience we have excluded them from our analysis.¹⁰

The Third Report lists 2256 individual strikes in Illinois, Massachusetts and New York between 1881 and 1886. Initial analysis of the data suggested a series of typographic errors affecting 77 strikes in Illinois.¹¹ Exclusion of these strikes generates a usable sample of 2179 observations. Table 1 presents some simple descriptive statistics for the sample, including breakdowns by state and year. Just over one-half of the strikes are from New York state, while 30 percent are from Illinois and 15 percent from Massachusetts. The annual number strikes is fairly stable from 1881-1885 and then shows a dramatic increase in 1886. Much of this increased strike activity grew out of the "8 hour day" campaign launched by the Federation of Trades and Labor Unions.¹² Figure 1 plots the number of strikes in each month of the sample period. The number of strikes in March and April of 1886 was over twice the average for these months in the preceding 5 years. In May 1886 there were as many strikes as in all of 1881 and 1882 combined.

Rows 2-6 of Table 1 show the average number of workers involved per strike, the average fraction of workers at each establishment involved in the strike, the average daily pre-strike wage of strikers, and the percent of strikes that are recorded as having been ordered by a labor organization (for simplicity we refer to these as "authorized" strikes). The mean size of

¹⁰A cursory examination suggests that lockouts were longer than strikes (mean duration 38 days versus 20 days) and were less likely to be sanctioned by a labor union (51 percent versus 77 percent), but were equally likely to result in a union "success" (49 percent versus 47 percent).

¹¹These strikes all involved laborers and wharf hands in the Chicago shipping industry. Details of our findings are presented in Appendix A.

¹²See David (1938, chapter 7), for example.

strikes is relatively large (245 workers), reflecting the presence of a few very large disputes. The median strike is much smaller (50 workers) -- for this reason, we also report the mean logarithm of the strike size and use the transformed variable in our statistical analysis below. On average, 80 percent of employees at the affected establishments participated in the strike.

The average wage of strikers in the 1880s was approximately \$2.00 per day although there is considerable variation across strikes, ranging from under \$0.75 to over \$4.00 per day. These rates are comparable to other wage data for the period. For example, Long's (1960) tabulations of manufacturing wages for 1880 show average daily wage rates of \$2.20 - \$2.45 for skilled occupations and \$1.32 for laborers. There is some variation across years, particularly in the average wage of strikers in 1884. The upward drift in average wages is consistent with overall wage trends during the 1880s: Lebergott's wage index (Historical Statistics of the United States, Table D-735) shows a 10 percent rise in nominal wages from 1881 to 1886.

On average three-quarters of the strikes were authorized by a union. The fraction of authorized strikes is higher in New York state, lower in Massachusetts, and shows an upward trend from 1881-1886, paralleling the steady growth of labor organizations during the early 1880s (Wolman (1936)).

Rows 7-9 of the Table report the fractions of strikes attributable to 3 major causes: disputes over workers' demands for a wage increase, disputes over the employer's demand for a wage cut, and disputes over workers' demands for a reduction in hours. Together these causes account for 77 percent of all strikes. The remainder are attributable to a variety of factors, including employee discharges, changes in work rules, and sympathy strikes. The

importance of the 8-hour day campaign in 1886 is illustrated by the unusually high fraction of hours-related strikes in that year. Relative to the pattern in earlier years, at least 300 extra hours strikes were reported in 1886.

Information on the duration of strikes is presented in rows 10 and 11. The distribution of strike durations is highly skewed, resulting in a mean duration considerably in excess of the median duration. The most frequent duration is 1 day (12 percent of all strikes). One-third of all strikes in the 1880s ended within 3 days. The means and medians suggest that strikes were relatively longer in Massachusetts than in the other two states, and were longer in 1885 than in other years.

Finally, the last five rows of Table 1 present information on strike outcomes. The authors of the Third Report (and later Bureau of Labor strike reports) coded strikes as union successes, failures, or partial successes. Close to one-half of all strikes were successful from the union's perspective, while 40 percent were failures. Perhaps surprisingly, only a small fraction of strikes (10 percent) were partially successful. Although the classification of strikes as successes, compromises, and failures may seem artificial from a modern perspective, earlier observers made extensive use of the classification (see for example Moore (1911, chapter 5)). As we show below, examination of the wage and hours changes associated with the different outcomes suggests that the distinction is relatively clear-cut.

Another useful measure of strike success is the extent to which new employees were recruited to replace the strikers. In 40 percent of strikes at least some replacements were employed at the end of the dispute, and in 7 percent of strikes all the strikers were replaced (or the employer closed

down).¹³ In cases where strike-breakers were used they account for an average of one-third of post-strike employment. As shown in the following table, however, this average conceals wide variation in the fraction of replacements:

Fraction of Post-Strike Employment	Percent of Strikes
.01 - .10	28.0
.10 - .20	18.3
.20 - .50	25.4
.50 - .90	15.9
.90+	12.4

In many strikes in which replacements were utilized the actual number of strike-breakers employed at the end of the strike was relatively small.

Information on the industry distribution of strikes and the characteristics of strikes in different industries is presented in Table 2. Most strikes are in the manufacturing industries, although building trades and transportation account for a significant share of disputes. In part, the industry composition in Table 2 reflects the states included in our data set: the apparel and textile industries were highly concentrated in New York state in the 1880s, while the boot and shoe industry was concentrated in Massachusetts.¹⁴ Comparisons of average size, average wages, and the causes, durations, and outcomes of strikes reveal significant differences across

¹³Unfortunately, we have no information on the timing of the decision to employ strike replacements -- only the number on hand at the end of the strike.

¹⁴See chapter 12 of U.S. Department of Commerce and Labor Manufactures 1905 (Part 1). In 1890 Massachusetts accounted for 53 percent of the output of the boot and shoe industry, while New York State accounted for 45 percent of the output of clothing industries.

industries. For example, strikes in the building trades involved relatively high-wage workers, were typically authorized by a labor organization, were relatively short, and enjoyed an above-average success rate. Strikes in the textile industry, on the other hand, involved relatively low-wage workers (over 60 percent female), were typically unauthorized, were relatively long, and had relatively low success rates.

The last three rows of Table 2 show the varying importance of the events in Spring 1886 across different industries. In two industries (machines and wood products) one-half of all strikes in our data set occurred in the first week of May 1886. In other industries (for example, building trades and food) there were relatively few strikes in 1886 prior to the Haymarket incident but a relatively high number in the post-Haymarket period.

Table 3 provides descriptive information on strikes by the cause of the dispute. The largest single category of strikes are those over wage increases. The characteristics of wage increase strikes are generally similar to the overall sample of disputes, although far fewer wage-related strikes occurred in the period from May 1 to May 7 1886. Strikes against wage cuts tended to be longer than wage increase strikes, but about equally as likely to result in a union success. Strikes over hours were heavily concentrated in the period around the Haymarket incident, and tended to be less successful (from the union's perspective) than other strikes.

Rows 11-13 show the average changes in wages and hours associated with strikes for various causes. The wage change measure is the difference between the post-strike wage of striking employees and their pre-strike wage. The hours change measure is a similar difference in the hours of workers affected by the strike. For strikes over wage increases, the average wage change is

relatively large and positive, while for strikes against wage cuts the average wage change is negative. In a significant fraction of strikes in either category, however, the wage was unchanged.

More insight into this fact is provided by Figures 2a and 2b. Here we have plotted the distributions of wage changes for strikes over wage increases and strikes against wage cuts, respectively. The overall distributions are bimodal, reflecting a mixture of strikes that ended with no wage change and strikes with mainly positive (in Figure 2a) or mainly negative (in Figure 2b) wage changes. Among the sample of strikes over wage increases, those with no wage change are almost exclusively union losses. Wage gains for the successful and compromise strikes tend to be strictly positive. Among the sample of strikes against wage cuts, those with no wage change are almost exclusively union successes, while the compromises and union losses have uniformly negative wage changes.

These plots suggest that strike outcomes in the 1880s were bifurcated. Over 90 percent of strikes were apparently resolved by a "winner-take-all" settlement. If the strike was successful, a strictly positive and often sizeable wage gain was achieved (in wage increase cases) or the wage rate was maintained (in wage cut cases). If the strike failed, no wage gain was achieved (in the wage increase cases) or a significant wage cut was enacted (in the wage cut cases). Mean wage changes by success status are shown in the following table (for strikes over wage increases):

Strike Outcome	Mean Wage Increase (%)
All	8.4
Union Win	13.6
Compromise	8.3
Union Loss	-0.1

On average, the wage gain for a successful strike was over 13 percent, while the gain for a failed strike was essentially 0. A similar pattern is evident in the distribution of hours changes for strikes over hours-related issues. Among the failed strikes, virtually all the hours changes are 0. Among the successful strikes, the distribution of hours changes is centered around -8 hours.

Table 4 presents information on wage outcomes by strike duration for the three major categories of strike causes. All three types of strikes show a systematic decline in union success rates with longer durations. The falling success rate is accompanied by an increase in the use of strike replacements. The entries in the 5th through 7th columns of the table show the conditional probabilities of a strike settlement in each duration interval, and the division of the overall settlement rate into union wins and other outcomes (losses and compromises). Over the broad duration intervals used in the table there is a uniformly decreasing conditional probability of strike settlements.¹⁵ A major factor in this decreasing hazard rate is the declining probability of a union victory: from a rate of 5-7 percent (per day) during the first 3 days of a strike to less than 1 percent per day after a month or more.

¹⁵Daily settlement patterns over finer time intervals reveal "spikes" in the settlement rates at 7, 14, and 21 days. See below.

The average wage (or hours) changes associated with strikes in each duration category are presented in the right-hand column of the table. The pattern of wage changes for the wage increase strikes in the upper panel of the table reveals a negative association between strike duration and wage increases. A similar negative relation between strike duration and the strike outcome (as viewed by workers) is also evident for strikes over wage cuts and hours cuts.

The relationship between the distribution of wage changes and strike duration (for wage increase strikes) is illustrated in Figure 3. This figure presents wage distributions similar to Figure 2a for each of the duration categories in Table 4. As in the pooled data in Figure 2a, wage settlements following strikes of different lengths are bimodal. For longer strikes, the probability of a union failure is higher, resulting in a bigger "spike" at the 0 wage change ordinate. Perhaps surprisingly, however, the distribution of wage increases conditional on a positive increase is not much different after shorter or longer strikes.

This raises the question of whether the negative relation between strike duration and average wage increases is entirely attributable to the decline in the likelihood of a union success, or whether the average wage change conditional on a union success is also affected by strike duration. A related question is whether the negative association between strike duration and average wage settlements is simply an artifact of the changing composition of the strikes in each duration category. Both questions are addressed by the multivariate regression models in Table 5.

Columns 1 and 3 of Table 5 present estimated coefficients from linear regressions of the average percentage wage increase and the incidence of a

union success on 9 duration class indicators. These models are estimated on the subset of strikes over wage increases. Comparable models for the subsample of strikes over wage cuts are presented in columns 6 and 8. In columns 2, 4, 7, and 9 we present models that include other control variables, including 11 industry dummy variables (corresponding to the industries identified in Table 2), 5 year dummy variables, measures of the size of the strike (the log of the number of workers involved) and the fraction of the employer's workers involved in the walkout, and indicators for the location of the strike, whether it was authorized by a labor organization, and whether it occurred in the first week of May 1886 or after May 7, 1886.¹⁶

Comparisons of the estimated duration coefficients in columns 1 and 2 (and columns 6 and 7) suggest that some of the negative relation between wage increases and strike duration is attributable to the effect of other covariates, particularly the authorization variable and measures of strike size and the fraction of workers involved. Nevertheless, there is still strong evidence of a negatively-sloped "resistance curve". Relative to 1-day strikes, wage increase strikes settled in the second month (29-60 days) have 4 percent lower wage gains. This effect is considerably larger than the effect estimated by McConnell (1989) for strikes in major U.S. labor contracts from 1970-81.

The linear probability models for the likelihood of a union success (in columns 3 and 8) also show a strong negative effect of strike duration. Again, the addition of other control variables dampens the duration effect. Even with a fairly rich set of controls, however, a strike of 1-2 months

¹⁶For simplicity we refer to the strikes after the first week of May 1886 as "post- Haymarket" strikes.

duration is estimated to have a 30 percentage point lower likelihood of success than a 1-day strike. Finally, columns 5 and 10 present estimated coefficients of models for the wage settlement conditional on a union success. Although several of the other covariates have significant effects, the duration coefficients are small and jointly insignificantly different from 0. Evidently, the effect of strike duration on observed wage increases works mainly through the probability of a union victory. Conditional on a successful strike, there is little systematic relation between strike duration and the actual wage settlement.

The coefficients of the control variables in the models in Table 5 reveal a number of interesting patterns. Increases in the size of a strike are associated with lower average wage increases. Comparisons of columns (2), (4), and (5) suggest that this difference is associated with a reduction in the wage increase conditional on a union victory, rather than with any effect on the likelihood of a union success. On the other hand, an increase in the fraction of the firm's employees involved in the strike increases both the probability of success and the wage increase conditional on a success. Both sets of effects are stronger among wage increase strikes.

In the 1880s, union-authorized strikes had systematically larger wage increases (or smaller wage cuts). Most of the overall effect of authorization is attributable to a significantly higher likelihood of a successful strike. (The strong connection between success rates and union authorization in these data was noted by Moore (1911)). We interpret this authorization effect not as a causal effect, but rather as a reflection of the process of union authorization. Throughout the late 19th century union leaders had an often

ambivalent attitude toward strikes.¹⁷ Janes (1916) and Ulman (1955) describe various mechanisms put in place by fledgling national unions in the 1880s and 1980s to prevent local union leaders and/or members from engaging in strikes. The goal of national union leaders seems to have been to prevent a strike when the likelihood of success was "too low". One very simple interpretation of these anti-strike policies is that union members under-valued the externality created by the continued existence of their union local. Since many of the strikers clearly earned little or no union wage premium, their private costs of engaging in a strike may have been quite low. To union leaders, on the other hand, the risk of possible extinction of the union local may have been an important additional cost of a strike.

Finally, the pattern of the coefficients for strikes that began during and after the Haymarket period suggest that these strikes were less likely to succeed than previous strikes. Strikes during the first week of May 1886 had a 10 percent lower success rate than earlier strikes, while those after May 7 had a 4.5 percent lower success rate. These patterns provide some support for the hypothesis that employer opposition was temporarily strengthened by the events leading up to and culminating in the Haymarket incident.

In summary, the descriptive evidence in Tables 1-5 and Figures 1-3 points to several tentative conclusions. First, the classification of strikes as either successes or failures was apparently justified by the nature of strike outcomes in the late 19th Century. Although a small fraction of strikes ended

¹⁷For example, in the 1886 Annual Report of the FOTLU (FOTLU, 1886, page 7), Gompers stated that "we do not, as a Federation, or as individuals, wish to be understood as advocates of strikes. On the contrary, it is well-known that the best regulated Trade Union has the least strikes". The Knights of Labor were in principle opposed to strikes, advocating instead arbitration and conciliation (Ware (1929, chapter 7)).

with a partial success or compromise, most ended with wage/hours outcomes that could be interpreted as a capitulation by one side or the other. Among strikes over wage increases, for example, almost all the successful strikes resulted in a significant wage increase, whereas almost all the failed strikes resulted in a return to work at the previous wage.

Second, close to one-half of strikes were successful. Although the gains achieved by a successful strike were potentially short-lived¹⁸, participation in a strike was by no means an exercise in futility. The large fraction of union successes in the first week of a strike is especially notable. One-third of strikes over wage increases ended in less than a week with a wage gain of 5 percent or more, and one-fifth ended in a week or less with an increase of at least 10 percent.

Third, as predicted by a variety of theoretical models (Ashenfelter and Johnson (1968); Hayes (1983); Sobel and Takahashi (1983)) longer strikes were associated with lower average wage increases or smaller hours reductions. Most of the duration effect is attributable to a decreasing likelihood of union success. Comparing 1-3 day strikes to 1-2 month strikes, for example, the union success rate was 25 percentage points higher among the shorter strikes. Strike outcomes conditional on a union success were not highly correlated with duration.

¹⁸Most observers agree, for example, that the hours reductions won by successful 8-hour day strikes in the spring of 1886 were largely reversed by the end of the year — for example David (1936, pp.539-540).

III. A Structural Model of Strike Durations and Strike Outcomes

Building on this descriptive evidence, we turn to a simple "structural" model of strike durations and strike outcomes. The model is motivated by the sharp distinction between successful and unsuccessful strike outcomes in the 1880s. For simplicity, we restrict our attention to strikes over wage increases. These strikes accounted for one-half of all disputes in the 1881-86 period and are broadly representative of other disputes in terms of durations and the likelihood of a union success. They also have the advantage that the strike outcomes are easily quantified.

The building blocks of our structural model are equations for the maximum strike durations that the employer and the workers can withstand.¹⁹ In light of the possibility of a partially successful strike, we also specify a time until the parties agree to a compromise. We assume that a strike ends when one of the maximum durations is exceeded. For example, if the maximum duration of the workers is smaller than either the time to a compromise or the maximum duration of the firm, then the strike ends as a failure (or the union "capitulates"). Likewise, a successful strike is concluded when the employer's maximum duration is less than either the time to a compromise or the workers' maximum duration. Finally, a compromise is reached when the time to a compromise is shorter than either of the maximum durations of the employer or the workers.

¹⁹We do not mean to imply that either parties' maximum strike duration is fixed ex ante. For example, employers may observe the level of support for the strike in the first few days of picketing and then decide whether to hold out or capitulate.

Formally, we specify three equations for the latent random variables D_u , D_f , and D_c , representing the maximum strike durations until the union capitulates, the firm capitulates, or a compromise is reached:

$$(1) D_u = X\beta_u + \epsilon_u$$

$$(2) D_f = X\beta_f + \epsilon_f$$

$$(3) D_c = X\beta_c + \epsilon_c$$

Here, X is a vector of observed attributes (industry and year effects, for example) and $(\epsilon_u, \epsilon_f, \epsilon_c)$ is a triplet of random error terms. Observed strike duration is D , where

$$(4) D = \min[D_u, D_f, D_c].$$

Equations (1)-(4) specify a competing risks model with 3 (possibly correlated) risks.²⁰

In this framework each party's bargaining power is summarized by its ability to withstand a strike. If one side or the other capitulates, the strike is "won" or "lost": either the workers return to work at their old wage (a failed strike) or the firm concedes to the workers' wage demand (a successful strike). The model corresponds to a simple war of attrition model in which the parties have private information over their own strike costs, and total strike costs are linear in strike duration.²¹ In such a model, once one of the parties is revealed to have a higher strike cost, the strike ends with the lower-cost bargaining party capturing all of the bargaining surplus. A difficulty with a pure war-of-attrition model, however, is the presence of compromise settlements. In light of this difficulty, and the inherent

²⁰See Kalbfleisch and Prentice (1980, chapter 7).

²¹See Kennan and Wilson (1988) for a discussion of this and other simple bargaining models of strikes. See Craig (1989) for an attempt to formally fit a war-of-attrition model to 1881 and 1891 strike data from New York state.

difficulty in distinguishing alternative theoretical models with the available data, we prefer to interpret equations (1) - (4) as a convenient statistical framework for analyzing the data, rather than as a parameterization of any particular model.

Given (1) - (4), the likelihood function for a random sample of strikes is:

$$\begin{aligned}
 (5) \quad L = & \prod_{FWIN=1} P(D = X_u\beta_u + \epsilon_u, D < X_f\beta_f + \epsilon_f, D < X_c\beta_c + \epsilon_c) \\
 & \prod_{UWIN=FWIN=0} P(D = X_c\beta_c + \epsilon_c, D < X_u\beta_u + \epsilon_u, D < X_f\beta_f + \epsilon_f) \\
 & \prod_{UWIN=1} P(D = X_f\beta_f + \epsilon_f, D < X_u\beta_u + \epsilon_u, D < X_c\beta_c + \epsilon_c)
 \end{aligned}$$

where UWIN=1 if the firm capitulates, FWIN=1 if the union capitulates, and $P(\cdot)$ is the appropriate probability statement. We begin by assuming that $(\epsilon_u, \epsilon_f, \epsilon_c)$ have a joint normal distribution, and that D_u , D_f , and D_c represent the logarithms of the union's capitulation time, the firm's capitulation time, and the time until a compromise, respectively. The likelihood function can then be written as:

$$\begin{aligned}
 (6) \quad L = & \prod_{FWIN=1} N(-Z_{fu}, -Z_{cu}; \rho_{fc.u}) 1/\sigma_u \phi[(D - X_u\beta_u)/\sigma_u] \\
 & \prod_{UWIN=FWIN=0} N(-Z_{fc}, -Z_{uc}; \rho_{uf.c}) 1/\sigma_c \phi[(D - X_c\beta_c)/\sigma_c] \\
 & \prod_{UWIN=1} N(-Z_{uf}, -Z_{cf}; \rho_{uc.f}) 1/\sigma_f \phi[(D - X_f\beta_f)/\sigma_f] .
 \end{aligned}$$

In this equation $N(\cdot, \cdot; \rho)$ is the cumulative density function for a standardized bivariate normal distribution with correlation ρ , $\phi[\cdot]$ is the standard normal density,

$$Z_{ij} = [(D - X_i\beta_i) - \rho_{ij} \sigma_i / \sigma_j (D - X_j\beta_j)] / (\sigma_i (1 - \rho_{ij}^2)^{.5}) \quad (i=u, f, c),$$

and $\rho_{ij,k}$ is the partial correlation between ϵ_i and ϵ_j , given ϵ_k .²² For example, the first line of equation (6) represents the product of the density function for the (observed) union capitulation time and the joint probability that the time to a compromise and time to a firm capitulation are both less than the observed strike duration.

In the special case where the error terms are independent, the likelihood function reduces to three independent Tobit equations, where the duration information for a particular capitulation time is censored if the strike ends with one of the other two possible outcomes. At the other extreme, the case where the three error terms are perfectly correlated reduces to a simple regression of log duration on the measured X variables. Perfect collinearity implies that the classification into union successes, partial successes, or failures is uninformative with respect to the determinants of strike duration.

The assumption that the latent determinants of strike duration are jointly normal yields a tractable model of strike durations that allows for unrestricted correlations between the unobserved determinants of the firm's and the union's capitulation times. A competing risks model with correlated heterogeneity is especially attractive in light of the numerous unobservable characteristics of the bargaining pairs and the bargaining environment. For example, we lack information on local labor market conditions and the demand for the firm's products. Both factors have been found to be useful in explaining strike durations in modern data (see Card (1990)). Another advantage of the joint-normality assumption is that it allows us to extend the model to include the wage settlement in the event of a union success (see

²²This partial correlation is equal to:

$$\rho_{ij,k} = [\rho_{ij} - \rho_{ik}\rho_{jk}] / [(1-\rho_{ik}^2)^{.5} (1-\rho_{jk}^2)^{.5}]$$

Section V below). On the other hand, a log-normal duration model imposes a strong functional form assumption on the hazard rate of strike settlements (see Kalbfleisch and Prentice (1980, pp. 24-25)). To check the sensitivity of our empirical findings to the log-normality assumption we present estimates from a uncorrelated competing risks model with semi-parametric baseline hazards in the next section.

The components of the X vector are the same control variables used in the regressions reported in Table 5. The model is also estimated with and without a dummy variable indicating the presence of strike replacements at the end of the strike. As noted earlier, strike replacements were used in about 40 percent of strikes in the 1880s. The availability of strike-breakers presumably lowered the firm's dispute costs and raised the firm's maximum strike duration. Strike replacements may have also been used as a threat to workers, with the goal of increasing the probability of a union capitulation. The problem with including this variable in equations (1)-(3) is that unless all the determinants of strike duration observed by the firm are included in X, the decision to hire replacements is likely to be endogenous with respect to the error terms (ϵ_u , ϵ_f , ϵ_u). For example, if firms are more likely to hire replacements in cases where the union can hold out longer, the estimated effect of replacements on the union's capitulation time is positively biased. A full investigation of the effect of strike replacements would require data on the timing of the decision to hire replacements, and some identifying information on which firms make this choice. Such an undertaking is beyond the scope of this paper and the available data. In the empirical results that follow, we report estimates both with and without the replacement variable, and allow the reader to evaluate both specifications.

a. Estimates of the Log-Normal Competing Risk Model

Tables 6 and 7 report OLS estimates of a linear regression model for the log of completed strike duration, and maximum likelihood (ML) estimates of the log-normal competing risks model described by equations (1)-(4). The models in Table 6 exclude the strike replacement dummy whereas the models in Table 7 include this variable. The linear regression models in the first column of each table describe strike durations without regard to the ultimate success or failure of the strike. The next three columns present the coefficients of the same explanatory variables in equations for the log of the firm capitulation time, the log of the union capitulation time, and the log of the time to a compromise settlement.

If the distinction between union victories, losses, and compromise settlements is uninformative, then the estimated coefficients in all 4 columns of each table should be similar. Inspection of the tables suggests that this is not the case. For example, the results in column 1 of Tables 6 and 7 show that union-authorized strikes were about 80 percent longer than unauthorized strikes.²³ The coefficients in columns 2-4 show that most of this effect is attributable to the positive effect of union authorization on the times to a union capitulation or a compromise. By raising the relative time to a union capitulation or compromise, authorization increases the likelihood of a union victory. This effect is evident in the linear probability models in Table 5, which show a 24 percent higher probability of a union victory in authorized strikes.

²³ This is equal to $\exp(\beta)-1$, where β is the estimated authorization coefficient reported in the table.

Analogous differences emerge in the effects of the other covariates. For example, an increase in the fraction of the firm's employees involved in a strike lowers the time to a firm capitulation and raises the time to a union capitulation, implying a strong positive effect on the probability of a union victory (see Table 5) but a statistically insignificant effect on mean log strike duration. The industry dummies (not reported in the Table) also show different effects in the equations for the three latent duration variables. We defer a discussion of the industry patterns to Section V.

The coefficients in the fifth and sixth rows of Tables 6 and 7 show the differences in strike durations and capitulation times associated with strikes beginning during and after the first week of May 1886.²⁴ Specifications with and without replacements show an increase in firm and worker capitulation times for strikes begun after May 1 1886, and a further increase for strikes initiated in the tumultuous first week of May. Relative to strikes that began in January-April 1886, firm capitulation times in the post-Haymarket period were 40-45 percent longer, while union capitulation times were 20 percent longer. Firm and union capitulation times for strikes that began in the first week of May were even further elevated: an additional 45-50 percent for firm times and 20 percent for union times.²⁵ As in Table 5, these estimates

²⁴The normalization of the coefficients in this table is different than in Table 5. The coefficients in row 6 show the effect for strikes that began on or after May 1 1886, while the coefficients in row 5 give the additional effect of strikes initiated between May 1 and May 6 of that year.

²⁵It is possible that the estimated coefficients for strikes in the first week of May 1886 simply reflect a seasonal effect common to all six years. To check we estimated independent tobit models for the firm and union capitulation times that included quarterly dummy variables, and other models that included a "May" dummy. While there is a slight seasonal effect on the firm's capitulation time in the years 1881-1885, controlling for this effect does not change the estimated effect associated with strikes in the first week of May 1886.

suggest that the nationwide eight-hour day strikes and the Haymarket incident in Chicago toughened employer resistance to union wage demands, and that part of this effect persisted to the end of 1886.

To summarize the effects of the major determinants of strike duration in the 3-equation competing risk model, we present predicted mean and median capitulation times for alternative values of the exogenous variables in Table 8. We also give the implied probabilities of a successful strike. The upper panel of the table presents results based on the models without the replacement dummy variable, while the lower panel presents results from the models that include this variable. Line (a) of the table shows the predicted mean and median capitulation times for an "average" union-authorized strike in New York State. The log-normality assumption implies that median capitulation times are shorter than mean capitulation times -- for example, the mean and median union capitulation times for an "average" strike are 36 and 20 days, respectively. Line (b) reports the means and medians for an unauthorized strike. As noted, union authorization has a strong positive effect on the union capitulation time, and a 22 percentage point effect on the likelihood of union success. Line (c) shows the capitulation times for a strike involving more workers. An increase in the size of the dispute increases both firm and union capitulation times, leading to an increase in mean observed durations and a reduction in the probability of a union success. In contrast to strike size, a change in the fraction of workers involved in the strike has opposite effects on the capitulation times of the parties (see row (d)).

Lines (e) and (f) of Table 8 show the dramatic differences between strike outcomes in Chicago and the rest of Illinois. The OLS estimates in Tables 6 and 7 suggest that the mean log strike durations in Chicago and the rest of

Illinois were not too different from those in New York. This similarity in mean durations, however, masks the fact that Illinois employers outside of Chicago were much tougher than New York employers, whereas unions were much weaker. In Chicago, on the other hand, mean and median capitulation times were roughly comparable to those in New York state.

The effects of hiring replacement workers on strike durations and the underlying capitulation times are reported in the second row of Table 7. The coefficient from the OLS duration model (in column (1)) suggests that mean log strike durations were significantly longer when replacements were used. This effect is attributable to the positive impact of replacements on the firm's capitulation time and on the time to a compromise. Replacements appear to have little "threat effect" on the capitulation time of the union. The magnitude of these various effects is illustrated in rows (g)-(j) of Table 8. A comparison of rows (g) and (i) shows the effect of replacements on capitulation times in union authorized strikes. A comparison of rows (h) and (j) shows the same effect in unauthorized strikes. In both cases the use of replacements resulted in longer firm capitulation times.

The parameters of the competing risk Tobit model specify both the capitulation times of the parties and the probability of a union success. This probability is a function of the difference between the firm and union capitulation times and is equal to the $\Pr(D_u > D_f, D_c > D_f)$. The last column of Table 8 shows the estimated probabilities of a union success for different types of strikes. For the "average" strike the probability of a union success is .7; it declines to .48 (row b) for an unauthorized strike and .24 (row j) if the strike is unauthorized and replacements are used.

The estimated correlations of the error terms in the three competing risks are presented in rows 12-14 of Tables 6 and 7. These correlations are all positive, although only the correlation between firm and union capitulation times is statistically significant at conventional levels.²⁶ This positive correlation suggests that there are unobserved characteristics of the bargaining pair (or the bargaining environment) that effect the capitulation times of both parties in the same direction. Such factors could include omitted labor or product market characteristics or, perhaps, differences in the parties' bargaining skill or "stubbornness".²⁷

Although most of the parameters in the firm and union capitulation equations are significantly different from zero, and many are significantly different between the firm and union capitulation equations, the effects of the observed covariates on the time to a compromise are far less precisely estimated. In part, this reflects the small number of compromise settlements in the data set. Except for the union authorization effect, none of the coefficients reported in the last columns of Tables 6 and 7 is statistically significant at the .05 level. The industry and year dummies are also imprecisely estimated in the compromise equation.

Three variants of the basic models presented in Tables 6 and 7 were estimated to further test whether the exogenous variables usefully predict the time to a compromise. First, a model that included only a constant term and the union authorization variable was estimated. Second, the parameters in the

²⁶We tested whether ρ_{cu} and ρ_{cf} were jointly equal to 0 in the model without replacements using a likelihood ratio test. The data do not reject this constraint ($\chi^2 = 2.578$, $p=.276$).

²⁷See Ashenfelter, Currie, Farber and Spiegel (1990) for an analysis of such "pair effects" in an experimental bargaining setting.

compromise equation (β_c) were constrained to be the same as those in the union capitulation equation (β_u). Finally, the parameters of the compromise equation were constrained to equal the parameters of the firm capitulation equation (β_f). The data reject all three sets of constraints:

Hypothesis	χ^2 value	p-value
$\beta_c=0$ except for a constant & union authorization effect	62.8	< .0001
$\beta_c=\beta_f$ except for a constant term	80.7	< .0001
$\beta_c=\beta_u$ except for a constant term	49.3	< .0001

Evidently, the variables that influence the capitulation times also effect the time to a compromise, but with magnitudes that differ from either of the capitulation equations.

One important difference between β_c and either β_u or β_f is in the coefficient of the constant term. For the "average strike" used in the calculations in Table 8 the median time to a compromise is 54 days -- over twice the median time to a firm capitulation and 5 times the median time to a union capitulation.²⁸ This relative ranking is surprisingly different from contemporary experience. Most modern strikes are settled quickly with a compromise between the demands of the union and the employer. An outright victory for one side or the other is extremely rare, and typically emerges only after a protracted dispute (often involving strike replacements).²⁹ In the 1880s, most strikes were won or lost relatively quickly. Compromise

²⁸This value is for the model without replacements.

²⁹"Classic" examples include the 10-year long strike of the United Automobile Workers against the Kohler Company, and the multi-year strike of the Machinists against Continental Airlines.

settlement were reached only after each side had discovered the other was willing and able to hold out longer than expected.

b. Simulation Tests of Model Specification

The statistical model for strike durations estimated in Tables 6 and 7 imposes the strong assumption of multivariate normality on the underlying disturbance terms $(\epsilon_u, \epsilon_f, \epsilon_c)$. Although this is a convenient assumption, it imposes a particular structure on the conditional probabilities of strike settlement that may not fit the data very well. We have two concerns with this specification. First, as shown in Table 4, the empirical hazard rates show a uniformly decreasing probability of settlement as the strike progresses. A log-normal specification cannot easily capture this feature of the data, since the hazard rate for a log-normal duration model necessarily follows an inverse U-shape (Kalbfleisch & Prentice, 1980). Second, examination of the daily settlement patterns shows "clustering" at strike durations of 7, 14, 21, ... days. This may reflect either recall errors or actual behavior tending to generate strike durations of exactly 1, 2, 3, ... weeks duration.³⁰ In either case, the smooth parametric hazard function of the log-normal distribution cannot capture spikes in the settlement rate at weekly intervals.

We pursue two strategies to assess these potential problems. First, we use a simulation to directly check the ability of the multivariate normal

³⁰Since the data were collected by the Bureau of Labor retrospectively in 1886 and 1887, we would expect to see more clustering in the early years of our sample period if recall errors are the main source of the concentration at week-long duration intervals. Checks of the data did not reveal significantly less clustering in the strikes from 1886 than in the 1881-1882 strikes.

model to reproduce the observed hazard rates. Second, in the next section of the paper we report results from a proportional hazard model with a fully non-parametric baseline hazard, and contrast the estimates from this model with the estimates from the normal model.

To evaluate the goodness-of-fit of the normal model we simulated 30 strikes for each observation using the maximum likelihood parameter estimates and the actual values of the exogenous variables for each observation.³¹ We then compared the distribution of strike durations generated by the simulation to the actual distribution of the data. In the data the mean strike duration is 15.5 days and the median is 7 days, whereas in the simulation the mean and median durations are 15.3 and 7.24 days. Likewise, the simulated union and firm win-rates (55.1 and 33.3 percent, respectively) compare very favorably to the observed 54.5 and 33.7 percent win rates.³²

Figures 4a-4d plot the actual daily settlement hazards for the sample and the implied hazards from the simulation exercise. The combined hazard rate for all types of settlements is shown in Figure 4a, while outcome-specific hazards are given in Figures 4b, 4c, and 4d. A comparison of the actual and simulated overall settlement rates in Figure 4a shows that the normal model substantially under-predicts the settlement rate on the first day of a strike. The parameter estimates imply about a 6 percent probability of settling on the first day, whereas the actual settlement rate was over 15 percent. As shown

³¹For each observation the predicted mean log duration times, $D_{u,\text{hat}}$, $D_{f,\text{hat}}$ and $D_{c,\text{hat}}$, were calculated using the estimates in Tables 6 and 7. Random error terms were then added to each predicted duration value by drawing from an appropriate trivariate normal distribution. The outcome of the strike and the observed duration were determined by the minimum of the three latent duration times.

³²These comparisons are for the model without the strike replacements variable. The model with strike replacements gives very similar results.

in Figure 4b, most of the bias arises from an under-prediction of the likelihood of a union victory on the first day of a strike. The model also has some difficulty reproducing the dip in the settlement rate in the 4th and 5th days of a strike. Apart from these shortcomings, the multivariate normal model seems to provide a reasonable fit to the observed distribution of strike durations.

IV. Semi-Parametric Estimates of Firm and Union Capitulation Times

To further investigate the limitations of the normal competing risks model, a proportional hazards competing risks model with fully nonparametric baseline hazards was estimated. Although it is unlikely that either the multivariate normal competing risks model or a proportional hazard model is the "true" model generating the strike data, a comparison of the substantive conclusions implied by the two models provides a useful check on the extent to which these conclusions are driven by functional form assumptions, rather than by the data.

The specification of the proportional hazard model is similar to the specification used by Moffitt (1985), and includes an unrestricted baseline hazard for the settlement rate in each day of the strike up to some maximum duration. In contrast to the normal model, however, we assume that the union and firm capitulation times and the time to a compromise are independent. The probabilities that the union and firm concede on day t (conditional on not conceding before t) are described by the following two equations:

$$(7a) \lambda_{ut} = \lambda_u(t) \exp(X\beta_u)$$

$$(7b) \lambda_{ft} = \lambda_f(t) \exp(X\beta_f),$$

where $\lambda_u(t)$ and $\lambda_f(t)$ are "baseline" probabilities that the firm or union will capitulate on day t , and deviations from this day-specific settlement rate are proportional to the values of the exogenous variables. Because the baseline probabilities are unrestricted, this model can capture the spikes in the settlement rates at weekly intervals as well as the shape of the settlement hazard in the first few days of a strike. On the other hand, it does not allow for unobserved heterogeneity in the hazard rates, and it also abstracts from any correlation between the unobserved components of the union and firm capitulation rates.³³

Equations (7a) and (7b) were estimated separately for the union and firm capitulations, using information on strikes lasting up to 21 days.³⁴ Eighty percent of wage increase strikes were settled in the first 21 days. After this point the number of surviving strikes is small and the daily settlement rates are also low, making it impossible to estimate separate baseline parameters for each day.³⁵

Table 9 and Figures 5a and 5b present the estimates of β_i and $\lambda_i(t)$ for the firm and union capitulation equations. Again we have estimated the models with and without a variable indicating the presence of replacement workers. For ease of interpretation of the baseline hazard parameters, we have re-scaled the variables measuring the number of workers on strike and the

³³Both limitations can be relaxed by specifically modelling unobserved components in the daily hazards that are fixed across bargaining pairs.

³⁴Details of the construction of the likelihood are presented in Appendix B.

³⁵ Even with a 21 day censoring point, it was not possible to estimate the union capitulation probability on day 19, because no strikes ended on day 19 by the capitulation of the union. The probability of a union capitulation on this day was set to zero.

fraction of the firm's workforce on strike to be deviations from the sample means. The estimated baseline parameters graphed in Figures 5a and 5b therefore represent the daily settlement rates for an unauthorized "average size" strike in New York State in 1881 in the miscellaneous industry category in which replacements were not used.

Since a higher hazard rate implies a lower expected capitulation time, the signs of all the coefficients in Table 9 are reversed relative to the coefficients in Tables 6 or 7. The magnitudes of the estimated coefficients, however, are generally similar to those of the normal competing risks model.³⁶ With the exception of the replacements variable, the statistical significance of the individual coefficients is also similar in the two specifications. In the normal competing risks model strike replacements do not have a significant impact on the union capitulation time, whereas in the proportional hazards model the use of replacements significantly increases the union capitulation rate, providing some evidence of a "threat effect".

The plots of the baseline hazards in Figures 5a and 5b show a steady decrease in the capitulation rates of both the firm and union, with large "spikes" in the hazards on the 7th and 14th days of a strike. As noted earlier, we suspect that these spikes represent a combination of reporting error (the tendency to report strike durations to the nearest week) and actual behavior induced by such institutional considerations as the scheduling of union strike votes. The estimated baseline settlement rate declines very quickly in the first few days of a strike. This measured decline may reflect

³⁶This is readily explained by the fact that $E(\log y) = c - X\beta$ (where c is a known constant) if y is exponentially distributed with hazard $X\beta$. The non-parametric baseline model does not impose a constant hazard. If the baseline parameters were all constant, however, it would be equivalent to an exponential model.

either true duration dependence, or the fact that the firms most likely to concede defeat do so in the early days of the strike.

There is no simple way to compare the magnitudes of the estimated effects from the log-normal competing risks model and the semi-parametric proportional hazard model (unless the estimated baseline is constant). Furthermore, it is impossible to translate the semi-parametric hazard estimates into expected capitulation times, since the model only uses information on strikes lasting up to 21 days.³⁷ Nevertheless, it is possible to use both models to calculate the probability of a capitulation before the 21st day of a strike, and the expected capitulation time, given that the strike ends in the first 21 days. We can also calculate how changes in the exogenous variables alter these predictions in the two different models. This analysis is shown in Table 10.

The first row of this table reports information for an "average strike" -- an unauthorized strike in New York State in 1881, in the omitted (miscellaneous) industry group, in which replacements were not used, and for which the log of the number of strikers and fraction of workers on strike equal their sample means. Columns (1)-(6) give the expected log capitulation times (for the firm and the union) assuming that a capitulation occurs within 21 days, while columns (7)-(12) give the probabilities of a capitulation after the first 21 days. We present estimates associated with 3 different models: the semi-parametric hazard model (columns (1), (4), (7), and (10)); a constrained version of the normal competing risks model in which the capitulation times are assumed to be independent (columns (2), (5), (8), and (11)); and the unconstrained normal model (columns (3), (6), (9), and (12)).

³⁷This is a general problem with any non-parametric baseline model.

For example, for an "average strike", the mean of the logarithm of the firm's capitulation time given that the capitulation time is ≤ 21 days is 1.53 using the proportional hazards model, and 1.47 using the unconstrained log-normal model. The corresponding union capitulation times are 1.93 (proportional hazard model) and 1.84 (unconstrained log-normal model).

Rows (b)-(h) report the increases in the conditional means (or the increases in the probability of not capitulating within 21 days) for a strike identical to the "average" strike except for the change described by the row label. For example, column (1) row (b) shows the increase in the conditional mean if the strike was authorized by a union, using the proportional hazards model. Column (3) row (b) shows the same increase implied by the unconstrained Tobit model. Rows (i)-(l) show the derivatives of the conditional means (or the derivatives of the probability of holding out for at least 21 days) with respect to the number of strikers and the fraction of the workforce on strike.

The entries in columns (1)-(6) of Table 10 suggest that the estimated effects of the exogenous variables on the capitulation times of the parties are very similar in the three different specifications. The unconstrained log-normal model consistently predicts larger changes than the proportional hazard model for firm capitulation times (compare column (3) versus column (1)). On the other hand, there is no uniform pattern in the predictions of the union capitulation times (compare columns (4) and (6)). The entries in columns (7)-(12) are even more similar across specifications. These findings lead us to conclude that a normal competing risks model and a proportional hazards model with a flexible baseline specification have very similar

implications with respect to the effects of the measured covariates on strike durations and outcomes.³⁸

V. Wage Settlements In the Event of a Successful Strike

Figure 2a shows that virtually all of the wage gains that occurred in wage increase strikes in the 1880s were confined to strikes that ended with a firm capitulation or a compromise. In 93 percent of cases where the union capitulated, the strike ended without a wage increase. Thus, to model wage settlements associated with strikes in the 1880s, we need only focus on explaining the wage increase conditional on a union success or a compromise. In view of the small number of compromises, we focus in this section on modelling wage settlements following a union strike success (i.e., a firm capitulation).

We proceed by expanding the statistical model of the preceding section to include wage changes in the event of a firm capitulation. The "structural" wage equation is:

$$(8) \Delta \ln(W_t) = X\beta_w + \nu_w,$$

where $\Delta \ln(W_t)$ represents the change in the logarithm of wages if the firm concedes the strike.³⁹ One possible interpretation of (8) is as the union's wage demand, which is fixed at the outset of the strike but only realized in the event that the firm concedes. The fact that the observed wage settlements

³⁸ In fact, Table 10 may overstate the differences between the proportional hazard and the normal competing risks estimates. The log-normal competing risks models underlying this table were estimated using information on all strikes, whereas the proportional hazard model is estimated by ignoring information on strikes that settled after 21 days (other than the fact that they lasted over 21 days).

³⁹ Note that an equation similar to (8) can be used to describe wage changes conditional on a compromise.

conditional on a union victory do not depend strongly on the duration of the strike (see Table 5) lends some support to this interpretation.

Equation (8) can only be estimated without bias on the subset of successful strikes if ν_f is uncorrelated with the unobserved determinants of duration. Otherwise, the conditional expectation of the wage change, given a union victory, is:

$$(9) \quad E(\Delta W_f) = X\beta_w + E(\nu_w \mid D=X_f\beta_f + \epsilon_f, D < X_u\beta_u + \epsilon_u, D < X_c\beta_c + \epsilon_c)$$

There are several approaches to dealing with the potential sample selection bias implied by the second term in this equation. One simple approach is the two-step procedure suggested by Heckman (1978). Since the results reported in Tables 6 and 7 imply that ρ_{uc} is approximately 0, Heckman's procedure in this case amounts adding two "selection terms" to the OLS regression equation, corresponding to the means of two truncated univariate normal distributions, $E(\epsilon_u \mid D < X_u\beta_u + \epsilon_u)$ and $E(\epsilon_c \mid D < X_c\beta_c + \epsilon_c)$. The values of these two selection terms were calculated using the estimates reported in Tables 6 and 7 and then used to estimate a selection-corrected version of Equation (8). These estimates, which are available from the authors, are very similar to the OLS estimates and provide no evidence that there is any correlation between ν_w and either ϵ_u or ϵ_c .

An alternative "full information" approach is to add equation (8) to the system of equations (1)-(3). This approach has the advantage of providing an estimate of ρ_{wf} , the correlation between unobserved factors that influence the wage settlement given a union win and unobserved factors that determine how long the firm will hold out before capitulating. An important disadvantage of a full information approach is computational complexity: the system of (1)-(3)

and (8) has over 100 parameters if we include unrestricted year and industry effects in all 4 equations. A reasonable compromise is to drop the equation for the time to a compromise settlement and treat "partially successful" strikes as independently censored observations. Following this approach it is possible to estimate β_w , β_f , β_u and the correlations between ϵ_u , ϵ_f and v_w using the entire sample of wage increase strikes.⁴⁰

The second columns in Tables 11 and 12 report maximum likelihood (ML) estimates of β_w based on the system of (1), (2) and (8).⁴¹ For comparative purposes we also report OLS estimates of the wage change equation (fit to the subset of successful strikes). The ML estimates are similar to the OLS estimates, although generally bigger in magnitude. This is consistent with the attenuation biases associated with a nonrandomly selected sample. The estimated correlations between ϵ_u , ϵ_f and v_w are presented in rows 12-14 of the tables, and suggest a positive correlation between the union capitulation time and the error in the wage change equation, and a negative correlation between the wage equation and the firm capitulation time. The estimate of ρ_{uw} is relatively small (and statistically insignificant) when the replacement variable is excluded from the model, but rises in magnitude when the replacement variable is included. The estimate of ρ_{uf} is positive and somewhat larger than the estimate obtained from the three outcome competing risks model that ignored wages (see Table 6 or 7).

⁴⁰Appendix C gives the complete expression for the likelihood of this extended model.

⁴¹ The estimates of β_u and β_f are not reported but available from the authors. The estimates are virtually identical to the estimates reported in Tables 6 and 7.

The effects of the observed variables on the wage settlement (conditional on union victory) show an interesting pattern. In cases where a variable increases the capitulation time of the union more than the capitulation time of the firm, the variable also tends to raise the wage settlement. For example, union authorization raises the capitulation times of both parties, but the effect on the union time is substantially larger. Union authorization is also associated with higher wage increases in the event of a union victory. To further illustrate this pattern we report in the third columns of Tables 11 and 12 the difference in the estimated effects of each variable on the union and firm capitulation times. Reading down the rows of the tables it is clear that variables for which $(\beta_u - \beta_f)$ is positive (negative) tend to have a positive (negative) β_w .

a. Union Bargaining Power and Wages Across Industries

Our estimates of union and firm capitulation times and the wage settlement conditional on a union victory can be used to evaluate the hypothesis that a union's bargaining power is related to its relative ability to withstand a strike. As shown in Tables 11 and 12, there is strong prima facie evidence in favor of this view. Measured variables that raise the union's capitulation time relative to the firm's also tend to raise the wage settlement, conditional on a union victory. Similarly, unobserved factors that raise the union's capitulation time tend to raise the wage settlement ($\rho_{uw} > 0$), whereas unobserved factors that raise the firm's capitulation time tend to lower the wage settlement ($\rho_{fw} < 0$).

This relationship between wage increases and relative bargaining power as measured by the difference in capitulation times extends to the pattern of the

industry effects included in equations (1), (2) and (8). The estimated values of the industry dummies in the union and firm capitulation time equations are presented in columns (1) and (2) of Tables 13 and 14. Their differences (and their associated standard errors) are reported in the third column of each table. Finally, the industry effects from the wage equations are presented in the 4th column of each table. The correlation across industries between the entries in columns (3) and (4) is 0.63.⁴² As is true for the other determinants of strike duration and wages, the patterns of the industry effects confirm a simple "relative bargaining power" hypothesis.

Figure 6a presents a plot of the estimated industry effects in Table 13. The industry effects from the wage settlement equation are graphed on the vertical axis against the estimate of $\beta_u - \beta_f$ for each industry on the horizontal axis. The 12 points in the figure correspond to the eleven industry effects presented in Table 13 and the omitted industry group.⁴³ The size of the circle used to graph each industry is proportional to the industry's share of strikes in our sample. Except for the tobacco industry, the data points lie on a positively sloped line, confirming the strong association between the relative ability of workers in each industry to withstand a strike, and their expected wage gain conditional on a union victory.

The null hypothesis that the industry wage differences are entirely explainable by differences in capitulation times can be tested by constraining the industry wage effects to be proportional to the difference in industry

⁴² The correlation is also .627 when the replacement dummy is included in the model.

⁴³ For the omitted industry, $\beta_u - \beta_f$ is the difference in the constant terms in the two capitulation equations.

effects in the union and firm capitulation equations. . Specifically, a simple version of the relative bargaining power hypothesis states that

$$\beta_{w,i} = \alpha (\beta_{u,i} - \beta_{f,i}),$$

where $\beta_{w,i}$ is the i th industry effect in the wage equation, $\beta_{u,i}$ and $\beta_{f,i}$ are the corresponding industry effects in the union and firm capitulation times, and α is a positive constant. The data strongly reject this restrictive hypothesis.⁴⁴ Much of the difficulty is attributable to the tobacco industry, which had relatively powerful unions (judging by their relative capitulation coefficients) but earned relatively small wage increases in the event of a union victory. Nevertheless, the estimate of α is positive and significant (the estimate is 0.18, $t = 2.6$), suggesting that differences in capitulation times were an important determinant of industry differences in the wage changes associated with a strike. Indeed, a one-standard deviation increase in $(\beta_{u,i} - \beta_{f,i})$ is associated with a 9.6 percent higher wage increase, conditional on a successful strike.

The unusual pattern of the tobacco industry could be due to one of several factors. Unlike most other workers, employees in the tobacco industry started from a position of relative union strength in the 1870s and saw that strength decline during the 1880s with technological innovations in the industry.⁴⁵ This trend could account for their relatively high "strike power" but relatively low wage increases. Another idiosyncratic factor in the tobacco industry during this period was the bitter rivalry between two different cigar makers unions in New York City -- the Cigar Makers International Union (lead

⁴⁴ The χ^2 value with 11 degrees of freedom equals 46.8 ($p < .001$).

⁴⁵ In the 1880s highly-skilled cigar rollers at many firms (particularly in New York City) were replaced by less-skilled operatives working at cigar molding machines. See Ware (1929, chapter 11).

by Adolph Strasser and later Gompers), and the Progressive Cigar Makers (backed by the Knights of Labor). In any case, the industry is clearly an "outlier" in our data.

One of the important implications of a relative bargaining power hypothesis is that it is the difference in capitulation times of the firm and the union that influence the wage, and not simply the average duration of strikes. It is possible to have two industries with similar values of $(\beta_{u,i} - \beta_{f,i})$ which have different average strike durations. Similarly, it is possible to have two industries with similar mean strike durations for which $(\beta_{u,i} - \beta_{f,i})$ is positive or negative. In fact, in our strike data there is no significant inter-industry relationship between the average duration of strikes and the average industry wage effect. This is shown in Figure 6b, which plots the industry effects from the wage equation against the estimated industry effects from a model for mean log strike duration.⁴⁶ In fact the plot shows a weak negative relationship ($\rho = -.186$) between the industry wage differentials and the industry mean duration differentials. A comparison of Figures 6a and 6b underscores the importance of clearly distinguishing between the capitulation times of the two parties in understanding the implications of strike duration for relative bargaining strength.

Finally, it is interesting to compare the pattern of wage and strike duration coefficients in Tables 13 and 14 with the predictions from a simple war-of-attrition model in which the wage settlement conditional on a union win is taken as exogenous. Higher wage settlements conditional on a union victory imply a bigger gain to workers if they win, and a bigger gain to the firm if

⁴⁶Industry effects for mean log duration are estimated from a simple Tobit model that treats compromise settlements as censored. Estimated effects from a simple OLS equation estimated to all the strike observations are very similar.

it wins. Hence, a war-of-attrition model predicts that $\beta_{u,i}$ and $\beta_{f,i}$ should both be positively related to $\beta_{w,i}$. It follows that overall strike durations in an industry should be related to the average size of the wage settlement. Similarly, any factor that raises the wage settlement conditional on a union win should raise both the firm and union capitulation times. Examination of the industry effects and many of the other coefficients in the duration and wage equations provides strong evidence against this prediction.

VI. Summary and Conclusion

In this study we take advantage of an unusually rich data set of strikes that was published over 100 years ago. These data, which are as detailed as many recent data sets on contemporary strike activity, allow us to investigate the role of strikes in determining the relative bargaining power of unions. We provide empirical support for the hypothesis that:

The relative degree to which workers and the employer are willing and able to sustain a strike is their *strike leverage*...Holding other things constant, then, the simplest measure of bargaining power is the amount of strike leverage each party holds (Kochan & Katz, 1988, p. 54).

The challenge in testing this hypothesis is to define bargaining power in terms other than the outcomes of bargaining. Otherwise, the definition is tautological: unions have bargaining power when we observe favorable bargaining outcomes, and favorable bargaining outcomes imply powerful unions.

Unlike most contemporary strikes, strikes in the 1880s almost always produced a "winner" and "loser". Observations on the duration of a strike and on which side "won" the strike therefore identify both the party with the greater bargaining power and the length of time the losing party was able to hold out before capitulating. We use this information to identify the factors that determine how long each side can withstand a strike. We then define the

relative bargaining power (or strike leverage) of the union as the difference between the time the union and firm can strike (or take a strike) before capitulating.

We find that union capitulation times were longer when: (a) the strike was ordered or authorized by the union, (b) a large number of workers participated in the strike and (c) the fraction of the establishment's workforce that struck was large. Firm capitulation times were longer when: (a) a large number of workers participated in the strike, (b) the fraction of the establishment's workforce that struck was small, (c) the firm hired strike replacements, and (d) the strike occurred on or after the first week of May 1886. We also find significant differences in capitulation times across different regions and industries.

Variables that alter the difference in the capitulation times of the parties are interpreted as changing their relative bargaining power. Factors having a statistically significant correlation with the bargaining power of the union are:

Variable	Estimated Effect on Union Bargaining Power
Union-authorized strike	Increase
Firm Hired Replacements	Decrease
Log (Number of Strikers)	Decrease
Fraction of workers Involved in Strike	Increase
Strike on/after May 1, 1886	Decrease
Strike in Illinois outside of Chicago (relative to Chicago)	Decrease

Each variable that increases union bargaining power (as measured by relative capitulation times) has a positive effect on the wage outcome in the event of a union victory, and each variable that lowers union bargaining power has a negative effect on the wage outcome, conditional on a union victory. A similar pattern is present across industries: in industries with a higher relative capitulation time for the union, the wage settlement conditional on a union success is higher. These results suggest a strong connection between the ability to weather a strike and the outcome of bargaining at the conclusion of a strike.⁴⁷

Because of the very different structure of labor markets and institutions in the 1880s, an investigation of strike activity during this period provides new evidence on the generalizeability of contemporary strike models. For example, many recent theoretical models focus on the strike costs that arise from the gap between a worker's productivity at the incumbent firm and his or her productivity at other firms. It is not obvious, however, that most workers in the 1880s were earning such firm-specific rents. On the other hand, aspects of observed behavior in the 1880s suggest an important factor absent from most recent theoretical strike models. In the late 19th century many unions tried to prevent strikes at the local level.⁴⁸ We suspect that this reflects a conflict between the interests of workers and the institutional needs of the union. Workers earning a competitive wage may have been willing to risk a strike with a low probability of success, since the

⁴⁷ It would be useful in future research to further explore this result and its implications for rivalry between the Knights of Labor and the AFL.

⁴⁸See Ulman (1955). Ulman's thesis is that the development of national control over strike policies was necessitated by the over-zealousness of local union organizations in calling strikes.

loss of a job was relatively costless to them. The union, however, needed to protect jobs to prevent the dissolution of the organization. This suggests a model of strikes that, like Ashenfelter and Johnson's (1969) model, emphasizes the conflicting objectives of the union leadership and membership.

Finally, our results suggest the importance of broader social and political variables in determining the relative bargaining power of the parties. We find substantial spillovers from the "nationwide" 8-hour day strike and the Haymarket Square incident to the bargaining behavior of firms later in 1886. Our estimates suggest these events increased employer resistance to union wage demands and reduced the likelihood of a successful strike for many months. It is interesting to speculate whether the example of tough bargaining set by Chicago employers (backed by political forces in the city) was enough to harden the resistance of employers in other states. If so, the Haymarket example lends credence to the role of the PATCO strike in determining the course of U.S. labor relations in the 1980s.

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Appendix A
Listing of Strikes Deleted From Final Data Set ^a

Case	Beginning Date		City	Number Strikers	Previous		Duration	New Wage	Days Closed	\$ Losses of Parties	
	Month	Day			Wage	Success				Workers	Firm(s)
<u>1881</u>											
553	6	1	Chicago	250	2.00	Y	2	2.25	2	1000	1000
554	6	1	Cairo	40	2.00	Y	3	2.25	3	240	250
555	6	1	Chicago	60	2.00	Y	1	2.25	1	120	200
556	6	15	Chicago	20	2.50	Y	2	3.00	2	100	500
557	6	15	Chicago	50	2.00	Y	3	2.25	3	100	200
558	7	15	Chicago	150	2.00	Y	3	2.25	3	900	1000
561	9	15	Chicago	200	2.25	N	3	2.25	3	1126	1200
562	10	1	Chicago	20	3.00	N	2	3.00	2	120	500
563	10	1	Cairo	40	2.25	N	3	2.25	3	270	250
564	10	15	Chicago	60	2.25	N	1	2.25	1	150	150
565	10	15	Chicago	250	2.25	N	2	2.25	2	1125	1000
<u>1882</u>											
634	6	1	Chicago	60	2.00	Y	1	2.25	1	120	200
635	6	1	Chicago	250	2.00	Y	2	2.25	2	1000	1000
636	6	1	Cairo	40	2.00	Y	3	2.25	3	240	250
637	6	15	Decatur	20	2.50	Y	2	3.00	2	100	500
638	6	15	Chicago	50	2.00	Y	3	2.25	3	300	200
640	7	15	Chicago	150	2.00	Y	3	2.25	3	900	1000
644	9	15	Chicago	200	2.25	N	3	2.25	3	1126	1200
646	10	1	Chicago	20	3.00	N	2	3.00	2	120	500
647	10	1	Cairo	40	2.25	N	3	2.25	3	270	250
648	10	15	Chicago	250	2.25	N	2	2.25	2	1125	1000
649	10	15	Chicago	60	2.25	N	1	2.25	1	150	200
<u>1883</u>											
706	6	1	Cairo	40	2.00	Y	3	2.25	3	240	250
707	6	1	Chicago	250	2.00	Y	2	2.25	2	1000	1000
708	6	1	Chicago	60	2.00	Y	1	2.25	1	120	200
710	6	15	Chicago	50	2.00	Y	3	2.25	3	100	200
711	6	15	Chicago	20	2.50	Y	2	3.00	2	100	500
712	7	15	Chicago	150	2.00	Y	3	2.25	3	900	1000
713	9	15	Chicago	150	2.25	N	3	2.25	3	1013	1000
714	9	15	Chicago	50	2.25	N	3	2.25	3	113	200
715	10	1	Cairo	40	2.25	N	3	2.25	3	270	250
716	10	1	Chicago	20	3.00	N	2	3.00	2	120	500
718	10	15	Chicago	60	2.25	N	1	2.25	1	150	200
719	10	15	Chicago	250	2.25	N	2	2.25	2	1125	1000

Appendix A, continued

Case	Beginning Date		City	Number Strikers	Previous Wage	Success	Duration	New Wage	Days Closed	\$ Losses of Parties	
	Month	Day								Workers	Firm(s)
<u>1884</u>											
761	6	1	Cairo	40	2.00	Y	3	2.25	3	240	250
762	6	1	Chicago	250	2.00	Y	2	2.25	2	1000	1000
763	6	1	Chicago	60	2.00	Y	1	2.25	1	120	200
764	6	15	Chicago	50	2.00	Y	3	2.25	3	300	200
765	6	15	Chicago	20	2.50	Y	2	3.00	2	100	500
766	7	15	Chicago	150	2.00	Y	3	2.25	3	900	1000
767	9	15	Chicago	150	2.25	N	3	2.25	3	1013	1000
768	9	15	Chicago	50	2.25	N	3	2.25	3	448	200
769	10	1	Cairo	40	2.25	N	3	2.25	3	270	250
770	10	1	Chicago	20	3.00	N	2	3.00	2	120	500
771	10	15	Chicago	60	2.25	N	1	2.25	1	135	200
772	10	15	Chicago	250	2.25	N	2	2.25	2	1125	1000
<u>1885</u>											
839	6	1	Mound Cty	40	2.00	Y	3	2.25	3	240	250
840	6	1	Chicago	60	2.00	Y	1	2.25	1	120	200
841	6	1	Chicago	250	2.00	Y	2	2.25	2	1000	1000
842	6	1	Cairo	40	2.00	Y	3	2.25	3	240	250
843	6	15	Chicago	50	2.00	Y	3	2.25	3	300	200
844	6	15	Chicago	20	2.50	Y	2	3.00	2	100	500
847	7	15	Chicago	150	2.00	Y	3	2.25	3	900	1000
848	9	15	Chicago	200	2.25	N	3	2.25	3	1338	1200
849	10	1	Cairo	80	2.25	N	3	2.25	3	540	500
850	10	1	Chicago	20	3.000	N	2	3.00	2	120	500
851	10	15	Chicago	60	2.250	N	1	2.25	1	135	200
852	10	15	Chicago	250	2.250	N	2	2.25	2	1125	1000

Appendix A, continued

Beginning Date			City	Number	Previous	Success	Duration	New	Days	\$ Losses	of Parties
Case	Month	Day		Strikers	Wage			Wage	Closed	Workers	Firm(s)
<u>1886</u>											
1137	4	20	Peoria	60	1.500	Y	2	1.75	2	180	100
1139	5	1	Chicago	150	2.000	Y	3	2.25	3	900	1000
1140	5	1	Chicago	60	2.000	Y	1	2.25	1	120	200
1141	5	1	Chicago	250	2.000	Y	2	2.25	2	1000	1000
1158	6	1	Cairo	40	2.000	Y	3	2.25	3	240	250
1159	6	1	Chicago	150	2.250	N	2	2.25	2	300	1000
1160	6	1	Chicago	60	2.250	N	1	2.25	1	150	200
1161	6	1	Mound Cty	40	2.000	Y	3	2.25	3	240	250
1162	6	15	Chicago	50	2.000	Y	3	2.25	3	300	200
1163	6	15	Chicago	20	2.500	Y	2	3.00	2	100	500
1164	7	15	Chicago	150	2.250	N	3	2.25	3	1013	1000
1165	9	15	Chicago	150	2.250	N	3	2.25	3	1013	1000
1166	9	15	Chicago	50	2.250	N	3	2.25	3	330	200
1167	10	1	Cairo	40	2.250	N	3	2.25	3	270	250
1168	10	1	Mound Cty	40	2.253	N	3	2.25	3	270	250
1169	10	1	Chicago	20	3.000	N	2	3.00	2	120	500
1170	10	15	Chicago	150	2.250	N	2	2.25	2	1125	1000
1171	10	15	Chicago	60	2.000	Y	1	2.25	1	120	200
1172	11	1	Chicago	250	2.250	N	2	2.25	2	1125	1000

^a All strikes are unauthorized, and involve laborers and wharf hands in Illinois.

Appendix B

Likelihood Function For the Semi-parametric Proportional hazard Model

To estimate equations (7a) and (7b) we consider all strikes still in progress after 21 days as censored at 21 days. 80.5 percent of the strikes were settled by one of the three outcomes by the end of day 21. After 21 days the number of daily settlements due to either a firm or union capitulation was very small, making estimation of $\lambda_i(t)$ impossible or very unreliable. The probability of a union capitulation on day T where $T \leq 21$ days is:

$$(8) \Pr(D_u=T) = \lambda_u(T) \exp(X\beta_u) \prod_{j=1}^{T-1} [1-\lambda_u(j) \exp(X\beta_u)].$$

A comparable equation describes the probability that a firm capitulation is observed on day T. If the union capitulation time was censored because the strike was settled sometime after day 21, or because of a compromise, or because of firm capitulation, we only observe that $D_u > T$. Letting $C=1$ if the observation is censored at T, the contribution of an individual observation to the likelihood function for the union capitulation times is:

$$(9) L_u = (\lambda_u(T) \exp(X\beta_u) \prod_{j=1}^{T-1} [1-\lambda_u(j) \exp(X\beta_u)])^{(1-C)} \prod_{j=1}^T ([1-\lambda_u(j) \exp(X\beta_u)])^{(C)}$$

Appendix C

Estimation of the Joint Model of Capitulation Times and Wage Change Given a Union Victory

The estimates reported in Tables 11 and 12 are the MLE for the following four equations:

$$(1) D_u = X\beta_u + \epsilon_u$$

$$(2) D_f = X\beta_f + \epsilon_f$$

$$(4) D = \min[D_u, D_f, D_c]$$

$$(8) \Delta \ln(W) = X\beta_w + \nu_w \text{ iff } D_u > D_f$$

where $(\epsilon_u, \epsilon_u, \nu_w)$ are assumed to be distributed trivariate normal. The compromise equation is dropped from the model because the results in Tables 5, 6 and the 2-step Heckman estimates of Equation (8) lead us to conclude that ϵ_c is independent of the other three error terms - ϵ_u , ϵ_u and ν_w . In the analysis we assume D_u , D_f and $\Delta \ln(W)$ are censored at time D_c when a compromise settlement occurs. The likelihood function for this model is

$$(C1) L = \prod_{FWIN=1} \Pr(\epsilon_u = D - X\beta_u) \Pr(\epsilon_f > D - X\beta_f \mid \epsilon_u = D - X\beta_u)$$

$$\prod_{UWIN=1} \Pr(\epsilon_u > D - X\beta_u \mid \epsilon_f = D - X\beta_f, \nu_f = \Delta \ln(W) - X\beta_w) \times$$

$$\Pr(\epsilon_f = D - X\beta_f, \nu_f = \Delta \ln(W) - X\beta_w)$$

$$\prod_{FWIN=UWIN=0} \Pr(D < X\beta_u + \epsilon_u, D < X\beta_f + \epsilon_f)$$

$$FWIN=UWIN=0$$

Figure 1
Number of Strikes by Month

Figure 2a
Distribution of Wage Increases
Strikes Over Wage Increases

Figure 2b
Distribution of Wage Increases
Strikes Over Wage Cuts

Figure 3

Distribution of Wage Settlements
By Strike Duration: Wage Increase Cases

Figure 4a
Simulated and Actual Hazard Rate For All Settlement Types

Figure 4b
Simulated and Actual Hazard Rate For Union Wins

Figure 4c
Simulated and Actual Hazard Rate For Firm Wins

Figure 4d
Simulated and Actual Hazard Rate For Compromises

Figure 5a
Baseline Hazard For Firm Capitulation

Figure 5b
Baseline Hazard For Union Capitulation

Figure 6a
Relative Union Strike Power & Wage Gains Across Industries
($\rho = .627$)

Figure 6b
Differences in Strike Duration & Wage Gains Across Industries
($\rho = -.186$)

Table 1
Mean Characteristics of Strikes, By State and Year

	All (1)	By State:			By Starting Year:					
		Illinois (2)	Mass (3)	New York (4)	1881 (5)	1882 (6)	1883 (7)	1884 (8)	1885 (9)	1886 (10)
1. Number Strikes	2179	705	300	1174	219	207	205	234	299	1015
2. Average Size	245	238	122	281	188	230	330	180	218	267
3. Average Log Size	4.04	4.44	3.66	3.90	4.25	4.29	4.27	3.63	4.10	3.98
4. Average Fraction Involved (%)	80.6	89.7	41.9	85.0	84.4	83.4	84.6	83.9	75.7	79.1
5. Avg Prev Wage (\$/day)	2.12	1.99	2.01	2.22	1.97	2.15	2.12	2.43	2.08	2.08
6. Pct Authorized	76.8	74.0	57.7	83.3	69.4	67.6	79.0	77.8	73.6	80.5
<u>Causes of Strike:</u>										
7. For a Wage Increase (%)	47.1	38.2	54.0	50.7	69.4	63.3	53.7	39.3	50.8	38.4
8. Against a Wage Cut (%)	11.6	12.1	15.7	10.2	10.0	11.6	18.5	24.8	18.4	5.4
9. For a Change in Hours (%)	18.3	30.9	3.3	14.6	5.4	4.3	1.0	11.5	1.0	34.1
<u>Duration:</u>										
10. Mean (days)	20.0	19.8	29.9	17.7	12.7	19.0	19.0	23.2	33.7	17.2
11. Median (days)	9.0	10.0	14.0	7.0	7.0	7.0	8.0	7.0	11.0	10.0
<u>Outcomes:</u>										
12. Union Win (%)	46.9	31.2	41.7	57.7	52.1	52.7	57.6	58.1	55.9	37.2
13. Union Loss (%)	43.8	56.6	49.0	34.8	34.2	41.1	37.1	39.3	40.5	49.9
14. Compromise (%)	9.2	12.2	9.3	7.3	13.7	6.3	5.4	2.6	3.7	12.7
15. Strike Breakers Employed (%)	38.7	38.6	33.0	40.3	41.0	36.7	47.8	34.6	37.1	38.2
16. All Strikers Replaced (%)	6.9	0.9	12.0	9.2	5.0	3.9	9.3	12.0	5.4	6.7

Notes: Sample is drawn from Third Annual Report of the Commissioner of Labor, and includes all strikes in Illinois, Massachusetts, and New York occurring between 1881 and 1886. See text for further discussion.

Table 2
Mean Characteristics of Strikes, By Industry

	Apparel (1)	Bldg Trades (2)	Food (3)	Machine (4)	Metals (5)	Mining (6)	Shoes (7)	Textiles (8)	Tobacco (9)	Transp (10)	Wood Prods (11)	Miscell (12)
1. Number Strikes	98	268	84	92	260	144	143	77	240	122	226	425
2. Average Size	545	275	433	201	116	200	91	283	320	524	150	198
3. Average Fraction Involved (%)	76.6	93.3	82.1	72.8	77.9	94.1	46.6	33.1	84.9	94.4	90.0	80.7
4. Avg Prev Wage (\$/day)	2.03	2.95	2.01	2.34	2.29	1.80	2.07	1.38	1.82	1.83	2.03	2.12
5. Pct Authorized	78.6	92.9	88.1	87.0	74.6	55.5	87.4	15.5	95.8	50.8	94.2	65.2
<u>Causes of Strike:</u>												
6. For a Wage Increase (%)	60.2	32.1	40.5	31.5	53.8	34.7	55.2	40.3	58.3	62.3	30.5	55.1
7. Against a Wage Cut (%)	14.3	3.7	3.6	5.4	10.0	27.1	9.8	26.0	24.2	5.7	0.9	12.7
8. For a Change in Hours (%)	4.1	25.4	38.1	50.0	21.2	0.0	2.1	0.0	1.7	8.2	58.8	10.4
<u>Duration:</u>												
9. Mean (days)	15.2	13.1	22.9	20.8	21.7	25.8	32.7	27.1	14.0	11.6	17.6	23.0
10. Median (days)	7.5	5.0	3.0	14.0	7.0	14.0	14.0	14.5	7.0	5.0	14.0	7.0
<u>Outcomes:</u>												
11. Union Win (%)	32.7	69.8	71.4	31.5	41.2	34.7	53.8	28.6	75.0	41.0	21.2	42.4
12. Strike Breakers Employed (%)	55.1	32.5	44.0	32.6	33.1	47.2	26.6	41.6	27.5	52.5	40.3	44.9
13. All Strikers Replaced (%)	8.2	10.1	8.3	3.3	4.2	1.4	11.2	6.5	8.8	9.8	3.1	7.3
<u>Geographic Distribution:</u>												
14. Illinois (%)	11.2	8.2	16.7	57.6	45.0	100.0	4.9	1.3	12.1	42.6	49.1	33.9
15. Mass (%)	8.2	3.7	1.2	1.1	10.8	0.0	67.8	72.7	11.7	4.9	1.8	14.3
16. NY City (%)	74.4	79.5	59.5	35.9	31.9	0.0	21.7	13.0	65.4	39.3	48.7	37.6
<u>Timing:</u>												
17. In 1886 (%)	67.4	42.9	58.3	65.2	41.1	25.0	55.2	40.3	27.9	42.6	76.5	42.3
18. Jan-April 1886 (%)	39.8	10.1	17.9	6.5	13.1	6.9	18.2	28.6	9.2	23.8	7.5	13.2
19. May 1-7 1886 (%)	15.3	6.7	9.5	48.9	17.7	4.9	10.5	0.0	4.2	13.1	50.9	9.9

Notes: See note to Table 1.

Apparel industry includes clothing (men's and women's), tailors, and furriers.

Food industry includes bakeries, butchers and meat packing establishments, and confectioners.

Machine industry includes agricultural equipment, metal fabrication, rolling stock, and shipbuilding.

Metal industry includes iron and steel foundries, hardware, nail, and wire works, and brass and copper foundries.

Textile industry includes cotton, wool and silk spinning and weaving.

Transportation industry includes railroads, streetcars, and longshoring.

Wood products industry includes furniture and box making and planing mills.

Table 3

Mean Characteristics, By Cause of Strike
(standard errors in parentheses)

	Cause of Strike:			
	Wage Increase	Wage Cut	Chg in Hours	Miscell
1. Number Strikes	1026	252	399	502
2. Average Log Size	4.02	4.24	4.11	3.90
3. Average Fraction Involved (%)	78.6	80.3	89.3	77.9
4. Avg Prev Wage (\$/day)	2.01	2.06	2.27	2.24
5. Pct Authorized	74.1	67.9	93.0	73.9
<u>Timing (Percent of Strikes):</u>				
7. Jan-April 1886	17.8	6.7	6.8	15.2
8. May 1-7 1886	8.0	2.0	57.1	4.4
9. Rest of 1886	12.2	13.1	22.8	25.1
<u>Duration:</u>				
6. Mean (days)	15.7	35.1	16.9	24.2
7. Median (days)	7.0	14.0	11.0	9.0
<u>Outcomes:</u>				
8. Union Win (%)	51.9	48.0	32.6	47.5
9. Strike Breakers Employed (%)	37.0	31.7	31.3	51.7
10. All Strikers Replaced (%)	5.4	8.7	4.8	10.8
11. Avg Chg Log Wage (%)	8.36 (0.28)	-4.94 (0.44)	0.10 (0.37)	0.58 (0.33)
12. Wage Change = 0 (%)	31.2	47.6	79.0	74.7
13. Avg Chg in Hours	-0.5 (0.1)	-0.1 (0.1)	-3.8 (0.3)	-0.1 (0.1)

Note: See note to Table 1.

Table 4
Strike Outcomes by Duration and Settlement Rates by Duration
(standard errors in parentheses)

	Number Strikes	Pct Union Wins	Use of Strikebreakers: Percent Replaced		Average Daily Settlement Rates			Average Percent Chg in Log Wage
			1-25%	>25%	Union Win	Other	Total	
<u>I. Strikes Over Wage Increases:</u>								
All	1027	51.9	21.0	16.0	--	--	--	8.36 (0.28)
1-3 days	360	59.4	15.0	17.2	0.069	0.047	0.117	10.76 (0.54)
4-7 days	224	61.6	17.9	8.4	0.052	0.032	0.084	8.26 (0.54)
8-14 days	168	41.7	29.8	13.1	0.023	0.032	0.054	7.16 (0.61)
15-28 days	135	43.0	28.9	17.8	0.015	0.020	0.035	6.44 (0.71)
29-90 days	121	37.2	19.8	24.0	0.005	0.009	0.014	6.26 (0.81)
90+ days	19	42.1	47.4	42.1	--	--	--	8.77 (2.40)
<u>II. Strikes Over Wage Cuts:</u>								
All	252	48.0	12.3	19.4	--	--	--	-4.94 (0.44)
1-3 days	60	60.0	5.0	35.0	0.048	0.032	0.079	-2.08 (0.88)
4-7 days	45	55.5	17.8	8.9	0.033	0.026	0.059	-4.45 (1.00)
8-14 days	42	50.0	9.5	9.5	0.020	0.020	0.041	-4.38 (1.11)
15-28 days	35	42.9	20.0	14.3	0.010	0.014	0.024	-4.94 (0.98)
29-90 days	50	34.0	8.0	24.0	0.004	0.008	0.012	-1.00 (0.81)
90+ days	20	35.0	25.0	15.0	--	--	--	-1.73 (2.40)
<u>III. Strikes Over Hours Changes:</u>								<u>Average Chg Hours/Wk</u>
All	399	32.6	17.0	14.3	--	--	--	-3.80 (0.34)
1-3 days	114	54.4	11.4	11.4	0.052	0.043	0.095	-7.00 (0.76)
4-7 days	67	38.8	22.4	7.5	0.023	0.036	0.059	-4.77 (0.79)
8-14 days	84	20.2	15.5	13.1	0.011	0.044	0.055	-2.20 (0.52)
15-28 days	80	16.3	21.3	15.0	0.007	0.036	0.043	-1.64 (0.48)
29-90 days	46	15.2	19.6	28.3	0.002	0.012	0.014	-0.28 (0.41)
90+ days	8	62.5	12.5	37.5	--	--	--	-10.40 (5.63)

Table 5

Estimated Resistance Curves, Strikes Over Wage Issues
(standard errors in parentheses)

	Strikes Over Wage Increases:					Strikes Over Wage Cuts:				
	Change in Log Wage of Strikers:		Probability of Union Win:		Wage Chg, Union Win	Change in Log Wage of Strikers:		Probability of Union Win:		Wage Chg, Union Win
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
<u>Strike Duration Classes:</u>										
1. 2 Days	-.023 (.011)	-.027 (.010)	-.066 (.065)	-.103 (.059)	-.015 (.010)	-.021 (.027)	-.007 (.028)	-.175 (.192)	-.177 (.183)	-.005 (.006)
2. 3 Days	-.032 (.013)	-.017 (.011)	-.059 (.071)	-.042 (.065)	-.013 (.011)	-.000 (.026)	.022 (.027)	-.042 (.182)	.037 (.176)	-.013 (.005)
3. 4-5 Days	-.023 (.012)	-.023 (.011)	-.059 (.067)	-.031 (.061)	-.021 (.011)	-.017 (.026)	-.004 (.026)	-.057 (.189)	.032 (.173)	-.003 (.005)
4. 6-8 Days	-.051 (.010)	-.033 (.009)	-.140 (.056)	-.146 (.052)	-.009 (.009)	-.033 (.020)	-.025 (.022)	-.403 (.143)	-.396 (.141)	-.010 (.005)
5. 9-14 Days	-.047 (.010)	-.026 (.009)	-.307 (.055)	-.254 (.053)	-.006 (.011)	-.029 (.020)	-.024 (.021)	-.375 (.141)	-.373 (.137)	-.000 (.004)
6. 15-28 Days	-.057 (.010)	-.047 (.010)	-.291 (.058)	-.299 (.054)	-.019 (.011)	-.034 (.020)	-.020 (.022)	-.446 (.144)	-.326 (.142)	-.003 (.005)
7. 29-60 Days	-.059 (.011)	-.039 (.010)	-.340 (.063)	-.307 (.060)	.006 (.012)	-.051 (.021)	-.025 (.023)	-.616 (.151)	-.419 (.147)	-.007 (.006)
8. 61-90 Days	-.058 (.019)	-.050 (.017)	-.388 (.106)	-.368 (.098)	-.025 (.023)	-.055 (.022)	-.032 (.024)	-.440 (.155)	-.347 (.155)	-.005 (.005)
9. Over 90 Days	-.029 (.022)	-.037 (.019)	-.277 (.121)	-.340 (.111)	-.037 (.024)	-.064 (.022)	-.029 (.024)	-.525 (.160)	-.263 (.157)	-.004 (.006)
<u>Other Control Variables:</u>										
10. Authorized by Union	--	.041 (.006)	--	.237 (.036)	.018 (.008)	--	.024 (.011)	--	.091 (.072)	-.002 (.003)
11. Fraction of Workers on Strike	--	.036 (.009)	--	.133 (.053)	.040 (.011)	--	.007 (.021)	--	-.189 (.134)	-.010 (.005)
12. Log of Number of Strikers Involved	--	-.006 (.002)	--	-.014 (.010)	-.007 (.002)	--	-.003 (.004)	--	-.076 (.024)	.001 (.001)
13. Strike in First Week of May 1886	--	-.031 (.011)	--	-.099 (.061)	-.012 (.015)	--	.025 (.036)	--	.016 (.233)	-.007 (.012)
14. Strike Post- Haymarket	--	-.010 (.009)	--	-.045 (.054)	-.013 (.011)	--	-.009 (.019)	--	.042 (.125)	.002 (.005)
15. Illinois	--	-.049 (.011)	--	-.302 (.061)	-.027 (.016)	--	-.017 (.018)	--	-.208 (.115)	-.006 (.005)
16. Massachusetts	--	-.010 (.011)	--	.023 (.065)	-.003 (.012)	--	-.011 (.023)	--	-.152 (.152)	-.020 (.006)
17. Chicago	--	.037 (.012)	--	.294 (.067)	.021 (.017)	--	-.012 (.021)	--	-.203 (.138)	.003 (.006)
18. 11 Industry and 5 Year Effects	NO	YES	NO	YES	YES	NO	YES	NO	YES	YES
19. R-squared	.052	.297	.067	.275	.335	.071	.268	.129	.427	.355
<u>Sample Characteristics:</u>										
20. Mean (std dev) of Dependent Variable		.084 (.089)		.545 (.498)	.136 (.072)		-.049 (.068)		.504 (.501)	.002 (.011)
21. Sample Size		971		1049	529		230		230	116

Note: Samples exclude strikes in which all the strikers were permanently replaced. Omitted duration class is 1 day strikes. Models in columns (5) and (10) are fit to subset of strikes which ended in a union "win" -- see text.

Table 6

OLS and Competing Risk Tobit Estimates of Time to Different Strike
Outcomes for Strikes Over Wage Increases

Variable	OLS Estimates	Competing Risk Tobit Estimates of Log(Duration Times)		
	Log(Duration)	Log(Firm Capitulation Time)	Log(Union Capitulation Time)	Log(Time to a Compromise)
	(1)	(2)	(3)	(4)
Authorized by Union	.600 (.026)	.284 (.162)	.945 ^a (.141)	.804 (.248)
"1" if Strike Replacements				
Log(No. of Strikers)	.164 (.026)	.201 (.035)	.122 ^a (.032)	.078 (.066)
Fraction of Workers on Strike	-.202 (.142)	-.473 (.216)	.247 ^a (.193)	-.487 (.416)
Strike in 1st week of May, 1886	.252 (.179)	.518 (.275)	.228 (.251)	-.023 (.321)
Strike on or after 1st week of May, 1886	.377 (.144)	.453 (.201)	.208 (.175)	.157 (.263)
Illinois	-.041 (.163)	.560 (.279)	-.552 ^a (.209)	-.313 (.372)
Massachusetts	.310 (.174)	.315 (.221)	.210 (.197)	.613 (.397)
Chicago	.155 (.180)	-.434 (.297)	.514 ^a (.219)	.477 (.401)
11 Industry & 5 Year Dummies	Yes	Yes	Yes	Yes
σ	1.172	1.339 (.077)	1.080 (.051)	1.333 (.259)
ρ_{UF}			.604 (.281)	
ρ_{UC}			.435 (.521)	
ρ_{FC}			.366 (.497)	
R ²	.171			
-Log L			2254.92	

N=971

Standard errors in parentheses.

^a Difference between column (2) and Column (3) is significant at the .05 level.

Table 7

OLS and Competing Risk Tobit Estimates of Time to Different Strike Outcomes for Strikes Over Wage Increases

Variable	OLS Estimates	Competing Risk Tobit Estimates of Log(Duration Times)		
	Log(Duration)	Log(Firm Capitulation Time)	Log(Union Capitulation Time)	Log(Time to a Compromise)
	(1)	(2)	(3)	(4)
Authorized by Union	.5955 (.0935)	.1618 (.1580)	1.0534 ^a (.3195)	.6377 (.2692)
"1" if Strike Replacements	.6146 (.0825)	1.1747 (.1518)	-.1572 ^a (.1729)	.9163 (.2766)
Log(No. of Strikers)	.1453 (.0253)	.1667 (.0338)	.1122 (.0352)	.0461 (.0680)
Fraction of Workers on Strike	-.1020 (.1391)	-.3599 (.2011)	.4112 ^a (.1964)	-.5658 (.3851)
Strike in 1st week of May, 1886	.1920 (.1741)	.4521 (.2795)	.2086 (.2777)	-.1441 (.3384)
Strike on or after 1st week of May, 1886	.3612 (.1405)	.3973 (.1925)	.1837 (.1945)	.1604 (.2677)
Illinois	-.0831 (.1586)	.7146 (.2877)	-.6798 ^a (.2104)	-.0932 (.4089)
Massachusetts	.4057 (.1693)	.4740 (.2168)	.1273 (.2267)	.7716 (.4054)
Chicago	.2796 (.1758)	-.4220 (.2991)	.5799 ^a (.2254)	.4640 (.4410)
11 Industry & 5 Year Dummies	Yes	Yes	Yes	Yes
σ	1.1393	1.2817 (.0649)	1.1345 (.0639)	1.3401 (.2250)
ρ_{uf}			.4151 (.2340)	
ρ_{uc}			.1145 (.3821)	
ρ_{fc}			.4107 (.4623)	
R ²	.2170			
-Log L			2182.586	

N=971

Standard errors in parentheses.

^a Difference between column (2) and Column (3) is significant at the .05 level.

Table 8
Predicted Mean and Median Capitulation Times (Days)
And the Probability of a Union Success

Strike Description	Mean ^a		Median ^b		
	Union	Firm	Union	Firm	Pr(Union Success) ^c
Estimates without Replacements					
(a) An Average Strike ^d	36.28	23.23	20.25	9.48	.70
(b) An Unauthorized Strike	14.10	17.49	7.87	7.13	.48
(c) Strike with Log(strikers) 1 S.D. above average	44.35	32.35	24.75	13.19	.65
(d) Fraction of wkers on strike is 1 SD below average	33.42	27.18	18.65	11.08	.64
(e) Strike in Chicago	34.94	26.35	19.50	10.75	.66
(f) Strike in Ill, not Chicago	20.89	40.68	11.66	16.59	.33
Estimates with Replacements					
(g) Average Strike w/o Replacements & authorized	58.51	14.51	30.74	6.38	.77
(h) Average strike w/o Replacements & unauth.	20.41	12.34	10.72	5.43	.56
(i) Average strike w. replacements & authorized	50.00	46.97	26.27	20.66	.48
(j) Average strike with replacements & unauth.	17.44	39.96	9.16	17.57	.24

^a These values are equal to $\exp(X\beta + .5\sigma^2)$.

^b These values are equal to $\exp(X\beta)$.

^c These values are equal to $\Pr(D_u > D_f, D_c > D_f)$

^d The values for the average strike in row (a) is calculated for an authorized strike in New York, with the mean value of log(strikers), the mean value of fraction of workers on strike and the values for the year dummies and industry dummies weighted by the sample proportions.

Table 9

Semi-Parametric Estimates of the Hazard Rate For
Union and Firm Capitulation Time

Variable	Firm Capitulation Hazard (1)	Union Capitulation Hazard (2)	Firm Capitulation Hazard (3)	Union Capitulation Hazard (4)
Authorized by Union	-.0162 (.1307)	-1.2416 (.1493)	-.0381 (.1329)	-1.2889 (.1520)
"1" if Strike Replacements			-1.0459 (.1273)	.6032 (.1384)
Log(No. of Strikers)	-.1371 (.0290)	-.1276 (.0505)	-.1256 (.0290)	-.1496 (.0510)
Fraction of Workers on Strike	.5485 (.1919)	-.7642 (.2248)	.4880 (.1815)	-.7129 (.2242)
Strike in 1st week of May, 1886	-.4838 (.3013)	-.0945 (.2962)	-.3986 (.3117)	-.1393 (.3130)
Strike on or after 1st week of May, 1886	-.4870 (.1918)	-.1921 (.2366)	-.4213 (.1875)	-.1373 (.2413)
Illinois	-1.0866 (.2814)	.5639 (.1869)	-1.0055 (.2826)	.5733 (.1873)
Massachusetts	-.2608 (.2198)	-.5112 (.2833)	-.3454 (.2179)	-.4482 (.2788)
Chicago	1.0155 (.3003)	-.4477 (.2467)	.8546 (.3014)	-.2943 (.2444)
11 Industry & 5 Year Dummies	Yes	Yes	Yes	Yes
-Log L	1568.25	945.63	1523.68	935.95

Standard errors in parentheses.

N=971

The estimates of the baseline hazard rates are shown in Figures 5a and 5b.

Table 10

Estimated Impact of Exogenous Variables on
Capitulation Time Using the
Semi-Parametric Estimates

Variable	E(Ln(cap. time) Ln(cap. time) ≤ Ln(21 Days))						Pr(Ln(Cap. Time) > Ln(21 Days))					
	Firm			Union			Firm			Union		
	(1) NP Est.	(2) Indep. Tobit	(3) Correl. Tobit	(4) NP Est.	(5) Indep. Tobit	(6) Correl. Tobit	(7) NP Est.	(8) Indep. Tobit	(9) Correl. Tobit	(10) NP Est	(11) Indep. Tobit	(12) Correl Tobit
(a) "Average Strike"	1.528	1.562	1.468	1.933	1.926	1.838	.275	.265	.182	.397	.384	.272
The data below show the change in the capitulation times relative to the "average strike" given a change in the dummy variable shown in the column heading.												
(b) Authorized by Union	.013	-.001	.095	.143	.403	.421	.014	.000	.036	.384	.382	.354
(c) "1" if Strike Replacements	.226	.539	.559	-.172	-.185	-.084	.372	.373	.322	-.222	-.114	-.044
(d) Strike in 1st week of May, 1886	.101	.260	.242	.001	.078	.004	.130	.146	.104	.002	.060	.002
(e) Strike on or after 1st week of May, 1886	.107	.203	.214	.001	.041	-.009	.138	.109	.090	.001	.031	-.005
(f) Illinois (not Chicago)	.213	.441	.373	-.215	-.465	-.410	.360	.285	.182	-.213	-.228	-.158
(g) Massachusetts	.103	.241	.260	.073	.022	.064	.133	.134	.114	.162	.017	.038
(h) Chicago	.050	.161	.167	-.066	-.104	-.053	.058	.084	.067	-.106	-.068	-.029
(i) $\partial E(D D \leq 21)/\partial(\log \text{stkr})$	.045	.090	.100	.030	.040	.058						
(j) $\partial \text{Pr}(D > 21)/\partial(\log \text{stkr})$							.048	.041	.034	.057	.033	.033
(k) $\partial E(D D \leq 21)/\partial(\text{fraction on Strike})$	-.174	-.200	-.217	.145	.230	.212						
(l) $\partial \text{Pr}(D > 21)/\partial(\text{fraction on Strike})$							-.185	-.094	-.074	.272	.170	.120

An average strike sets all of the variables to zero and log(no of Strikers) and Fraction on strike to the sample means. This corresponds to a strike in New York, in 1881, in the omitted industry grouping that was unauthorized and replacements were not used.

The values in columns (2), (3), (5) and (6) equal $X\beta_i$, $\hat{\epsilon}_i + E(\epsilon_i | X\beta_i, \text{hat} + \epsilon_i \leq 21)$ where i refers to either the firm or union. The values in columns (8), (9), (11) and (12) equal $\text{Pr}(X\beta_i, \text{hat} + \epsilon_i > 21)$. The parameter estimates for the calculations in columns (3), (6), (9) and (12) come from Table 7. The parameters estimates used in columns (2), (5), (8) and (11) are from tobit estimates that assume the errors are uncorrelated. These parameter estimates are available from the authors.

Using the semi-parametric estimates, the values in columns (1) and (4) equal $\Sigma[\ln(D_i) \text{Pr}(\text{Duration} = D_i)] / \text{Pr}(\text{Duration} \leq 21)$ where $i = 1, 2, \dots, 21$. In this calculation $\text{Pr}(\text{Duration} = 0_i)$ is calculated using Equation (x) in the text.

Table 11

Estimates of Change in Log(wage) When the Union Wins
Strikes Over Wage Increases

Variable	$\Delta \ln W$ OLS Estimates	$\Delta \ln W$ MLE Estimates	$(\beta_u - \beta_f)$
	(1)	(2)	(3)
Authorized by Union	.0184 (.0080)	.0367 (.0089)	.6611 (.2190)
"1" if Strike Replacements			
Log(No. of Strikers)	-.0069 (.0018)	-.0083 (.0023)	-.0792 (.0400)
Fraction of Workers on Strike	.0380 (.0108)	.0534 (.0132)	.7205 (.2691)
Strike in 1st week of May, 1886	-.0026 (.0162)	-.0126 (.0189)	-.2902 (.2566)
Strike on or after 1st week of May, 1886	-.0145 (.0113)	-.0176 (.0111)	-.2442 (.2129)
Illinois	-.0252 (.0163)	-.0659 (.0149)	-1.1119 (.3602)
Massachusetts	-.0048 (.0121)	-.0044 (.0160)	-.1044 (.2291)
Chicago	.0190 (.0172)	.0595 (.0170)	.9482 (.3396)
11 Industry & 5 Year Dummies	Yes	Yes	Yes
σ_w	.0611	.0700 (.0029)	
ρ_{uw}		.2837 (.1742)	
ρ_{fw}		-.2365 (.0607)	
ρ_{uf}		.7971 (.2834)	
R^2	.3209		
-Log L		1138.51	
Sample Size	529	971	

Standard errors in parentheses.

The results shown in column 4 are based the estimates reported in Table 6

Table 12

Estimates of Change in Log(wage) When the Union Wins
Strikes Over Wage Increases

Variable	$\Delta \ln W$ OLS Estimates (1)	$\Delta \ln W$ MLE Estimates (2)	$(\beta_u - \beta_f)$ (3)
Authorized by Union	.0184 (.0081)	.0341 (.0091)	.8916 (.2213)
"1" if Strike Replacements	.0000 (.0073)	-.0242 (.0098)	-1.3320 (.2653)
Log(No. of Strikers)	-.0069 (.0018)	-.0073 (.0023)	-.0544 (.0430)
Fraction of Workers on Strike	.0380 (.0109)	.0456 (.0134)	.7711 (.2628)
Strike in 1st week of May, 1886	-.0026 (.0164)	-.0083 (.0192)	-.2435 (.3150)
Strike on or after 1st week of May, 1886	-.0145 (.0114)	-.0165 (.0105)	-.2136 (.2431)
Illinois	-.0252 (.0164)	-.0574 (.0155)	-1.3944 (.3689)
Massachusetts	-.0048 (.0122)	-.0088 (.0159)	-.3467 (.2818)
Chicago	.0190 (.0172)	.0481 (.0172)	1.0019 (.3530)
11 Industry & 5 Year Dummies	Yes	Yes	
σ_w	.0612	.0662 (.0026)	
ρ_{uw}		.4158 (.1597)	
ρ_{fw}		-.2031 (.0633)	
ρ_{uf}		.6326 (.2371)	
R^2	.3209		
-Log L		1078.22	
Sample Size	529	971	

Standard errors in parentheses.

The estimates shown in the third column are from Table 7.

Table 13

Industry Effects on Time to Union and Firm Capitulation
And Union Wage Demands

Industry	Time to Union or Firm Capitulation			Union Wage Demand Relative to the Misc. Industries
	β_u (1)	β_f (2)	$\beta_u - \beta_f$ (3)	(4) ^a
Apparel	-.4811 (.2569)	.2062 (.3283)	-.6873 (.3177)	-.0498 (.0153)
Building Trades	-.3172 (.2717)	-1.2155 (.2308)	.8983 (.3356)	.0569 (.0119)
Food	.0422 (.2662)	-.3854 (.2632)	.4276 (.3474)	.0551 (.0175)
Machines	.2868 (.2848)	-.0267 (.2883)	.3136 (.3139)	-.0133 (.0333)
Metals	-.1450 (.1610)	-.4917 (.1785)	.3467 (.1920)	.0081 (.0128)
Mining	.5631 (.2429)	.2319 (.3301)	.3311 (.3264)	.0177 (.0233)
Shoes	.4183 (.2576)	-.3579 (.2337)	.7762 (.3178)	.0503 (.0120)
Textile	.1048 (.2515)	-.1095 (.3418)	.2144 (.3048)	.0159 (.0225)
Tobacco	.2247 (.2415)	-.6780 (.2023)	.9027 (.2968)	-.0359 (.0145)
Transportation	-.3837 (.2182)	-.5762 (.2060)	.1925 (.2291)	.0339 (.0126)
Wood	-.5213 (.2522)	.0528 (.2451)	-.5741 (.2868)	-.0707 (.0203)
Other Industries			-.4367 (.2854)	-----
With Replacement Dummy Variable	No	No	No	No

^a Industry parameters for the results shown in column (2), Table 11.

Table 14

Industry Effects on Time to Union and Firm Capitulation
And Union Wage Demands

Industry	Time to Union or Firm Capitulation			Union Wage Demand Relative To Misc. Industries
	β_u (1)	β_f (2)	$\beta_u - \beta_f$ (3)	(4) ^a
Apparel	-.5568 (.2594)	.0716 (.3093)	-.6284 (.3233)	-.0394 (.0152)
Building Trades	-.1542 (.2745)	-1.0302 (.2191)	.8760 (.3216)	.0483 (.0115)
Food	.1726 (.3064)	-.3353 (.2470)	.5079 (.3935)	.0531 (.0158)
Machines	.3744 (.3123)	.0396 (.2952)	.3348 (.3676)	-.0218 (.0362)
Metals	-.0698 (.1757)	-.4451 (.1777)	.3753 (.2185)	.0044 (.0127)
Mining	.5591 (.2701)	.1661 (.3446)	.3930 (.3866)	.0132 (.0241)
Shoes	.6082 (.2562)	-.0758 (.2204)	.6839 (.3093)	.0406 (.0118)
Textile	.1986 (.2710)	-.0612 (.3526)	.2598 (.3527)	.0135 (.0227)
Tobacco	.3828 (.2327)	-.4167 (.1899)	.7995 (.2667)	-.0464 (.0139)
Transportation	-.3336 (.2429)	-.6714 (.2065)	.3378 (.2807)	.0363 (.0124)
Wood	-.5328 (.2796)	.0417 (.2440)	-.5745 (.3280)	-.0671 (.0208)
Other Industries			.0823 (.3335)	---
With Replacement Dummy Variable	Yes	Yes	Yes	Yes

Industry parameters are for the model reported in column (2), Table 12.