STRIKES AND WAGES: A TEST OF AN ASYMMETRIC INFORMATION MODEL*

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This paper describes a simple model of labor disputes based on the hypothesis that unions use strikes to infer the profitability of the firm. The model posits the existence of a negatively sloped resistance curve between wages and strike duration. In addition, it offer a series of predictions relating wage and strike outcomes to changes in the expected profitability of the firm and changes in the alternative opportunities of striking workers. These implications are tested using data on wage outcomes, strike probabilities, and strike durations for a large sample of collective bargaining agreements.

I. INTRODUCTION

It has been long recognized that any consistent economic model of strikes must appeal to some form of imperfect information.[1] Recently, a great deal of progress in the theoretical analysis of disputes has been made by focusing on a particularly simple case: that of one-sided asymmetric information over the size of the bargaining surplus.[2] In the application of this framework to strikes and lockouts, it is usually assumed that the profitability of the firm is unknown to union members. A strike is then viewed as a screening device that allows workers to extract higher wages from more profitable employers. In addition to providing a simple explanation for the existence of disputes, this class of models yields a rich set of empirical implications for observed wage settlements and the probability and duration of strikes.

This paper is an attempt to test these implications using wage

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1. The role of imperfect information in generating disputes was emphasized by Hicks [1964] and many subsequent authors, including Ashenfelter and Johnson [1969]. Kennan [1986] provides a brief summary of the historical development of theoretical models of strikes.

2. See, in particular, Hayes [1984], Morton [1983], Sobel and Takahashi [1983], Fudenberg and Tirole [1983], Kennan and Wilson [1988], and Hart [1989]. Cramp-ton [1984] considers the case of two-sided asymmetric information. Recent applied studies motivated by this class of models include Tracy [1987]; Fudenberg, Levine, and Ruud [1985]; and McConnell [1987, 1989]. Farber [1978a] presents an earlier microeconometric study of strikes and wages based on Ashenfelter and Johnson's [1969] model.

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and strike outcomes for a sample of collective bargaining agreements from the Canadian manufacturing sector. In the first part of the paper, a simple theoretical model is presented that describes negotiated wage rates and strike outcomes in terms of a small set of underlying variables: the mean and dispersion of the unobserved component of profitability, the opportunities available to striking workers, and the risk preferences of union members. The model predicts the existence of a simple resistance curve relating lower wage settlements to longer strikes. The model also describes the effects of the predetermined variables on wage settlements, strike probabilities, and strike durations.

In the second part of the paper, these implications are tested against the contract data. Industry-specific input and output prices are used to form a measure of the expected profitability of the firm, while regional unemployment rates and wages are used to proxy the earnings opportunities of striking workers. Models are fit for negotiated wage rates, the probability of disputes, and the conditional duration of work stoppages. The econometric models abstract from any permanent differences across firm and union bargaining pairs.

The empirical analysis provides only limited support for the class of one-sided asymmetric information strike models. Very long strikes are associated with lower wage settlements. The trade-off between negotiated wages and shorter strikes, however, is generally positive. And while the effects of unemployment on wages and strike outcomes are consistent with the model, the effects of industry-specific prices are not. These findings suggest that a richer model may be needed to capture all the features of the wage and strike outcomes in this data set.

II. A Simple Model of Strikes

This section outlines a model of disputes based on the hypothesis that unions use strikes as a mechanism to price discriminate against more profitable employers. For simplicity, bargaining power is vested in the union. With complete information the union is assumed to be able to capture all the bargaining surplus. When profits are not directly observable, however, the union cannot rely on the firm to reveal its private information. Nevertheless, the union may be able to improve upon the strategy of a fixed wage demand by offering the firm a downward sloping wage-strike schedule. Faced with such a resistance curve, the firm chooses a shorter strike and a higher wage in high-profit states, and a longer strike and a lower wage in low-profit states. Strikes therefore enable the union to distinguish between more and less profitable employers on the basis of their willingness to endure delays in production.

Following Hayes [1984], I assume that the union can commit itself to an arbitrary resistance schedule. Most of the recent literature on bargaining disputes has adopted an alternative assumption of limited commitment. In the sequential bargaining models of Sobel and Takahashi [1983], Fudenberg and Tirole [1983], and Hart [1989], bargaining consists of a series of offers made by the union that the firm can either accept or reject. If an offer is rejected, the strike continues, and a new offer is presented after the passage of a fixed interval of time.[3] As Hart [1989] has pointed out, even with one to three days between offers these models have considerable difficulty explaining the existence of strikes longer than a few weeks. Hart reconciles the existence of longer disputes by assuming that the costs of continuing a strike increase discretely after two to three months.

The assumption of full commitment eliminates the difficulty of "too short" strikes, and also abstracts from the mechanical features of the bargaining process, such as the time between offers or whether the firm can make counteroffers. Full commitment presents its own difficulties, however. Foremost among these is the observation that if a union can commit to an arbitrary wage-strike schedule, then the schedule that maximizes the expected income of the union is a take-it-or-leave-it wage demand with the threat of an indefinite strike (see below). Since the median duration of strikes in North American contract negotiations is on the order of a month, and most strikes are settled within a year, there is strong evidence against the joint hypotheses of full commitment and expected income maximization. Sequential bargaining models generate shorter strikes by assuming that the union can only delay the bargaining process for a short time, and that after each delay a new offer is presented. An alternative assumption pursued here is that the union acts in a risk-averse manner in the formation of its full commitment resistance curve. This leads to a very simple formulation in which the duration of strikes is related to the risk aversion of union members, rather than to the time between offers. It is an open question whether there are testable differences between the implica-

3. Some noninfinitesimal delay is necessary to generate strikes, see Hart [1989] for a discussion of the limiting case.

tions of a full commitment model with risk aversion, on the one hand, and a sequential bargaining model with more limited commitment, on the other.

A. Description of the Model

The driving variable in the model is the net value of output per worker, θ, which is assumed to be a random variable whose realization is known to the firm but unknown to workers. For technical convenience θ is assumed to be uniformly distributed on the interval $[\theta_1, \theta_2]$.[4] Bargaining involves the determination of a wage rate w and a strike length s. It is convenient to model s as the fraction of an exogenous contract period lost to a work stoppage.[5] During a strike workers earn a constant alternative wage a, while the firm produces nothing. Thereafter, each worker produces θ and is paid w. The firm's profits per worker, given θ, w, and s, are therefore proportional to

$$(\theta - w)(1 - s), \tag{1}$$

while the total earnings of a worker are proportional to

$$r \equiv (1 - s)\, w + s\, a. \tag{2}$$

I assume that the alternative wage is less than θ_1, the lower bound of θ, so that the joint surplus from any bargain $(\theta - a)$ is strictly positive. Workers are assumed to evaluate a particular distribution of earnings according to the expected value of $u(r)$, where u is a constant absolute risk aversion utility function with risk parameter R.

The union's problem can be thought of as one of choosing a wage-strike schedule $w(s)$ that maximizes $E(u(r))$ subject to the constraint that for any given schedule $w(s)$ and any realization of productivity the firm will choose a profit-maximizing strike length. The union is assumed to have information on the value of a and on the parameters of the distribution of θ. Variation over time in wage and strike outcomes therefore reflects predictable variation due to changes in these parameters, and unpredictable variation due to the specific realization of θ.

4. For much of the theoretical analysis it is sufficient to have θ distributed on a closed interval with a strictly positive density and an increasing hazard function.

5. If a contract is to run from the present (period 0) to some period T in the future, and if a work stoppage lasts S periods, then the fraction of time lost is $(1 - \exp(-\delta S))/(1 - \exp(-\delta T))$, where δ is an appropriate discount rate. For values of S and T such that S/T is relatively small (say less than 1/3), this fraction is approximately S/T.

Analytically, it is more convenient to express the union's problem as one of choosing a pair of functions $w(\theta)$ and $s(\theta)$ to maximize the expected utility of union receipts, subject to the incentive compatibility constraint that the firm is willing to declare θ truthfully, and subject to the individual rationality constraint that profits are large enough in every state to induce the firm to participate in an agreement. Let $\Pi(\tilde{\theta},\theta)$ denote the profits of the firm when productivity is θ, and it declares a level of productivity $\tilde{\theta}$, and let $\Pi(\theta) = \Pi(\theta,\theta)$. Then

$$\begin{aligned}\Pi(\tilde{\theta},\theta) &= (1 - s(\tilde{\theta}))\,(\theta - w(\tilde{\theta}))\\ &= \Pi(\tilde{\theta}) + (1 - s(\tilde{\theta}))\,(\theta - \tilde{\theta}).\end{aligned}$$

Incentive compatibility requires that

$$\Pi(\theta) \geq \Pi(\tilde{\theta},\theta) = \Pi(\tilde{\theta}) + (1 - s(\tilde{\theta}))\,(\theta - \tilde{\theta}).$$

Reversing the roles of θ and $\tilde{\theta}$ leads to a conformable expression that may be combined with this one to yield

$$(1 - s(\theta))\,(\theta - \tilde{\theta}) \geq \Pi(\theta) - \Pi(\tilde{\theta}) \geq (1 - s(\tilde{\theta}))(\theta - \tilde{\theta}). \tag{3}$$

Since strike length is bounded between 0 and 1, equation (3) implies that $\Pi(\theta)$ is nondecreasing in θ. A comparison of the right- and left-hand expressions of (3) also shows that $s(\theta)$ is nonincreasing in θ. Furthermore, (3) implies that $\Pi(\theta)$ is convex, and therefore continuously differentiable almost everywhere, with derivative $\Pi'(\theta) = (1 - s(\theta))$.

Assuming that the firm can earn zero profits simply by closing down, the individual rationality constraint is $\Pi(\theta) \geq 0$ for all θ. Since Π is nondecreasing, this condition is satisfied if and only if $\Pi(\theta_1) \geq 0$. Thus, necessary conditions for incentive compatibility and individual rationality are $\Pi(\theta_1) \geq 0$, $s(\theta)$ nonincreasing (and between 0 and 1) and

$$\Pi(\theta) = \int_{\theta_1}^{\theta} (1 - s(t))\, dt. \tag{4}$$

It is straightforward to show that these three conditions are also sufficient for incentive compatibility and individual rationality.

If θ and $\tilde{\theta}$ are two points with $\theta > \tilde{\theta}$ and $1 > s(\tilde{\theta}) > s(\theta)$, then equation (3) implies that

$$-\frac{\theta - w(\theta)}{1 - s(\tilde{\theta})} \leq \frac{w(\theta) - w(\tilde{\theta})}{s(\theta) - s(\tilde{\theta})} \leq -\frac{\tilde{\theta} - w(\tilde{\theta})}{1 - s(\theta)}.$$

Thus, incentive compatibility and assumption of positive profits in each state imply that observed wage-strike combinations lie on a negatively sloped resistance curve.

These conditions also imply a bound on the maximum wage advantage associated with a given strike mechanism. For any incentive-compatible wage and strike functions satisfying $\Pi(\theta_1) = 0$, let θ^+ represent the lowest value of θ such that $s(\theta) = 0$. The probability of a strike is then prob $(\theta \leq \theta^+)$. Incentive compatibility requires that the firm pay the same "no-strike" wage $w(\theta^+)$ for all $\theta \geq \theta^+$. The wage that the union can achieve without any threat of strikes is θ_1. The maximum wage advantage associated with the pair of functions $w(\theta)$, $s(\theta)$ is therefore $w(\theta^+) - \theta_1 = \theta^+ - \Pi(\theta^+) - \theta_1$. Convexity of the profit function implies that

$$\Pi(\theta^+) \geq \Pi(\theta_1) + (\theta^+ - \theta_1)\,\Pi'(\theta_1) = (\theta^+ - \theta_1)\,(1 - s(\theta_1)).$$

Therefore,

$$w(\theta^+) - \theta_1 \leq \theta^+ - \theta_1 - (\theta^+ - \theta_1)\,(1 - s(\theta_1)),$$

or

$$(w(\theta^+) - \theta_1)/\theta_1 \leq s(\theta_1)\,((\theta^+ - \theta_1)/\theta_1).$$

The union can capture at most a fraction $s(\theta_1)$ of the productivity differential between θ_1 (the lowest level of productivity) and θ^+ (the pth percentile of productivity, where p is the probability of a dispute).

The problem of maximizing the expected utility of union members subject to incentive compatibility and individual rationality is equivalent to

$$\max_{s(\theta)} \int_{\theta_1}^{\theta_2} u(\theta(1 - s(\theta)) - \Pi(\theta) + s(\theta)a)f(\theta)\,d\theta$$

subject to

$$\begin{aligned}
&\Pi(\theta_1) \geq 0,\\
&1 \geq s(\theta) \geq 0,\\
&\Pi'(\theta) = 1 - s(\theta), \text{ and}\\
&s(\theta) \text{ nonincreasing,}
\end{aligned}$$

where $f(\theta) = 1/(\theta_2 - \theta_1)$ denotes the density function of θ.

Provided that the monotonicity constraint on s is never binding, this problem can be solved by conventional optimal control techniques, treating s as the control variable and Π as the state variable.[6] I shall proceed under this assumption and then show that this constraint is, in fact, not binding. The Hamiltonian function is

$$H(\Pi,s,\theta) = u(\theta(1 - s) - \Pi + sa)\, f(\theta) + \mu(\theta)\,(1 - s(\theta)),$$

where μ is the costate variable. The necessary conditions for an optimum are

$$\frac{\partial H}{\partial s} = (a - \theta)u'(r(\theta))\, f(\theta) - \mu(\theta) = 0 \qquad (0 < s < 1), \tag{5a}$$

$$\frac{\partial H}{\partial \Pi} = -\mu'(\theta) = -u'(r(\theta))\, f(\theta), \tag{5b}$$

and

$$\mu(\theta_2) = 0. \tag{5c}$$

The Hamiltonian is concave in s if u is a concave function, or equivalently, if the index of absolute risk aversion is positive.

Using (5b) and (5c), the value of the costate variable can be written as

$$\mu(\theta) = -\int_{\theta}^{\theta_2} u'\,(r(t))\, f(t)\, dt.$$

Substituting this expression into (5a), the first-order condition for an interior strike length can be written as

$$(\theta - a)\, u'(r(\theta))\, f(\theta) = \int_{\theta}^{\theta_2} u'\,(r(t))\, f(t)\, dt. \tag{6}$$

Notice that if u' is constant (i.e., workers are risk neutral), then this expression is independent of s and implies that $(\theta - a)\, h(\theta) = 1$, where $h(\theta) \equiv f(\theta)/(1 - F(\theta))$ is the hazard function of θ. In the risk-neutral case the union makes a single take-it-or-leave-it wage demand. If the solution to equation (6), say θ^+, is less than θ_1, then the union demands θ_1, and there are no strikes. Otherwise, the wage demand is $\theta^+ > \theta_1$, which is accepted by the firm if $\theta > \theta^+$, and rejected if $\theta < \theta^+$, resulting in a strike of length 1.

Under the assumption that θ is uniformly distributed, the expression for the critical value θ^+ has the simple form $\theta^+ = (\theta_2 + a)/2$. Let $m = (\theta_1 + \theta_2)/2$ represent the mean of the distribu-

6. In the case of a risk-neutral union, the problem of choosing $s(\theta)$ can be easily converted into a point-wise optimization problem by substituting for $\Pi(\theta)$ from equation (4) and changing the order of integration.

tion of net profits per worker; let $g = (m - a)/m$ represent the percentage gap between expected productivity at the firm and the wage earned during a strike, and let $d = (\theta_2 - \theta_1)/2m$ index the dispersion of θ.[7] The condition for the occurrence of strikes can then be written as $d > g/3$, which is more likely, the greater the dispersion in potential profitability and the smaller the expected bargaining surplus. Assuming that this condition is met, the probability of a work stoppage is just the probability that $\theta < \theta^+$, which is $3/4 - g/4d$. For example, if $g = 0.4$, then the index of dispersion in profitability must be 0.167 to generate a 15 percent probability of strikes. This value implies a coefficient of variation of net productivity of approximately 10 percent.

In contrast to the risk-neutral case, in the risk-averse case the first-order condition (6) is a function of s. Nonetheless, the critical value of productivity that distinguishes the strike and no-strike states is the same as in the risk-neutral case. In particular, if θ^+ satisfies equation (6) in the risk-neutral case, then $s(\theta^+) = 0$ in the risk-averse case. To see that this is true, observe that if $s(\theta^+) = 0$, then $s(\theta) = 0$ for all $\theta \geq \theta^+$: thus, union receipts are fixed, and $u'(r)$ is constant for all $\theta > \theta^+$. It follows that $s(\theta^+) = 0$ at $\theta^+ = (\theta_2 + a)/2$ is a solution to (6) for any value of the risk aversion parameter R. Thus, the probability of strikes does not depend on the degree of risk aversion.[8]

In the risk-averse case, however, there is an interval of realizations of productivity below θ^+, say (θ^*,θ^+), in which strike length is strictly positive and less than unity. For any θ in this interval, the derivative of strike length with respect to θ may be obtained by differentiating the first-order condition (6) with respect to θ. The result can be written as

$$s'(\theta) = -\,2/(R(\theta - a)^2),$$

yielding the solution,

$$\begin{aligned} s(\theta) &= 2/(R(\theta - a)) - 4/(R(\theta_2 - a), && \theta^+ \geq \theta \geq \theta^*, \\ &= 0, && \theta > \theta^+, \\ &= 1, && \theta < \theta^*, \end{aligned} \tag{7}$$

where $\theta^* = \max\,[\theta_1, a + 2(\theta_2 - a)/(R(\theta_2 - a) + 4)]$. This solution clearly satisfies the (ignored) monotonicity constraint on $s(\theta)$ and is therefore a solution to the fully specified optimization problem.

7. The coefficient of variation of θ is equal to $3^{-1/2}d$. The condition $a < \theta_1$ implies that $g > d$.

8. This conclusion does not depend on the assumption of a uniform distribution for θ.

As an empirical matter, strike durations rarely exceed one year in negotiations for contracts over a three-year term (see Section II below). In the context of the model this suggests that workers are significantly risk averse. Maximum strike length is $s(\theta_1)$, which can be written as

$$s(\theta_1) = \frac{2}{R\,\theta_1}\frac{p}{1-p}\frac{1}{g_1},$$

where p is the probability of a dispute and $g_1 = (\theta_1 - a)/\theta_1$ is the minimum proportional difference between the alternative wage and productivity in the firm.[9] The term $R\ \theta_1$ in this expression has the interpretation of the index of relative risk aversion evaluated at the lowest possible wage rate. If the probability of disputes is 15 percent and $g_1 = 0.3$, then the index of relative risk aversion must be at least 3.5 to ensure strike durations of less than one third. This is within the range of estimates of the index of relative risk aversion obtained by Farber [1978b] in his study of wages and pensions in the coal mining industry, and only slightly higher than the estimates in Card [1986] based on the indexation provisions of labor contracts in the Canadian manufacturing sector.[10]

Provided that maximum strike duration is less than unity, the conditional mean duration of strikes can be written as

$$\frac{4}{R\,m}\left\{\frac{1}{(3d-g)}\log\left[\frac{d+g}{2(g-d)}\right]-\frac{1}{g+d}\right\}.$$

For example, if $g = 0.4$, $d = 0.167$, and $R\ m = 4$, then expected duration of a strike (conditional on one occurring) is 0.18, which translates roughly to a 120-day stoppage in negotiations over a two-year contract. The conditional mean duration of strikes can be shown to be increasing in the measure of the dispersion in unobserved productivity (d) and decreasing in the measure of the expected bargaining surplus (g). Since the probability of disputes is also increasing in d and decreasing in g, the model implies that changes in the distribution of profitability and changes in the alternative wage shift the probability and conditional duration of strikes in the same direction.

An expression for the profits of the firm can be obtained from equation (7) using the incentive-compatibility constraint $\Pi'(\theta) = 1 - s(\theta)$ and the initial condition $\Pi(\theta_1) = 0$. This expression can

9. In terms of the parameters g and d, $g_1 - (g - d)/(1 - d)$.
10. Farber [1978a, Table 1] reports estimates of 3.0 and 3.7, while Card [1986, Tables 2, 3, 4] reports estimates ranging from 1.9 to 2.8.

then be combined with equation (1) to calculate $w(\theta)$. Assuming that the maximum strike length is less than 1, the wage in the absence of a strike can be written as

$$w(\theta^{+}) = \theta_1 \left\{ 1 + \frac{2}{R\theta_1} \log \left[\frac{g+d}{2(g-d)} \right] - \frac{2}{R\theta_1} \left[\frac{3d-g}{g+d} \right] \right\}.$$

This wage is approximately unit-elastic with respect to the mean of productivity, and is increasing in both the dispersion of unobserved productivity and the level of alternative wages. The predicted elasticity of the no-strike wage with respect to the alternative wage, however, is relatively small. For example, if the probability of strikes is 15 percent and the gap between the mean of productivity and the level of earnings for striking workers is 40 percent, the predicted elasticity of $w(\theta^{+})$ with respect to the alternative wage is 0.15.[11]

The wage resistance curve may be obtained by combining the expressions for $w(\theta)$ and $s(\theta)$ and making use of the fact that $w(\theta_1) = \theta_1$. A useful summary of the resistance curve is the maximum wage concession, which is the difference between the lowest possible wage $w(\theta_1)$ and the no-strike wage $w(\theta^{+})$:

$$\frac{w(\theta^{+}) - w(\theta_1)}{w(\theta_1)} = \frac{2}{R\theta_1} \left\{ \log \left[\frac{g+d}{2(g-d)} \right] - \left[\frac{3d-g}{g+d} \right] \right\}.$$

This difference can be shown to be increasing in d, decreasing in g, and decreasing in R. To get some idea of the magnitude of the wage concessions implied by the model, if $d = 0.167$ and $g = 0.4$ (implying a strike probability of 0.15), and if the index of relative risk aversion at $w = m$ is 4 (implying a maximum strike duration of 0.37 and a mean duration of 0.18), then the gap between the no-strike wage and the wage after the longest possible strike is 1 percent. By comparison, if workers were risk neutral, they would demand a wage 6 percent above the minimum realization of productivity and strike for the full contract period otherwise.

To summarize, this model of strikes identifies two important determinants of strike incidence and duration: the dispersion in the unobserved component of productivity, and the gap between expected productivity at the firm and the wage available to union members during a work stoppage. Increases in dispersion of the unobserved component of productivity increase the probability and

11. The elasticity is $s(\theta_1) \cdot 2d(1-g)/((1-d)(g+d))$, where $s(\theta_1)$ is the length of the longest strike.

duration of strikes. As predicted by simple "joint-cost" models of strikes [Kennan, 1980], increases in the expected bargaining surplus, generated by either increases in the mean level of productivity at the firm or decreases in the alternative wage, reduce both the probability and duration of strikes.

The model implies that negotiated wage rates depend on the same set of exogenous variables. Wages are predicted to increase with the mean and dispersion of profitability and increase with the alternative wage. Furthermore, holding constant these variables, wages are a decreasing function of strike duration. Given the range of empirical estimates of mean and maximum strike duration, however, the model suggests that the gap between wage settlements reached with and without a work stoppage may be relatively modest.

III. Empirical Analysis

This section presents a series of evidence on wage settlements and strike activity based on information from a sample of collective bargaining agreements in the Canadian manufacturing sector. I first investigate whether or not negotiated wage rates are systematically linked to measures of strike intensity. The existence of a negatively sloped resistance curve is the most direct implication of the hypothesis that strikes arise from one-sided asymmetric information on profitability. I then compare the effects of changes in expected profitability and changes in the outside opportunities of striking workers on negotiated wage outcomes, strike probabilities, and strike durations. The simple model presented in the previous section yields a readily testable set of predictions on the effects of these variables.

A. Sample Description

The empirical results are based on a sample of union contracts negotiated in Canada between 1964 and 1985.[12] The sample is derived from a file of wage settlements made available by Labour

12. Canadian laws governing strikes and lockouts during contract negotiations are similar to those in the United States, although in some provinces the parties are required to pass through one or two stages of compulsory mediation before declaring a work stoppage. Several jurisdictions also require a mandatory strike vote (union-supervised or government-supervised) prior to a strike. See Sufrin [1970]. A cross-sectional analysis of the effects of these provisions on the incidence of disputes is presented by Gunderson, Kervin, and Reid [1986]. Gunderson and Melino [1988] present a parallel analysis of the duration of work stoppages.

Canada, and includes contracts involving more than 500 employees in 19 two-digit manufacturing industries.[13] I have restricted attention to the set of agreements written by bargaining pairs with four or more consecutive contracts in the file. The resulting sample contains 2,258 contracts negotiated by 299 different firm-union pairs. More information on the derivation of the sample is presented in Appendix 1.

The Labour Canada file reports the effective date and expiration date of each contract, and monthly wage rates throughout the contract. The file also reports the occurrence of a strike, but gives no information on the duration of any work stoppage. Beginning and ending dates of all labor disputes involving more than 100 workers are published annually in *Strikes and Lockouts in Canada.* I have therefore merged duration information from this publication into the contract sample.[14] I have also merged a variety of other aggregate, regional, and industry-specific information into the contract sample, including the Consumer Price Index (CPI), regional wage rates and unemployment rates, and industry-specific intermediate input prices and selling prices. Sources for these data are presented in Appendix 1.

Table I provides a cross-sectional overview of the contract sample. The table shows the number of distinct firm-union pairs in each two-digit industry, the total number of contracts from each industry, and the average contract duration, contractual wage rate, and strike probability in each industry.[15] Two measures of strike duration are also presented in the table: the median duration of disputes, and the ratio of the mean duration of strikes in the industry to the mean duration of contracts in the industry. The latter provides a rough measure of the fraction of the contract period lost to a work stoppage.

The sample is "unbalanced" in the sense that a given bargaining pair may have as few as four consecutive contracts in the sample,

13. I am grateful to Labour Canada for supplying these data. Previously, Lacroix [1986] and Gunderson, Kervin, and Reid [1986] have used the Labour Canada file to study strikes. Riddell [1980] analyses wage and strike outcome in an earlier file of Canadian contracts.

14. The listing of disputes in *Strikes and Lockouts in Canada* also provides a check on the accuracy of the strike information in the Labour Canada file, which turns out to be incorrect in about 2 percent of cases. See Appendix 1.

15. The wage measure is a geometric average of the monthly real wage rates throughout the contract, sampled at six-month intervals.

TABLE I
CHARACTERISTICS OF NEGOTIATED WAGE RATES AND MEASURES OF STRIKE ACTIVITY BY INDUSTRY

	Number of bargaining pairs	Number of contracts	Average contract length (months)	Average real wage rate during contract	Strike probability (percent)	Median strike duration (days)	Mean days lost as percent of mean contract duration
1. Food and beverages	37	321	23.6	7.88	15.6	39	6.2
2. Tobacco	5	38	22.9	8.51	5.3	29	4.1
3. Rubber products	11	61	33.8	7.25	26.2	19	6.2
4. Leather products	4	28	28.2	4.81	14.3	43	3.5
5. Textiles	14	103	27.8	6.22	29.1	30	5.9
6. Apparel	17	134	27.7	4.91	7.5	11	1.2
7. Wood products	6	47	23.8	8.91	34.0	45	6.9
8. Furniture	3	21	20.2	6.66	28.6	24	5.9
9. Paper products	40	308	26.2	8.49	21.8	71	10.3
10. Printing	9	83	21.3	7.66	8.4	40	9.3
11. Primary metals	34	246	28.8	7.92	24.4	41	6.7
12. Metal fabrication	8	55	29.2	7.65	10.9	14	5.1
13. Machinery	13	92	25.7	7.88	29.3	33	4.5
14. Transportation equip.	35	260	29.3	8.05	35.8	30	6.1
15. Electrical equip.	34	240	26.0	6.68	23.7	30	4.9
16. Nonmetallic mineral products	14	100	25.8	7.66	21.0	65	7.7
17. Petroleum products	1	6	19.7	8.30	16.7	147	23.7
18. Chemicals	10	82	22.2	7.21	23.1	38	6.9
19. Miscellaneous mfg.	4	33	24.5	6.21	18.2	42	6.3
20. All industries	299	2,258	26.3	7.50	22.1	38	6.7

Note. Sample is described in text and Appendix 1. Average real wage rate (in 1981 dollars) is regression-adjusted for the year in which the contract is effective.

or as many as thirteen. Furthermore, the contracts in a given industry are often irregularly distributed over the sample period. For this reason, the average real wage measures in the fourth column of the table are regression-adjusted for the effective dates of the contracts in each industry.[16]

The average strike probability in the sample is 22 percent but varies widely by industry, ranging from a low of 5 percent in the tobacco products industry to a high of 36 percent in the transportation equipment industry. The median duration of work stoppages is 38 days, and the mean duration, as a fraction of the mean duration of contracts, is 6.7 percent. The two measures of strike duration are highly correlated across industries, and both are positively correlated with the probability of disputes.[17] There is a positive correlation between real wage levels and the probability of strikes across industries, and there is also a positive correlation between wage levels and the median duration of strikes.

Table II summarizes the time-series variation in contract characteristics and strike activity during the sample period. The average real wage rate of contracts effective in each year shows an increasing trend until 1977, and relatively little growth threafter.[18] By comparison, the probability and median duration of strikes show remarkable year-to-year variation. Strike probabilities exceeded 30 percent in 1974 and 1975, fell off substantially during the 1976–1978 period of wage and price controls, and then returned to pre-controls levels in 1979.[19] The last few years of the sample suggest a more recent decline in the probability and duration of disputes. The correlation over time between real wage rates (adjusted for industry composition) and strike incidence is small and negative. The correlation between the probability and median duration of disputes is also close to zero.

16. The industry average wages represent estimated industry coefficients from a linear regression of the average real wage during each contract on industry and year effects, with the year effects normalized to sum to zero.

17. Among 18 industries (excluding petroleum, which has only one strike) the correlation between the probability and the median duration of strikes is 0.15, while the correlation between the probability of disputes and the normalized measure of duration is 0.26.

18. To account for year-to-year variation in the industry composition of contract negotiations, the average wage rates in Table II are actually estimated year effects from a regression of average real wage rates on year and two-digit industry effects, with the industry effects normalized to sum to zero.

19. See Riddell [1986] for a description of wage and price controls during this period and a survey of evidence on their effects.

TABLE II
[illegible]TERISTICS OF NEGOTIATED WAGE RATES AND MEASURES OF STRIKE ACTIVITY BY YEAR

Year	Number of contracts	Average contract duration (months)	Average real wage during contract	Strike probability (percent)	Median strike duration
1964	34	35.0	5.53	11.8	17
1965	84	31.9	5.75	22.6	20
1966	72	27.3	5.68	16.7	35
1967	72	28.3	5.90	33.3	41
1968	115	27.4	6.26	19.1	32
1969	78	26.6	6.18	23.1	60
1970	118	28.7	6.72	19.6	37
1971	98	29.1	7.20	26.5	32
1972	101	26.2	6.71	15.8	48
1973	127	27.6	7.26	29.9	28
1974	112	26.3	7.68	36.6	38
1975	128	24.7	7.84	35.9	92
1976	129	23.3	7.88	19.4	45
1977	133	20.5	8.14	14.3	12
1978	164	22.4	7.85	12.8	41
1979	102	25.8	7.85	33.3	45
1980	136	27.0	8.15	17.6	35
1981	77	26.5	8.11	27.3	51
1982	109	24.6	8.41	13.8	45
1983	94	24.8	8.07	24.5	8
1984	109	29.1	8.67	13.8	31
1985	66	28.9	8.35	19.7	20

Note. See note to Table I. Average real wage (in 1981 dollars) is regression-adjusted for two-digit industry composition of contract negotiations in each year.

B. Measurement of Negotiated Wage Outcomes

An analysis of wage outcomes for collective bargaining agreements requires some measure of the wage associated with each contract. Since multiperiod labor contracts typically specify a schedule of wage rates over the life of the contract, there is a variety of alternative choices. One natural measure is the average real wage rate during the term of the agreement. The wage measures in Tables I and II are of this form. Except in very rare instances, however, the real wage rates during the term of a contract are not set directly by the bargaining parties. Instead, the parties specify a series of nominal wage rates, often in conjunction with an indexation formula. This suggests that a more appropriate ex ante measure of the wage rate is the *expected* average real wage rate, conditional on information available at the settlement date of the contract.[20]

Expected real wage rates are not directly observable. Nevertheless, given information on the indexation provisions of a contract and estimates of the expected future price level, it is straightforward to construct estimates of the expected real wage rate. In a nonindexed contract, for example, the logarithm of the real wage rate in month m of the contract, $w(m)$, differs from its expectation as of the negotiation date, $w^*(m)$, by the forecast error in the logarithm of the price level in month m:

$$w(m) = w^*(m) - (p(m) - p^*(m)).$$

Similarly, in an escalated contract with an indexation formula that increases nominal wages by λ percent for each 1 percent increase in the CPI, actual and expected real wages are related by

$$w(m) = w^*(m) - (1 - \lambda)\,(p(m) - p^*(m)).$$

Although most escalated labor contracts do not specify a fixed elasticity of indexation, this equation is approximately correct when λ is replaced by $\lambda(m)$, the marginal elasticity of contractual wages in month m of the contract evaluated at $p = p^*(m)$.[21]

Given an estimate of the parties' expected price level in month m, $\hat{p}(m)$, and an estimate of the marginal elasticity, say $\hat{\lambda}$, an estimate of the expected real wage in month m is

$$\begin{aligned} \hat{w}(m) &= w(m) + (1 - \hat{\lambda})(p(m) - \hat{p}(m)) \\ &= w^*(m) + (1 - \hat{\lambda}) \cdot (\hat{p}(m) - p^*(m)) \\ &\quad - (\hat{\lambda} - \lambda(m)) \cdot (p(m) - p^*(m)). \end{aligned} \tag{8}$$

This estimate differs from the parties' true expectation by two terms: one that depends on the difference between $\hat{p}(m)$ and the parties' expected price level, and another that depends on any errors in approximating the indexation formula in the contract. Provided that these errors are uncorrelated with other observable variables, they can be treated as classical measurement errors.

In the absence of detailed information on the indexation clauses in the sample, I estimate the elasticity of indexation by the ratio of the total escalated wage increases over the life of the contract (in percentage terms) to the total increase in prices over

20. This measure was first used by McConnell [1989].

21. See Card [1983] for a descriptive analysis of indexation provisions in Canadian labor contracts.

the life of the contract. This estimator introduces some noise into the calculation of expected real wage rates, particularly for contracts with restricted escalation clauses. I also use a very simple price-forecasting equation based on the rate of change of consumer prices in the 12 months prior to the effective date of the contract.[22] I have experimented with other price-forecasting equations and found few differences in the resulting estimates.

C. Models of Contractual Wages

The first stage of the empirical analysis is an investigation of the determinants of contractual wage rates. Let w^*_{ij} represent the averge expected value of the logarithm of the real wage rate in the jth contract written by the ith bargaining pair. The model of Section II suggests that w^*_{ij} is a function of the expected profitability of the firm, the labor market opportunities of striking workers, and the dispersion in the unobserved component of profitability. In addition, there is a variety of unobservable determinants of wages, including the skill characteristics of employees who earn the base wage rate.[23] On the assumption that these characteristics are stable over time, a reasonable model for w^*_{ij} is a "fixed effects" model of the form,

$$w^*_{ij} = a_i + X_{ij}b + u_{ij},$$

where a_i is a parameter representing the time-invariant component of wages in the ith bargaining unit, X_{ij} is a vector of measured covariates, b is a vector of coefficients, and u_{ij} is a contract-specific disturbance.

The measured expected average real wage differs from w^*_{ij} by a stochastic component reflecting errors in measuring expectations and errors in measuring the elasticity of indexation. Thus, measured expected average wages follow an equation of the form,

$$\hat{w}_{ij} = a_i + X_{ij}b + \nu_{ij},$$

where ν_{ij} includes both measurement errors and the contract-specific disturbance u_{ij}.

22. The precise forecasting equation for the one-year ahead price level (in logarithms) is $p(t + 12) = p(t) + 0.026 + 0.613^* (p(t) - p(t - 12))$. This equation was selected by fitting various models to the deferred wage increases in two-year nonindexed contracts in the sample, and assuming that the parties designed these increases to maintain the real wage rates as of the start of the contract.

23. In most contracts the base wage rate is earned by a relatively low-skilled group: often janitors or cleaners.

A convenient technique for eliminating the fixed effect a_i is to first-difference over consecutive contracts, yielding

$$(9) \qquad \Delta\hat{w}_{ij} \equiv \hat{w}_{ij} - \hat{w}_{ij-1} = \Delta X_{ij}b + \Delta\nu_{ij}.$$

This equation can be estimated by conventional least-squares methods. However, differencing introduces a first-order moving average component in the residuals of the differenced equation that necessitates a two-step procedure for estimation of consistent standard errors.[24]

It remains to describe the measured determinants of wages. I have assembled data on two proxies for the alternative values of workers' time: a province-specific unemployment rate (measured in the effective month of the contract); and a province-specific wage index for nonproduction laborers (measured in the effective year of the contract). On the firm side I have constructed a series of real value-added price indexes at the three-digit industry level to proxy expected profitability. These indexes are constructed from published annual series of output selling prices and intermediate input price indexes using the approximation,

$$\Delta\log v(t) = (\Delta\log q(t) - h\,\Delta\log r(t))/(1 - h),$$

where $v(t)$ is the value-added price index, $q(t)$ is the price index for gross output, $r(t)$ is the intermediate inputs price index, and h is the average share of intermediate inputs in the gross value of output.

While the value-added price index provides a potentially useful measure of expected profitability, there is no readily available proxy for the uncertainty associated with the unobserved component of profitability.[25] It is interesting to note that the variance of the detrended rate of growth of the value-added price index (estimated over the 1961–1983 period) is significantly correlated with the probability of strikes across industries.[26] This measure of uncertainty is time-invariant, however, and is absorbed by the fixed effects in the statistical model. In the empirical analysis it is therefore necessary to assume that firm-specific variation over time in the dispersion of profitability is negligible.

24. See Holtz-Eakin, Newey, and Rosen [1988] for a complete discussion.
25. Tracy [1986, 1987] uses the variability of firm-specific stock returns in the period prior to contract negotiations as a proxy for the variability of profits. Similar data are available only for a subset of the firms in the contract sample in this paper.
26. Across two-digit industries the correlation of this measure of uncertainty with the probability of strikes is 0.24.

Three other sets of variables are included in the vector of determinants of wages. The first is a series of unrestricted year effects for the effective date of each contract. These year effects control for any aggregate variation in wage growth: due to changing trends in productivity growth, for example, or the imposition of wage-and-price controls during 1976–1978. They also control for any economy-wide changes in the level of uncertainty associated with contract negotiations. A second covariate is the level of real wages at the expiration of the previous contract. The presence of this variable reflects the possibility of "spillovers" between consecutive contracts, resulting from either true state dependence or the presence of slowly changing unobservable determinants of wages.[27] Finally, the third set of explanatory variables consists of measures of the incidence and duration of strikes. According to the one-sided asymmetric information model, wage settlements should be systematically related to the duration of any stoppage.

Estimation results for the first-differenced wage equation (9) are presented in Table III. The first two columns of the table present wage equations that exclude any measures of strike activity, while columns 3–7 add various measures of the incidence and duration of strikes. The estimates in column 1 suggest that wages are negatively related to unemployment rates, positively related to the industry value-added price index, and positively related to the wage at the end of the previous contract. The measure of regional wages, by comparison, has only a negligible effect on negotiated wages.[28] The estimated coefficient of the industry price index is small but very precisely determined. In an effort to verify the robustness of the estimate, I decompose the change in the value-added price index into its two constituent parts: the change in the output selling price (divided by one minus the average share of intermediate inputs h), and the change in the intermediate input price index (multiplied by $h/(1 - h)$). The results yield surprisingly strong support for the value-added specification: the estimated coefficient of the selling price index is 0.042 (with a standard error of

27. Since the level of real wages at the end of the previous contract is essentially a lagged dependent variable, OLS estimates of its coefficient are negatively biased in the first-differenced specification (9). (See Holtz-Eakin, Newey, and Rosen [1988].) As instruments I use year effects for the effective date of the previous contract and the real wage rate in manufacturing industries at the effective date of the previous contract.

28. The estimated coefficient of this variable is larger and statistically significant when the year effects are excluded from the regression, but the overall fit of the equation is significantly worse.

TABLE III

ESTIMATED WAGE DETERMINATION EQUATIONS

(STANDARD ERRORS IN PARENTHESES)

	Dependent variable: first difference of expected average real wage rate						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1. Regional unemployment rate[a]	−0.344	−0.365	−0.328	−0.296	−0.308	−0.280	−0.305
	(0.113)	(0.113)	(0.113)	(0.112)	(0.113)	(0.111)	(0.113)
2. Regional wage of nonproduction laborers[b]	0.013	0.014	0.012	−0.008	−0.010	−0.005	0.010
	(0.044)	(0.043)	(0.044)	(0.043)	(0.043)	(0.044)	(0.043)
3. Industry value-added price index[c]	0.045	0.045	0.043	0.044	0.045	0.046	0.045
	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)	(0.008)
4. Real wage at end of previous contract[d]	0.349	0.309	0.345	0.337	0.341	0.349	0.340
	(0.066)	(0.083)	(0.066)	(0.065)	(0.066)	(0.067)	(0.066)
5. Unexpected component: real wage at end of previous contract[e]	—	0.057					
		(0.059)					
6. Strike indicator	—	—	0.004	0.004	—	—	—
			(0.002)	(0.002)			

	Dependent variable: first difference of expected average real wage rate						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
7. Strike duration (yrs.)	—	—	—	—	0.003 (0.011)	—	—
8. Indicators for strike duration classes:							
(a) 1–7 days (18%)	—	—	—	—	—	−0.001 (0.004)	—
(b) 8–28 days (22%)	—	—	—	—	—	0.002 (0.004)	—
(c) 29–49 days (21%)	—	—	—	—	—	0.004 (0.003)	—
(d) 50–98 days (21%)	—	—	—	—	—	0.012 (0.005)	—
(e) 99–147 days (10%)	—	—	—	—	—	0.007 (0.005)	—
(f) 148+ days (8%)	—	—	—	—	—	−0.011 (0.006)	—
9. Strike duration/mean duration in industry	—	—	—	—	—	—	0.001 (0.002)
10. Aggregate strike prob. in previous 3 months	—	—	—	0.017 (0.009)	0.017 (0.009)	0.018 (0.009)	0.017 (0.009)
11. Standard error	0.039	0.039	0.039	0.039	0.039	0.039	0.039

Notes. Sample consists of 1,467 observations on contracts negotiated between 1966 and 1983. Equations are estimated in first-difference form using changes between consecutive contracts. Standard errors are corrected for heteroskedasticity and moving-average error induced by differencing. Mean and standard deviation of the dependent variable are 0.0438 and 0.0619, respectively. All equations include 17 unrestricted year effects.

a. Seasonally adjusted provincial unemployment rate in effective month of contract.
b. Real wage of nonproduction laborers in province during effective year of contract.
c. Real price index for three-digit level industry value-added.
d. Instrumented with year effects for effective year of previous contract and average wage in manufacturing.
e. Instrumented with same instruments as above.

0.008), while the estimated coefficient of the input price index is −0.038 (with a standard error of 0.011).

The estimated equation in column 2 of Table III presents a second check of the specification in column 1. This regression introduces the unexpected component of the real wage at the end of the previous contract as an additional explanatory variable for negotiated wages.[29] If the bargaining parties set wages to fully "catch up" for any unexpected wage changes during the previous contract, then the coefficient on the unexpected component of real wages should be equal and opposite to the coefficient on the level of real wages, implying no long-run effect of unexpected real wage changes. In fact, the coefficient on the unexpected component of real wages is insignificantly different from zero, suggesting that the degree of persistence of unexpected wage changes is about equal to the degree of persistence of fully anticipated changes.[30]

The third column of Table III introduces a simple indicator variable for the ocurrence of a strike. Contrary to the implications of the model, the estimated coefficient is positive, although only marginally significant. This finding must be interpreted carefully, since the theoretical prediction is conditional on a complete specification of the observable determinants of wages. Failure to adequately measure the alternative wage opportunities of striking workers, for example, could be expected to lead to a positively biased partial correlation between wages and the occurrence of strikes.

The specification in column 4 introduces an additional control variable that may in principle help to control for aggregate-level variation in alternative wage opportunities (or other determinants of wages known to both bargaining parties but unknown to an outside data analyst). This variable is the aggregate probability of strikes among contracts effective in the three months prior to the effective date of the current contract. If the estimated correlation of wages and strikes is biased by the omission of some aggregate-level variable, then conditioning on the overall probability of strikes in the recent past should reduce or eliminate this bias. Despite the fact that wages are positively correlated with the probability of strikes

29. To account for possible measurement error in the calculation of the unexpected component of real wages, this variable is instrumented with the same instruments used for the level of real wages.

30. The implied "catch-up" coefficient is about 0.65—very similar to estimates obtained by Riddell [1979] and Christofides, Swidinsky, and Wilton [1980a, 1980b] in Phillips-style equations for nominal wage changes.

in recent contract negotiations, the addition of this variable has no significant effect on the estimated coefficient of strike incidence.

The fifth column of Table III substitutes a measure of the unconditional duration of work stoppages for the strike indicator variable in columns 3 and 4. Again there is some evidence of a positive effect of strikes on wages, although the duration measure gives a poorer fit than the simple indicator variable. To further explore the relation between wages and strike durations, the regression in column 6 introduces a series of dummy variables for strikes of different lengths. This model includes indicator variables for strikes in each of the first four quintiles of the distribution of strike durations, and separate dummies for strikes in the highest 10 percent and the next highest 10 percent of duration. The estimated coefficients suggest a nonlinear effect of strike duration on wages: the effect is positive and increasing until the fourth quintile (i.e., strikes of 50–98 days duration) and then becomes negative for very long strikes. The estimated wage effect associated with strikes in the fourth quintile is 1.2 percent, and is significantly different from zero at conventional levels. The estimated effect of strikes longer than 148 days is about -1 percent, and is marginally significant.

This specification and the one in column 5 restrict the effects of strikes of equal length to be the same across different industries. There is little reason to expect that the economic impact of strikes is the same in different industries, however.[31] Depending on the ability to stockpile output, for example, a 90-day strike may have a large or small effect on the ability of a firm to fulfill its supply obligations. One simple specification that controls for heterogeneity in the effect of strikes is presented in column 7. Here, I have normalized the duration of a strike by the mean duration of all strikes in the same industry.[32] This specification yields a marginal improvement in fit relative to the simple linear specification in column 5, but is still less successful than the specification in column 6. A similar finding emerged when I normalized strike durations by the mean duration of contracts in each two-digit industry.

I have also fit the specification in column 6 to subsamples of contracts from five of the larger two-digit industries. The results are

31. McConnell [1989] finds that strikes in nonmanufacturing industries have a larger effect on wage settlements than those in manufacturing.

32. Since some industries have very few strikes, I actually use a weighted average of the overall mean duration of strikes (with weight 2) and the mean duration of strikes in the industry (with weight equal to the number of strikes in the industry).

presented in Appendix 2, and indicate the same pattern of coefficients as the pooled data in Table III. In four of the five industries the effect of shorter strikes is positive, while the effect of longer strikes is negative. The fifth industry had no strikes in the longest duration category, but shows a generally positive effect for shorter strikes.

Taken at face value, these findings are inconsistent with a simple one-sided asymmetric information model of strike activity. It is possible, however, that the results simply reflect the omission or mismeasurement of a key determinant of strike activity and wages. Potential candidates include the alternative wage of strikers, the level of uncertainty regarding potential profitability, or other time-varying contract-specific factors that are positively correlated with wages and strike intensity. It is also possible that the findings reflect a more complicated model of strikes. For example, strikes may serve in some instances to reveal the profitability of the firm, and in other instances to reveal private information held by the union. In this case the predicted effect of a strike varies with the reason for the stoppage. As a crude check on this possibility, I augmented the specifications in columns 4 and 5 of Table III with interactions between the strike indicator and two variables: the regional unemployment rate, and the deviation of the value-added price index from its long-run trend. The resulting interaction terms were poorly determined and did not affect the results in Table III. A complete explanation for the pattern of strike effects in these data awaits further investigation.

D. Models of Strike Incidence and Duration

Although the effects of strike incidence and duration on wages are inconsistent with the simple model of strikes outlined in Section II, the pattern of the effects of unemployment rates and industry prices is more favorable. This section presents statistical models that explore the effects of these same variables on the likelihood and duration of strikes. A comparison of the effects of these variables on wages, strike probabilities, and strike durations allows a much stronger test of the asymmetric information model.

Table IV presents estimates of two alternative models of strike incidence. Both models abstract from any pair-specific heterogeneity that might bias the cross-sectional correlation between strike probabilities and the observable explanatory variables. The first is a simple linear probability model. According to this model the

TABLE IV
ESTIMATED STRIKE PROBABILITY EQUATIONS
(STANDARD ERRORS IN PARENTHESES)

	First-differenced linear probability models[a]			Conditional logit models[b]		
	(1)	(2)	(3)	(4)	(5)	(6)
1. Regional unemployment rate	−3.53	−3.57	−3.36	−4.12	−4.43	−4.27
	(1.81)	(1.82)	(1.83)	(2.55)	(2.52)	(2.55)
2. Regional wage of nonproduction laborers	0.59	0.43	0.38	0.84	0.28	0.24
	(0.54)	(0.27)	(0.28)	(0.65)	(0.26)	(0.27)
3. Industry value-added price index	0.49	0.50	0.49	0.25	0.27	0.27
	(0.15)	(0.15)	(0.15)	(0.13)	(0.13)	(0.13)
4. Real wage at end of previous contract	−0.39	−0.43	−0.38	−0.19	−0.28	−0.24
	(0.28)	(—)	(—)	(0.28)	(—)	(—)
5. Unexpected component: real wage at end of previous contract	—	—	−0.41	—	—	−0.42
			(0.47)			(0.68)
6. Standard error	0.56	0.56	0.56	—	—	—
7. −2 Log likelihood	—	—	—	529.48	530.43	530.06

Notes. See notes to Table III for description of explanatory variables. All equations include unrestricted year effects.

a. Estimated on sample of 1,467 contracts negotiated between 1966 and 1983. Equations are estimated in first-difference form using changes between consecutive contracts. Standard errors are corrected for heteroskedasticity and moving-average error component induced by differencing. The mean and standard deviation of the dependent variable are −0.0095 and 0.5733, respectively.

b. Estimated on subsample of 1,110 contracts for 222 bargaining pairs with at least five contracts in the sample. For comparability with the linear probability coefficients, the estimated coefficients and their standard errors are multiplied by a factor of $p\ (1 - p)$, where $p = 0.2613$ is the average strike probability in the subsample.

probability of a work stoppage in the jth contract negotiation of the ith bargaining pair (p_{ij}) is given by an equation of the form,

$$p_{ij} = \alpha_i + X_{ij}\beta,$$

where α_i represents a permanent pair-specific effect and X_{ij} represents a vector of covariates. While the functional form of the linear probability model is open to criticism, it is a convenient one for panel data because it leads to a linear regression for the first-difference of measured strike incidence that is free of pair-specific heterogeneity. Specifically, the linear probability model implies that

$$\Delta y_{ij} = \Delta X_{ij}\beta + \Delta\phi_{ij},$$

where y_{ij} is an indicator for a strike in the jth contract negotiation of pair i, and ϕ_{ij} has the interpretation of a conditionally heteroskedastic zero-mean residual.[33] Estimates of this equation are presented in the first three columns of Table IV, using the same sample of contract negotiations as in Table III. The vector X consists of the regional unemployment rate, the regional wage of nonproduction laborers, the industry-specific value-added price index, and the level of real wages at the end of the preceding contract. As in Table III the estimated models include a set of unrestricted year effects for the effective date of each contract.

The estimates in column 1 suggest that strike probabilities are significantly negatively related to regional unemployment rates and significantly positively related to the industry-specific value-added price index. While the former effect is consistent with the theoretical model, the finding that strike probabilities increase with the real value of industry output stands in sharp contrast to the prediction of the model. As a check on the coefficient I entered the two constituent parts of the change in value added (the change in the real price of gross output and the share-weighted change in the real price of intermediate inputs) separately in the linear probability model. Again the results lend strong support to the restricted specification: the estimated coefficient of the output price index is 0.51 (with standard error of 0.15), while the estimated coefficient of the intermediate inputs price index is -0.53 (with a standard error of 0.20).

Although the estimated coefficients of the wage variables in column 1 are poorly determined, their signs and magnitudes suggest that it is the change in *relative* contractual wages that affects the probability of disputes. This hypothesis is embodied in the specification in column 2, which restricts the effects of outside wages and wages at the end of the previous contract to be equal and opposite. The restriction easily passes a conventional test, but its imposition yields very little improvement in the estimated standard errors associated with the wage variables.

Finally, the model in column 3 introduces the unexpected part of the real wage at the end of the previous contract as a separate explanatory variable.[34] This specification allows a test of the

33. By construction the residual term $\Delta\phi_{ij}$ exhibits first-order moving average serial correlation. The estimated standard errors in columns 1–3 of Table IV account for both heteroskedasticity and first-order correlation.

34. For reasons described above, the unexpected component of real wages may be measured with error. In recognition of this fact I reestimated the model in column 3 by two-stage least squares using the same instruments as in Table III. The resulting coefficient estimate was very similar to the one reported in the table.

hypothesis that unexpected changes in real wages (arising from unexpected inflation during the previous contract) affect the likelihood of strikes more or less than predictable movements in real wages. The estimates are relatively imprecise; nevertheless, there is little indication of a differential effect between expected and unexpected wage changes.

An alternative to the linear probability model that also permits an unrestricted specification of heterogeneity is a logistic probability model with individual effects:

$$\log (p_{ij}/(1 - p_{ij})) = \alpha_i + X_{ij}\beta.$$

The individual effects in this model can be conveniently eliminated by maximizing a conditional likelihood function, as described by Chamberlain [1980]. The conditional likelihood is rather unwieldy for an unbalanced panel, however. For convenience, I therefore selected a balanced subsample of contracts drawn from the set of bargaining pairs with complete data on five or more contract negotiations. The resulting panel contains 222 pairs and a total of 1,110 contracts. The characteristics of the subsample are generally similar to those of the overall sample. The average strike probability in the subsample is 26.1 percent.

Estimates from the logistic probability model using the same configuration of explanatory variables as in columns 1–3 are presented in columns 4–6 of Table IV. For ease of comparison between the linear probability and logistic specifications, I have renormalized the logistic coefficients to represent the estimated effects of the explanatory variables on the average probability of a strike.[35] Overall, the estimates and inferences from the two specifications are very similar. Both models confirm that strike probabilities are reduced by higher unemployment and increased by higher prices for net industry output.[36] The latter finding must be viewed as evidence against the simple model presented in Section II, and similarly against any model that embodies the "joint cost" hypothesis of strike incidence.

Table V presents a series of estimated models for the completed duration of work stoppages. The sample of strikes is drawn from the set of contract negotiations underlying the wage equations

35. This is accomplished by multiplying the estimated logistic coefficients and their standard errors by $p(1 - p)$, where p is the sample average strike probability.

36. As in the linear probability specification, the restrictions implied by the use of the value-added price index are easily passed in the conditional logit specification.

TABLE V
ESTIMATED STRIKE DURATION EQUATIONS
(STANDARD ERRORS IN PARENTHESES)

	Dependent variable: log of completed strike duration					
	18 two-digit industry effects			203 bargaining pair effects		
	(1)	(2)	(3)	(4)	(5)	(6)
1. Unrestricted year effects	No	Yes	Yes	No	Yes	Yes
2. Regional unemployment rate	−1.30	7.36	8.38	−7.17	−4.49	−0.37
	(3.22)	(5.17)	(5.18)	(5.91)	(13.79)	(14.06)
3. Regional wage of nonproduction laborers	0.77	−0.28	−0.28	3.08	0.08	0.34
	(0.73)	(1.04)	(1.04)	(1.64)	(3.83)	(3.82)
4. Industry value-added price index	1.06	0.98	1.03	1.33	1.20	1.23
	(0.26)	(0.27)	(0.27)	(0.57)	(0.61)	(0.61)
5. Real wage at end of previous contract	−1.09	−0.35	−0.19	−1.64	−1.03	−0.49
	(0.44)	(0.46)	(0.46)	(1.38)	(1.51)	(1.55)
6. Unexpected component: real wage at end of previous contract	—	—	−5.18	—	—	−5.36
			(2.65)			(3.81)
7. Standard error	1.25	1.20	1.20	1.30	1.24	1.24

Notes. See notes to Table III for description of explanatory variables. Sample consists of 403 strikes occurring in negotiations of 203 bargaining pairs between 1966 and 1983. The mean and standard deviation of the dependent variable are 3.383 and 1.294, respectively.

in Table III and the linear probability models in Table IV. The sample contains a total of 403 strikes from 203 distinct bargaining pairs.

The regression functions in Table V can be interpreted as estimates of the expected log strike duration function. Alternatively, on the assumption that strike durations are exponentially distributed, these estimates can be interpreted as estimates of the logarithm of the inverse hazard function for strike settlements.[37]

In view of the relatively small sample of strikes, I have fit models with both two-digit industry effects (columns 1–3) and a complete set of bargaining-pair effects (columns 4–6). Columns 1 and 4 report specifications that exclude year effects, while the remaining models include an unrestricted set of year effects for the

37. If the completed duration of strikes (S) is exponentially distributed with hazard μ, then $E(\log(S)|\mu) = \text{constant} - \log(\mu)$. See Jones [1988, p. 747].

effective year of the contract whose negotiation resulted in the strike. The other explanatory variables are the same as in Table IV.

A comparison of models with and without year effects suggest that they provide a significant improvement in fit.[38] In contrast, there is no indication that the addition of bargaining-pair effects provides a significant improvement over and above the industry-level controls. This may well reflect the small number of strikes relative to the number of bargaining pairs, however.

With the exception of the coefficients of the industry price variable, none of the other estimated coefficients is consistently different from zero, and many of the coefficients vary substantially across specifications. The positive effect of the industry selling price variable is consistent with its positive effect on strike incidence, but inconsistent with the implications of the theoretical model. There is some indication that strike durations fall with higher unemployment rates, although this inference depends on the treatment of year and bargaining-pair effects. Unlike the models for wages and strike probabilities, the model for strike durations suggests that the conditional duration of work stoppage is affected differentially by the expected and unexpected components of the real wage at the end of the previous contract. Again, precise inferences are made difficult by the small sample size.[39]

IV. Summary and Conclusions

This paper has presented and tested a simple model of strikes based on the hypothesis that costly disputes arise from one-sided asymmetric information over the profitability of the firm. The model predicts the existence of a negatively sloped resistance curve relating lower wage settlements to longer strikes. The model also specifies the joint effects of changes in the expected profitability of the firm and changes in the alternative opportunities of striking workers on wages, strike probabilities, and the conditional duration of disputes.

In common with many other theories of strikes, the model predicts that strike incidence and duration will decrease when the expected surplus from a bargain is raised. Thus, increases in the

38. The F-tests are significant at levels below 0.1 percent.

39. Gunderson and Melino [1988] and Harrison and Stewart [1989] estimate strike duration models using a much larger file of Canadian strikes. Both sets of authors find that strike durations are countercyclical.

expected profitability of the firm are predicted to reduce the probability and duration of strikes, while increases in the alternative opportunities of striking workers are predicted to increase the probability and duration of disputes. The model also predicts that increases in expected profitability and increases in workers' outside opportunities will increase negotiated wages.

These implications are tested using a large sample of collective bargaining agreements from the Canadian manufacturing sector. A simple model of contractual wage outcomes is developed using the expected average real wage rate during the term of the agreement. Negotiated wages are found to depend negatively on unemployment rates and positively on the real price of industry output. Contrary to the premise of the model, however, and contrary to the findings of McConnell [1989] with U. S. data, wage outcomes are not related to strike durations in a simple monotonic fashion. Rather, the data suggest a nonlinear relation, with small positive effects (1 to 1.5 percent) of moderately long strikes and small negative effects (−1 percent) of very long disputes.

Simple models are also estimated for the probability and duration of disputes. As predicted by the theory, increases in unemployment, which are interpreted as reductions in the alternative opportunities of striking workers, reduce the probability of strikes. The effect on strike duration is imprecisely measured but may also be negative. In contrast to the predictions of the theory, however, strike probabilities and conditional durations are positively related to the real price of industry value-added.

On balance, the evidence in favor of the model is weak. While it is possible that the empirical findings are biased by the omission of unobserved determinants of wages and strike outcomes, it may also be that a richer model of strikes, allowing for asymmetries on both sides of the bargaining table, is needed to fully describe the data.

APPENDIX 1

A. Derivation of the Contract Sample

The contract sample was derived from the wage settlement file made available by Labour Canada in three stages. First, I extracted all 2,868 manufacturing agreements on the December 1985 version of the file. Second, the sample was checked for duplicate contracts between the same firm and union covering employees in different establishments. Two contract chronologies were merged together if they had the same starting and ending dates and identical wage and strike information. There are several instances of multiple-employer bargaining units that are coded separately for some contract negotiations and jointly for other negotiations. In these cases I merged together the related contracts in all years to form a single chronology for the multiple-employer bargaining unit. There are also some cases where the Labour Canada identification number for a given bargaining unit changes between negotiations. In these cases I concatenated the contract chronologies to form a single continuous chronology. The final step of the sample derivation was to extract the subset of contracts from bargaining pairs with at least four consecutive contract negotiations in the file.

In some cases the base wage rate definition changes between consecutive contracts in the Labour Canada file (for example, between "janitors and sweepers" in one contract and "assemblers" in the next). The ending wage rate for each contract was checked against the wage reported in the next contract as the "old rate." In cases where a change of definition occurred, the base wage series were index-linked to form a consistent wage series.

The merging of strike duration information into the contract sample involved two steps. First, firm and union names, locations, and settlement dates were listed for all contracts in the sample that were recorded as settling after a work stoppage. The appropriate issue of *Strikes and Lockouts in Canada* (*SLC*) was then checked for information on the duration of the dispute. Second, in order to identify strikes that were reported in *SLC* but not recorded on the Wage Tape, every strike listing in *SLC* from 1964 to 1984 was checked against the list of firm and union names generated from the contract data set.

The results from the first step of the merging process revealed a coding error in the Labour Canada file for settlements in 1980–1981. In these two years 51 agreements were coded as settling at the stage of "bargaining after a work stoppage." In none of these cases was there a matching strike listing in *SLC*, or a record of the strike in the contract extract published in Labour Canada's *Collective Bargaining Review*. I therefore recoded these agreements as settling without a work stoppage.

The results from the second step of the merging process revealed that in approximately 5 percent of cases where no strike was recorded in the Labour Canada file, a strike actually occurred during the contract negotiations. These strikes were distinguished from intracontract wildcat strikes by their dates and by *SLC* information on the cause of the dispute.

The following table shows the distribution of final strike outcomes by their original recording status in the Labour Canada File for the entire sample of 2,868 manufacturing contracts.

		Original status	
		No strike	Strike
Final status	No strike	2,145	51
	Strike	100	572

There was a total of five strikes for which no strike duration information was found in *SLC*. There are also 28 instances of strikes that occurred in two or more spells. In most of these cases, the initial spell(s) lasted less than one week. For these strikes I recorded the strike duration as the duration of the longest spell.

B. Aggregate Data

The following aggregate data were merged to each contract listing, by the effective date of the contract.

a. Consumer price index, all items, 1961 = 100. January 1961–November 1985: Cansim D484000, from the 1985 University Base Tape. December 1985 to June 1986: Cansim D484000, from the *Bank of Canada Review,* November 1986.
b. Unemployment rates, seasonally adjusted. Rates for January 1966–November 1983 were obtained from the 1983 Cansim University Base Tape. Rates for December 1983–December 1985 were obtained from the *Bank of Canada Review,* November 1986. The following series were used: Quebec-Cansim D768478; Ontario-Cansim D768648; British Columbia-Cansim D769233; all other provinces-Cansim D767611 (national rate).
c. Industry output and input prices and indexes of output. Three-digit industry data for 1961–1971 were taken from Statistics Canada, *Real Domestic Product by Industry 1961–71* (Ottawa: Statistics Canada). These data are classified by industry on the basis of the 1960 standard industrial codes. Data on a 1971 industry code basis for 1971–1983 were taken from the 1978 and 1984 issues of *Gross Domestic Product by Industry* (Ottawa: Statistics Canada). The 1960

and 1971 industry codes were then matched, and the price and output series spliced at 1971. There were 31 (of 65) three-digit industries for which data were not available on a consistent basis. For these industries, appropriate two-digit industry data were used.

d. Average hourly earnings of nonproduction laborers, by province. Annual hourly earnings for selected occupations are available by city. I matched data for the following cities to their respective provinces: Halifax, St. John, Montreal, Toronto, Winnipeg, Regina, Edmonton, Vancouver. The wage rates used are listed as wages for "general male labourers" between 1966 and 1977, for "general labourers in service occupations" between 1978 and 1981, and for "non-production labourers" between 1982 and 1985. Data for 1969–1972 are taken from *Wage Rates, Salaries, and Hours of Labour* (Ottawa: Canada Department of Labour), 1966–1972 editions. Data for 1973–1985 are taken from *Canada Year Book* (Ottawa: Statistics Canada), various editions. For contracts that cover two or more provinces, I used a weighted average of rates for Montreal, Toronto, and Vancouver, with weights 0.35, 0.55, and 0.10, respectively.

APPENDIX 2:
ESTIMATED WAGE DETERMINATION EQUATIONS FOR FIVE MAJOR INDUSTRIES
(STANDARD ERRORS IN PARENTHESES)

	Industry:				
	Food and beverages	Pulp and paper	Primary metals	Transport equipment	Electrical products
	(1)	(2)	(3)	(4)	(5)
A. Sample characteristics					
1. Sample size	223	200	154	164	156
2. Strike probability	14.3	26.0	27.9	34.8	25.6
3. Mean of dependent var.	0.041	0.045	0.054	0.044	0.042
4. Std. deviation of dependent var.	0.062	0.059	0.058	0.071	0.057
B. Estimated coefficients					
1. Regional unemployment rate	−0.095	0.031	−0.654	0.060	−0.755
	(0.295)	(0.209)	(0.417)	(0.490)	(0.357)
2. Regional wage of non-production laborers	0.170	0.040	−0.057	−0.051	0.024
	(0.095)	(0.026)	(0.136)	(0.125)	(0.114)
3. Industry value-added price index	0.044	0.026	0.141	−0.026	0.004
	(0.041)	(0.039)	(0.049)	(0.030)	(0.047)

APPENDIX 2:
(CONTINUED)

	Industry: Food and beverages	Pulp and paper	Primary metals	Transport equipment	Electrical products
	(1)	(2)	(3)	(4)	(5)
4. Real wage at end of previous contract	0.237	0.568	0.231	−0.055	0.368
	(0.119)	(0.110)	(0.133)	(0.125)	(0.119)
5. Indicators for strike duration classes:					
(a) 1–7 days	−0.009	0.013	0.010	−0.016	0.003
	(0.015)	(0.014)	(0.013)	(0.011)	(0.010)
(b) 8–28 days	0.027	0.003	0.007	0.010	0.013
	(0.014)	(0.005)	(0.010)	(0.011)	(0.009)
(c) 29–49 days	0.013	0.002	0.023	−0.016	−0.006
	(0.008)	(0.008)	(0.015)	(0.012)	(0.008)
(d) 50–98 days	0.007	0.002	−0.012	0.014	0.006
	(0.012)	(0.006)	(0.011)	(0.011)	(0.011)
(e) 99–147 days	0.035	0.010	0.001	0.007	0.036
	(0.028)	(0.007)	(0.012)	(0.016)	(0.013)
(f) 148+ days	−0.001	−0.011	−0.006	−0.024	—
	(0.023)	(0.007)	(0.019)	(0.017)	
6. Aggregate strike prob. in previous 3 months	0.030	0.015	0.007	−0.036	0.022
	(0.019)	(0.026)	(0.023)	(0.033)	(0.026)
7. Standard error	0.037	0.023	0.039	0.046	0.034

Notes. See notes to Table III for description of explanatory variables. All equations are estimated in first difference form using changes between consecutive contracts, and contain unrestricted year effects. Dependent variable in each case is the first difference of the expected average real wage rate. Estimated standard errors are not corrected for moving average error component induced by differencing.

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