CYCLICAL VARIATION IN INDIVIDUAL CONDITIONAL STRIKE-SETTLEMENT PROBABILITIES*

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ABSTRACT

Empirical work on strikes has tended primarily to concern itself with strike incidence, with the recurring finding that strikes occur procyclically. Less attention has been paid to strike duration and the cyclical fluctuations it might display, although a variety of theoretical models have been proposed suggesting that strike duration will vary with the business cycle. This paper presents an empirical investigation of this issue, using Canadian microdata to estimate flexible models of conditional strike-settlement probabilities. Two contrasting hazard models are estimated and compared. Both utilise a flexible polynomial representation for the duration dependence and explicitly incorporate unobserved heterogeneity. The first is a beta-logit model recently proposed by Kennan (1984), and the second is a discrete analogue of the familiar proportional-hazards model with a Gamma mixing distribution. Score tests for a number of potential specification errors are also constructed for both models.

The models are estimated on a sample of individual contract strikes in manufacturing over wage issues during the period 1971-1980. The results strongly support the hypothesis that strike duration varies countercyclically, with conditional settlement probabilities at the peak of the cycle being up to twice as high as at the trough. A supplementary finding is that this cyclical sensitivity is even more marked in the paper and printing industries, and that individual conditional settlement probabilities there are much lower, ceteris paribus, than elsewhere in manufacturing. The hypothesis of homogeneity across industries and unions in the remainder of manufacturing is accepted by the data.

I. INTRODUCTION

That strike activity varies positively with the business cycle has become part of economic folklore. A feature of much of the research on this subject, however, is a tendency to concentrate solely on strike incidence without considering strike durations. Additionally, the majority of studies are at a highly aggregated level, being either time-series studies of the macro-economy (or the manufacturing sector) or industry cross-section studies. This aggregation severely limits the inferences that can be drawn and also potentially biases the estimates.

In parallel there has developed a small statistical literature on the specification and estimation of the distribution of individual strike durations.[1] Although the data are sometimes partitioned by broad industry or issue group, the impact of exogenous factors such as variations in economic activity over the business cycle is not typically examined.[2] This is the issue addressed in this paper. More specifically, we present an investigation of models for individual strike settlement probabilities and examine the impact on them of various exogenous factors, including cyclical fluctuations.

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1. See, *inter alia*, Lancaster (1972) and Newby and Winterton (1983).

2. A notable exception is Kennan (1984), who finds that strike duration is counter-cyclical. Kennan's paper is, however, primarily concerned with methodology, and the robustness of the finding is not investigated.

An investigation of this type depends crucially on the specification of the hazard function, representing the conditional settlement probability of the strike. The framework for such models is outlined in section II and two contrasting flexible models described. Because of the importance of using a satisfactory specification of the hazard (including choice of exogenous variables), diagnostics for the estimated models are particularly important. To this end a number of score tests against possible misspecifications are constructed, also in section II.

To undertake our investigation, we make use of a particularly rich set of data on strikes in Canada, covering all stoppages that involved three or more workers and lasted more than half a day. Brief details of the data set are given in section III. As indicated above, one particular focus of the paper is the impact of fluctuations in the business cycle, and in section III we also briefly discuss models that suggest why strike durations may vary cyclically, and which cyclical variables they may be most closely related to. The appropriate measurement of the business cycle is also considered.

Besides cyclical effects, we are interested too in the impact of other factors (associated with the plant, the strike, the union, the industry, the macro-economy, etc.).

The theoretical literature on strikes is not particularly helpful in specifying an operational list of the key factors influencing the durations of individual strikes, but what assistance it does offer is reviewed in section III.

The results for the two contrasting models are presented and discussed in section IV. The proposed score test statistics are examined and some reformulations of the model considered as a result. All the models estimated support the hypothesis that conditional strike-settlement probabilities vary pro-cyclically and hence that strike durations vary counter-cyclically, and this and our other main conclusions are summarised in section V.

II. MODELS FOR STRIKE-SETTLEMENT PROBABILITIES

In this section we discuss the general framework for analysing individual strike-settlement probabilities and present two contrasting models within this framework. A preliminary issue that needs to be addressed concerns the question of whether the model should be cast in discrete or continuous time. Heckman and Singer (1984a) argue that the use of continuous-time models is to be preferred for two reasons. First, because there is no natural time interval within which agents make decisions, it is more appropriate to treat them as if they make them continuously; and secondly, even if there were a natural time interval, this would not in general accord with the interval of measurement of the data.

While these arguments are valid in the context of the type of data typically available on individual labour-market transitions that Heckman and Singer have in mind, it is not obvious that the same considerations apply in the case of an investigation of strike durations. In the case of the potential settlement of a strike, agents do not take decisions continuously, but rather determine precisely the observed variable in our case, namely the length of the strike measured in working days lost; many strikes (and probably most of those occurring at the renegotiation of a contract) are called to start at the beginning of a working day, and when a settlement is reached, the return to work most often likewise takes place at the beginning of a working day. We have therefore opted for a discrete-time framework.

We now turn to the general formulation of discrete-time, strike-duration models with unobserved heterogeneity. Consider a strike with an elapsed duration of t days. Let the conditional settlement probability, or hazard, at t days be $h(t;x,v)$, where x is the vector of observed characteristics (of the plant, the strike, the union, the industry, the macro-economy, etc.) and v is a scalar representing the unobserved heterogeneity. The conditional continuation probability for a strike that has lasted t days is thus

$$p(t;x,v) = 1-h(t;x,v),$$

and the survivor function for an individual strike, the probability that it will last more than d days, is given by

$$S(d;x,v) = \prod_{t=0}^{d} \{1-h(t;x,v)\}.$$

The likelihood of a strike with characteristics x and v having completed duration D days is then

$$L(D;x,v) = S(D-1;x,v)-S(D;x,v).$$

Integrating the unobserved v out of the likelihood function gives

$$L(D;x) = S(D-1;x)-S(D;x),$$

where $S(D;x) = \int S(D;x,v)\cdot f(v)\cdot dv$, $f(v)$ being the density of v (conditional on x). This likelihood can also be written as

$$L(D;x) = h(D;x) \prod_{t=0}^{D-1} \{1-h(t;x)\},$$

where

$$h(t;x) = 1-[S(t;x)/S(t-1;x)]$$

can be regarded as the "aggregate" hazard for strikes that have lasted t days. A wide variety of models then result according to the functional form chosen for the hazard, $h(t;x,v)$, and the distribution assumed for the unobserved heterogeneity component, v.

There is of course an identification problem in attempting to distinguish between omitted heterogeneity and duration dependence. More than one combination of specifications of $h(t;x,v)$ and the distribution of v (i.e. the mixing distribution) can give rise to any given form for the conditional distribution of observed durations. It therefore seems unwise to infer too much from the estimates of a single model without considering how robust they are to the assumptions made. If, despite this problem, very different specifications without too many restrictions on them give rise to similar conclusions with respect to the impact of the exogenous factors and the direction of the duration dependence, one might reasonably be more confident about the robustness of any inferences drawn. It seems unlikely that any robust conclusions will be possible with regard to either the exact shape of the hazard over the elapsed duration or the relative magnitude of the impact of the omitted heterogeneity in the present context.[3]

In the light of these remarks it seems appropriate (a) not to restrict our analysis to a single form of model, and (b) to adopt a very flexible specification for the duration dependence in the hazard function. We now consider two contrasting flexible models within this framework.

A Beta-Logit Model

This model, considered by Kennan (1984), assumes a beta distribution for the unobserved heterogeneity, with a mean dependent on x, takes a logit form for the representative hazard[4] and utilises a flexible polynomial representation for the duration dependence. The heterogeneity is assumed to enter the continuation probability multiplicatively:

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3. If one has a sufficiently strong prior belief about the form of the individual hazard function (presumably derived from an underlying economic model) then it is possible to estimate such models without imposing any (except very weak) distributional assumptions on the form of f(v). Heckman and Singer (1984b) propose a non-parametric, maximum-likelihood estimator for the case of a continuous-time, proportional-hazards model in this context.

4. Use of the term "representative" should not be taken to imply that the associated strike is in any sense "typical". Rather it refers to the (hypothetical) strike whose individual hazard is at the mean of the mixing distribution.

$$1-h(t;x,v) = p(t;x,v)$$
$$= v\cdot p_1(t;x); \quad (0 < v, p_1(t;x) < 1).$$

v is taken to have a beta distribution with mean $\Pi(x)$ and variance $[\alpha/(1+\alpha)]\Pi(x)(1-\Pi(x))$, with α assumed to be a positive constant that does not depend on x. The individual continuation probability for the representative strike is specified as

$$p(t;x,\Pi(x)) = 1/[1+\exp(x\beta+\theta(t))],$$

where $\theta(t)$ is a polynomial in t (without a constant term).[5] The variation around this is then determined by that of the initial continuation probabilities, so that v is identified with p(0;x,v) and hence

$$\Pi(x) = 1/[1+\exp(x\beta)].$$

This yields

$$p_1(t;x) = [1+\exp(x\beta)]/[1+\exp(x\beta+\theta(t))].$$

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5. Nickell (1979) uses this logit form with a quadratic imposed for $\theta(t)$. He specifies a discrete two-point density for the omitted heterogeneity term.

Integration then gives the aggregate hazard as

$$h(t;x) = 1 - \frac{1+\alpha s(1+\exp(x\beta))}{(1+\alpha s)(1+\exp(x\beta+\theta(t)))} .$$

A Gamma-Proportional-Hazards Model

In this alternative model we introduce an unobserved heterogeneity component into the proportional-hazards model in an analogous way to that in which the observed heterogeneity enters. In the continuous-time proportional-hazards model (without unobserved heterogeneity) introduced by Cox (1972), the conditional survivor function can be written

$$S^*(d;x) = \{S_0^*(d)\}^{\exp(x\beta)} ,$$

where $S_0^*(d)$ is the base-line survivor function and asterisks are used to indicate components of the homogeneous model. A discrete analogue of the proportional-hazards model can be obtained by applying this survivor function relationship directly to a discrete model (see Kalbfleisch and Prentice (1973)). If the hazard function corresponding to S_0^* has contribution λ_t at t $(0<\lambda_t<1)$ then

$$S_0^*(d) = \prod_{t=0}^{d} (1-\lambda_t),$$

and hence

$$S^{*}(d;x) = \prod_{t=0}^{d} (1-\lambda_t)^{\exp(x\beta)} .$$

The hazard at t given the vector of characteristics x can then be written

$$h^{*}(t;x) = 1 - (1-\lambda_t)^{\exp(x\beta)} .$$

It can be shown that this discrete model is also the one that arises from grouping the continuous model (see Kalbfleisch and Prentice (1980, p.37)). In some sense it is therefore the natural analogue.[6] We allow unobserved heterogeneity to enter the model in the same way as observed heterogeneity enters, and hence write the hazard with unobserved heterogeneity as

$$h(t;x,u) = 1 - (1-\lambda_t)^{\exp(x\beta+u)} ,$$

where u is a scalar random variable representing the unobserved heterogeneity,[7] and the survivor function as

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6. There are other possibilities; see, inter alia, Cox and Oakes (1984, p.101).

7. For small h, $-\log(1-h) \approx h$; for small λ, $-\log(1-\lambda) \approx \lambda$. Hence $h(t;x,u) \approx \exp(u)\cdot\exp(x\beta)\cdot\lambda_t$, so that there is approximate proportionality in the hazard.

$$S(d;x,u) = \prod_{t=0}^{d} (1-\lambda_t)^{\exp(x\beta+u)} .$$

A convenient distributional assumption here is to take $v=\exp(u)$ to have a Gamma distribution with unit mean and variance σ^2. Then integrating v out gives the aggregate survivor function

$$S(d;x) = \{1+\sigma^2 \cdot \exp(x\beta) \cdot \sum_{t=0}^{d} \exp(\gamma_t)\}^{-1/\sigma^2} ,$$

where we have reparameterised the λ_t in terms of unconstrained parameters by using $\gamma_t=\log(-\log(1-\lambda_t))$.[8]

<u>Score Tests for Specification Errors</u>

We now turn to the construction of score tests for specification errors in the two models outlined above. These diagnostics are particularly useful in the present context since estimation of the models is relatively expensive when the number of parameters is large, and construction of the score test statistics requires only estimation under the null together with a simple calculation by means of a linear regression.

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8. This use of a Gamma mixing distribution is similar to the introduction of unobserved heterogeneity into a continuous-time model by Lancaster (1979) and Lancaster and Nickell (1980).

We consider first the omission of relevant explanatory variables. Suppose that under the alternative hypothesis $x\beta$ should be replaced by $x\beta+z\rho$ in the specification of the hazard, where z is a k-vector of omitted explanatory variables. We construct here a score test for $H_0 : \rho = 0$. Define

$$g_i = \partial \log L_i / \partial \rho,$$

$$c_i = \partial \log L_i / \partial \phi,$$

where L_i is the contribution of the ith observation to the likelihood and ϕ is the vector of parameters (β plus the heterogeneity parameter) that remain in the model under the null. We can then define the efficient score for ρ as

$$g_* = (1/\sqrt{n})\cdot(\partial \log L/\partial \rho) = (1/\sqrt{n})\cdot \sum_{i=1}^{n} g_i .$$

The asymptotic distribution of g_* is then k-variate normal with mean zero and variance Ω (see Rao (1973, pp.415-6)). The information matrix on $\theta=(\phi',\rho')'$ in a single observation can be written, under the usual regularity condition (see Rao (1973)), as

$$J = E\{(\partial \log L_i/\partial \theta)\cdot(\partial \log L_i/\partial \theta')\}.$$

Hence, using the notation above we can write

$$J = E\left\{\begin{pmatrix} c_i \\ g_i \end{pmatrix}\begin{pmatrix} c_i \\ g_i \end{pmatrix}'\right\} = E\begin{bmatrix} c_ic_i' & c_ig_i' \\ g_ic_i' & g_ig_i' \end{bmatrix} \quad ,$$

and Ω^{-1} will be the bottom right hand block of the inverse of this matrix.

If $\hat{\Omega}$ is a consistent estimator of Ω, the score test statistic

$$\xi = \hat{g}_*'\hat{\Omega}^{-1}\hat{g}_* \quad ,$$

where $\hat{g}_*$ is evaluated at the maximum-likelihood estimates under the null, will be asymptotically $\chi^2(k)$ under $H_0 : \rho = 0$. A consistent estimator of J is provided by

$$(1/n)\begin{bmatrix} \hat{C}'\hat{C} & \hat{C}'\hat{G} \\ \hat{G}'\hat{C} & \hat{G}'\hat{G} \end{bmatrix} = (1/n)(\hat{C}:\hat{G})'(\hat{C}:\hat{G})$$

$$= (1/n)(\hat{B}'\hat{B}) \quad ,$$

where $\hat{C}$, $\hat{G}$ are matrices with ith rows $\hat{c}_i'$, $\hat{g}_i'$ respectively, which are in turn c_i', g_i' evaluated at the maximum-likelihood estimates under the null. A consistent estimator of Ω^{-1} is therefore given by the bottom right hand block of $n(\hat{B}'\hat{B})^{-1}$. Thus since $\ell'\hat{C} = 0$ and $\ell'\hat{G} = \sqrt{n}.\hat{g}_*$, where ℓ is the unit n-vector, the score test statistic can be written

$$\xi = \iota'\hat{B}(\hat{B}'\hat{B})^{-1}\hat{B}'\iota,$$

which is the explained sum of squares (or n times the uncentred R^2) from the regression of ι on $\hat{B}$.

We now turn to the elements of $\hat{B}$ in the two models presented in the previous section. In both cases g_i can be written as a multiple of the z-vector:

$$g_i = \Psi_i z_i.$$

Since Ψ_i is the contribution of the ith observation to the score for the "intercept" in β, it is extremely straightforward to output from the maximum-likelihood program.

In the case of the beta-logit model,

$$\Psi_i = \sum_{t=0}^{D-1} \{\lambda_i(t) - \mu_i(t)\} + \lambda_i(D),$$

where, under the null,

$$\lambda_i(t) = 1/\{1+\exp(x_i\beta+\theta(t))\}$$

and

$$\mu_i(t) = (1+\alpha t)/\{1+\alpha t(1+\exp(x_i\beta))\}.$$

$\lambda_i(t)$ is the individual continuation probability for a representative strike and $\lambda_i(t)/\mu_i(t)$ is the aggregate continuation probability.

In the case of the gamma-proportional-hazards model,

$$\psi_i = \frac{1}{\sigma^2}\left[\frac{S(D-1;x)^{1+\sigma^2} - S(D;x)^{1+\sigma^2}}{S(D-1;x) - S(D;x)} - 1\right]$$

for $\sigma^2 \neq 0$.

The elements of c_i corresponding to the β-vector will be ψ_i times the x-vector in both models. The elements of c_i corresponding to the remaining parameters in the model (the coefficients in the polynomial in elapsed duration and the omitted heterogeneity parameter) are more complicated in form and are more conveniently output directly from the maximum-likelihood program as the individual contributions to the score.

Score tests can also be used on the order of the polynomial in elapsed duration. In this case the estimated score contributions corresponding to the omitted higher powers of t are again most conveniently calculated within the maximum-likelihood program. Hence an alternative approach to the calculation of score test statistics is employed. The score test statistic can be constructed by applying the Wald formulation to the estimates from one iteration of Fisher's scoring algorithm from the maximum-likelihood estimates under the null (see Breusch and Pagan (1980) or Engle (1984)). Since the negative Hessian provides a consistent estimator of the information matrix, an iteration of the Newton-Raphson algorithm can be used instead. This is the method employed for the tests of the order of the polynomial in elapsed duration in the empirical work reported here.[9]

Turning to second-order misspecification, we might appropriately consider alternatives in which there is variation in α or σ^2. In the beta-logit model, if we write $\alpha=\alpha_0\exp(za)$ under the alternative, the score test statistic can be calculated as the explained sum of squares from a regression of ℓ on the score contributions under the null together with $(\partial \log L_i/\partial\alpha)z$. A score test for the constancy of σ^2 in the gamma-proportional-hazards model can be constructed in a parallel way.

III. EMPIRICAL IMPLEMENTATION OF THE MODELS

The Micro-Data Set

The micro-data on individual strikes employed in this

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9. The adequacy of the score test in this connection was examined by comparing the result with that obtained from unconstrained estimation of models with polynomials of different order; see footnote 17 below.

study are taken from a data set (supplied to us by Labour Canada) covering all stoppages (strikes and lockouts) in Canada between 1946 and 1980 that involved three or more workers and lasted more than half a day.[10]

The sample utilised covers strikes[11] between 1971 and 1980 in the manufacturing sector and is further restricted to strikes over wages occurring at the conclusion of an existing contract and the negotation of a new one. There are strong theoretical reasons for expecting the processes determining strike durations to differ considerably between the manufacturing sector and non-manufacturing, particularly service, industries. One of the main differences might reasonably be expected to be associated with the role of cyclical factors, a major concern in this paper. Similarly, the theoretical model for the duration of a strike occurring during the regular bargaining process at the end of the term of an agreement, and over the central issue of wages, is likely to need some modification to encompass strikes arising primarily over working conditions or because of inter-union disputes and strikes during the term of an agreement or prior to the first agreement.[12]

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10. Details of a certain amount of tidying-up of the original data required to yield a data set suitable for our analysis are available from the authors on request.

11. Lockouts are excluded from the analysis in this paper, as are rotating strikes.

12. Considerable evidence exists on the heterogeneity of strikes across issues and contract statuses. See, for instance, Kennan (1980) and Newby and Winterton (1983).

As mentioned above, the data are left-censored at one day. However, since the models are specified in terms of conditional settlement probabilities, we can simply measure elapsed duration relative to this censoring point. The sample considered here is further confined to strikes beginning before 28th March 1980, ensuring that our sample contains only strikes where the date of completion is known. Hence issues of right-censoring are not considered in this paper.

Factors influencing the conditional settlement probability

We now consider the various factors that the theoretical literature on strikes might lead us to expect would influence the conditional settlement probabilities, beginning with the primary consideration of our paper, the cyclical influence. An early suggestion of such an effect is provided by Ashenfelter and Johnson (1969), who use a model of strike duration to provide pointers for empirical work on the probability of a strike. In the model, the firm is viewed as choosing the length of the strike conditional upon the (known) shape of the workers' concession curve. A certain amount of criticism has, however, been directed at their theoretical work, in particular its basic assumption of irrational workers, and furthermore, the model provides no strong prediction of the direction of any cyclical impact on strike duration.[13]

More pertinent to the present study is the work of Kennan (1980) and Reder and Neumann (1980). The focus here is on the likelihood of a strike being settled, conditional on it being already underway, which precisely mirrors our empirical work. Kennan proposes the notion that the firm and the workers will display Pareto-optimal tendencies, so that strikes with higher total costs to the parties involved will be likely to be settled more quickly. This leads to the suggestion that strike duration will vary counter-cyclically: at the peak of the cycle, when the opportunity cost of a strike is highest, there is maximum incentive on both sides to settle quickly and thereby avoid the loss of the rent shared by the parties in the bilateral monopoly.

Such a view implies that the main cyclical influence will be the level of demand for the firm's products; in our empirical work, however, the cyclical variable is a macroeconomic one, and the strategy adopted for dealing with the possibility that there may also be industry-specific factors, representing differences between an individual

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13. There are actually a number of ways in which the business cycle may have an impact in their model, the combined effect of which is ambiguous. Ashenfelter and Johnson concentrate on only one, suggesting that the remaining factors need not be considered because the variables in question "change only slowly through time" (1969, p.40).

industry's cycle and the aggregate one, is to use the score tests outlined in the previous section to test for the omission of such factors in the form of dummy variables. The formulation is then modified as necessary.

Various opportunities exist, of course, for lowering the costs of a strike. Kennan (1980) mentions the role of inventories in this context, and the extent to which unfilled orders can be accumulated is another relevant factor. Both of these contain a strong industry-specific element, and we therefore regard them as being incorporated in the industry-specific factors discussed above.

Additional cyclical influences arising from the effect of labour-market conditions on the costs to striking workers are discussed by Tracy (1984), who views strikes as a learning mechanism necessitated by asymmetric information.[14] Like Kennan, Tracy notes that when an industry is doing well, strikes in that industry should be settled more quickly, but he then argues that favourable conditions in the local labour market will have the reverse effect since the workers' opportunity cost is lowered by the enhanced prospects for temporary employment.

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14. See Fudenberg, Levine and Ruud (1983), Hayes (1984), and Morton (1983) for other papers that model strikes along similar lines.

Turning next to variables other than cyclical ones that might have an impact on strike conditional settlement probabilities, it bears repeating that the literature on the economics of strikes does not yield a particularly rich harvest. We have, however, identified three variables whose inclusion in the x-vector has at least some validity, in addition to which there exists a further set of specific factors the importance of whose omission can be investigated with score tests.

Reder and Neumann introduced the notion of a "bargaining protocol", which defines, among other things, the issues that both sides recognise as open to negotiation. If, then, issues arise outside this protocol, this may delay settlement. Such a possibility is explored in our empirical work by the inclusion of a dummy variable, ISSD, indicating strikes involving secondary issues besides wage increases, and those primarily concerned with "wage adjustments" (e.g. overtime or incentive pay rates) rather than basic wage rates; the expectation is that the conditional settlement probabilities for these strikes will be lower.

A variable that some have argued may affect the total cost of a strike, and therfore its duration, is the size of the strike, that is, the number of workers involved.[15] In

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15. See Gunderson, Kervin and Reid (1984) for a number of reasons why this might be true.

our empirical work, the variable WS, which measures the number of workers on strike at its start, is included to test for this possibility.

Thirdly, we have also constructed a variable, STFR, measuring the strike start frequency (for the type of strike under analysis: those during contract renegotiation and involving wage issues) during the month in which each strike began. (To allow for the seasonality of contract expiration, monthly seasonal components are removed.) This is taken to measure, in part, the industrial relations climate at the time of the strike. Its effect on strike duration is, however, unclear; some argue that a spate of strikes might encourage worker resistance and prolong strike duration, but it could equally be asserted that, if this is the case, firms will recognise this and, behaving in much the same way as they do in Ashenfelter and Johnson's (1969) model, could deem it optimal to face a shorter strike.

Finally, there exists the possibility that our analysis may omit consideration of union-specific factors, for instance the varying strengths and resources of different unions, that could affect the cost of a strike to the workers. We therefore adopt the same strategy in this case that we proposed earlier for industry-specific factors, namely using score tests to test for omitted factors in the form of dummy variables, and modifying the formulation as necessary.

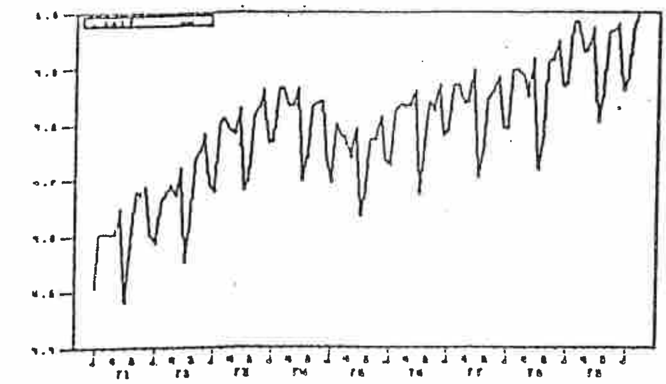
Measurement of the aggregate cycle

Finally in this section we turn to the identification of the aggregate business cycle. The graph in panel (a) of figure 1 shows a plot of the logarithm of the monthly index of industrial production in Canada (Statistics Canada CANSIM database number D144484) between 1971 and 1980. The cyclical pattern is obscured by the strong seasonal factors and general upward trend. The residuals from a regression on a set of monthly dummies and a linear time trend, shown in panel (b) of figure 1, are one possible measure of the cyclical component in the series over this period.

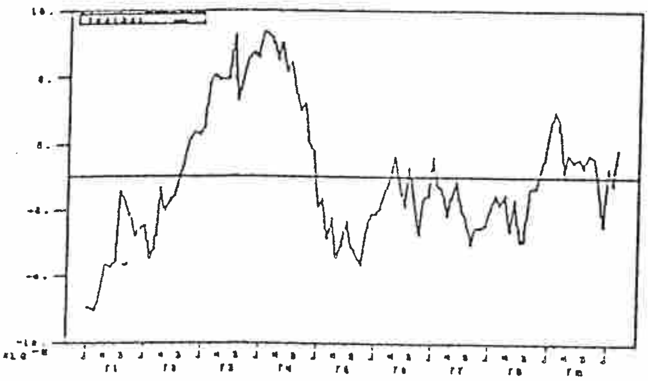
One of the problems surrounding this procedure concerns the nature of the detrending. Is it sufficient, for instance, to include only a linear trend? Some evidence contributing to an answer to this question is presented in panel (c) of figure 1, which portrays residuals from a regression of the variable plotted in panel (a) on seasonal dummies, time and time squared. Some differences in magnitude between the residuals in panel (b) and those in panel (c) are apparent, but it can be seen that the general patterns of the two plots are very similar.

A further problem that arises is that both sets of residuals are rather noisy. If, for example, we were interested in determing the turning points in these series

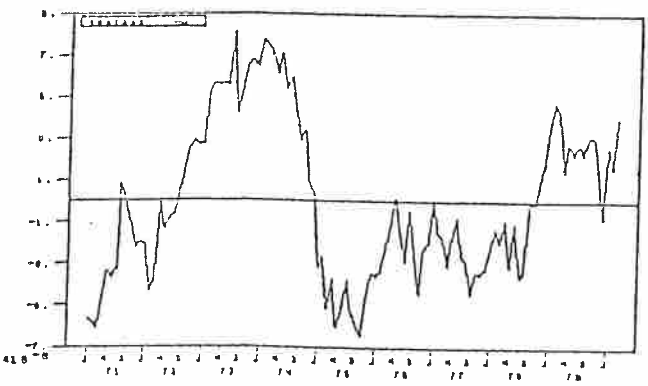
FIGURE 1



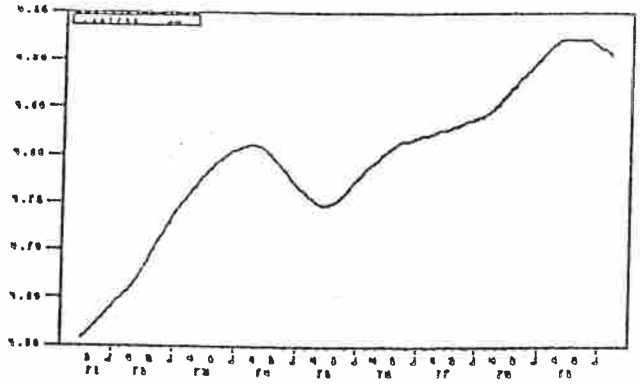
(a) logarithm of industrial production



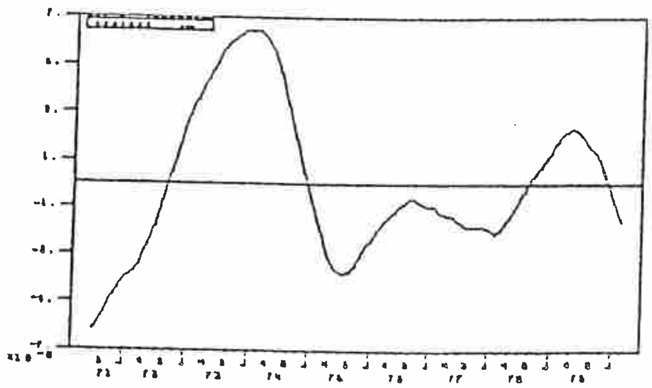
(b) residuals from regression of (a) on constant, seasonal dummies and time



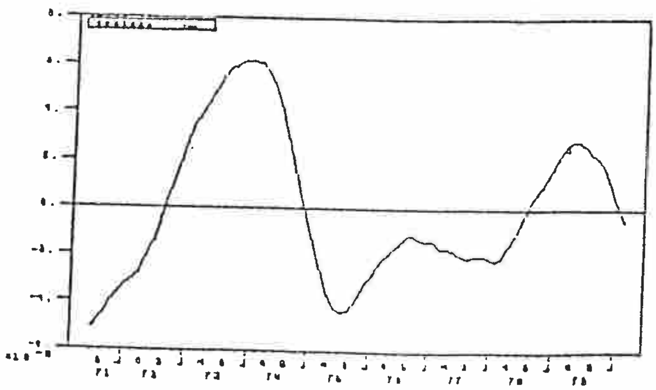
(c) residuals from regression of (a) on constant, seasonal dummies, time and time squared



(d) 12-month centred moving average of (a)



(e) residuals from regression of (d) on constant and time



(f) residuals from regression of (d) on constant, time and time squared precisely it is not at all clear that this would be a straightforward task. As an alternative, a smoothed version of the original series might be used instead. An obvious candidate is a moving average; selected appropriately, this has the added advantage of dealing in an alternative way with the strong seasonal pattern. Panel (d) of figure 1 illustrates a centred 12-month moving average of the series plotted in panel (a), panel (e) shows the residuals resulting from a regression of this moving average on a linear time trend, and panel (f) plots the residuals when time squared is added to the equation.

Several comments are apposite. First, panel (d) seems to confirm that the trend is basically a linear one. Secondly, either panel (e) or panel (f) offers a better basis than panels (b) and (c) for identifying some of the essential elements of the cycle, namely upswings, downswings, and turning points. Thirdly, differences in magnitude between the residuals in panels (e) and (f) are relatively minor, as they were when comparing panels (b) and (c).

In the light of this, the detrended moving average of the series (panel (e) of figure 1) is adopted as our main cyclical variable for the empiricial analysis in the next section; the issue is not viewed as being entirely resolved however, so we do also examine the question of how sensitive the results are to the selection of the cyclical variable.

IV. THE RESULTS

In this section we present the results of maximum-likelihood estimation of the two alternative models formulated in section II, using the sample of strikes starting during the period July 1971 to March 1980 inclusive (a total of 1429 strikes).[16] Although central attention is given to examination of the impact of the cyclical variable, the specification also incorporates three other exogenous variables, details of which are given in the preceding section. The strategy adopted, as outlined there, is to estimate the models with this basic x-vector and then construct score tests for inter-industry and inter-union variation and other omitted factors.

Table 1 presents results for the beta-logit model with this x-vector. The first column may be regarded as the base specification. A 4th-order polynomial in elapsed duration is used for $\theta(t)$. This was always accepted in tests against higher order polynomials. The score test against a 10th-order polynomial gave a $\chi^2(6)$-statistic of 9.7, and the additional term was insignificant when the model was estimated with a 5th-order polynomial.[17] The cyclical

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16. A modified Newton-Raphson algorithm, written by Mark Stewart, was used. Convergence, measured in terms of the number of iterations, was generally fairly quick.

17. The score and likelihood-ratio tests for the exclusion of a 5th-order term were in close agreement.

variable used, MIPT, is the centred 12-month moving average of the logarithm of the index of industrial production with trend removed. Its coefficient is positive and significant (asymptotic t-ratio = 5.8), implying that strike duration is indeed counter-cyclical.

Some summary statistics for assessing the impact of the cyclical variable are presented in table 2. The estimated impact on the settlement probability for the representative strike is quite considerable. If the strike starts at a time when the index of production is one standard deviation above trend, it has an estimated conditional settlement probability roughly 30% higher than if it starts when the index is one standard deviation below trend, with the result that in the latter case, the strike has an estimated expected duration approximately one-third longer. Alternatively, the estimated conditional settlement probability at the 1974 production peak is of the order of 70% higher, *ceteris paribus*, than at the initial low (in 1971), generating an estimated expected duration some 65% longer in the latter case.

The estimated individual conditional settlement probability for the representative strike is rising up to 99 days. (The corresponding fitted survivor function with all variables set to their means has fallen to .036 by then, i.e. such a strike has a 96.4% predicted probability of

TABLE 1

Maximum-Likelihood Estimates of Beta-Logit Model

Constant	-3.134 (0.116)	-3.281 (0.102)	-3.363 (0.118)	-3.331 (0.116)	-3.381 (0.118)
MIPT	4.488 (0.779)	4.232 (0.773)			
IPT			3.434 (0.656)		
MIPTSQ				4.586 (0.780)	
IPTSQ					3.592 (0.660)
ISSD	-0.214 (0.058)	-0.207 (0.058)	-0.216 (0.058)	-0.218 (0.058)	-0.223 (0.058)
STFR	-0.410 (0.164)	-0.353 (0.162)	-0.321 (0.165)	-0.367 (0.164)	-0.264 (0.166)
WS	0.296 (0.253)	0.277 (0.252)	0.278 (0.253)	0.301 (0.252)	0.290 (0.251)
t	0.354 (0.131)		0.352 (0.130)	0.353 (0.131)	0.350 (0.132)
t^2	-0.107 (0.038)		-0.106 (0.038)	-0.107 (0.038)	-0.106 (0.038)
t^3	0.0101 (0.0040)		0.0101 (0.0040)	0.0101 (0.0040)	0.0100 (0.0040)
t^4	-0.00029 (0.00014)		-0.00029 (0.00014)	-0.00029 (0.00014)	-0.00029 (0.00014)
α	0.0104 (0.0091)	0.0032 (0.0008)	0.0106 (0.0092)	0.0103 (0.0091)	0.0105 (0.0092)
log L	-6704.4	-6712.2	-6707.3	-6703.7	-6706.1

TABLE 1 (contd.)

NOTES

1. Asymptotic standard errors in parentheses.
2. Sample: strikes in manufacturing during contract renegotiation involving wage issues and starting during the period July 1971 - March 1980 inclusive. Sample size = 1429.
3. Variables: IPT = index of industrial production with seasonals and trend (T only) removed; IPTSQ = as IPT but with trend involving T and T^2; MIPT = centred 12-month moving average of index of industrial production with trend (T only) removed; MIPTSQ = as MIPT but with trend involving T and T^2; ISSD = issue not solely wage rate (dummy variable); STFR = strike start frequency for the month, deseasonalised ($\times 10^{-2}$); WS = number of workers on strike ($\times 10^{-4}$).
4. Coefficients on the t-polynomial are scaled by the mean duration of the sample.

TABLE 2

Summary Statistics for the Cyclical Effect in the Estimated Beta-Logit Model

MIPT=	initial 1971 low	1 st.dev. below trend	zero	1 st.dev. above trend	1974 peak
Initial hazard for representative strike=	1.9%	2.2%	2.5%	2.9%	3.3%
Hazard at 65 days=	2.6%	3.0%	3.5%	4.0%	4.5%
Mean duration=	43 days	39 days	34 days	29 days	26 days
Median duration=	32 days	29 days	25 days	22 days	19 days

NOTES

1. For the representative individual strike, defined by $v=\Pi(x)$, its mean,

   $$h(t) = 1 - \{1+\exp(x\beta+\theta(t))\}^{-1}.$$

   The survivor function is calculated using

   $$S(D)=\{1-h(D)\}\cdot S(D-1)$$

   and the mean duration is given by

   $$E = \sum_{D=0}^{\infty} S(D).$$

2. All other variables in the x-vector are evaluated at their means.

having finished by then.) The rate of change is, however, extremely slow. The conditional settlement probability is estimated to rise by about 4% of itself per week at the start. The highest it gets is 3.6%, compared with 2.5% at the start. This is outweighed by the variation in the initial conditional settlement probability over the cycle.

Strikes involving secondary issues in addition to wages or being primarily concerned with "wage adjustments" (ISSD=1) have significantly lower conditional settlement probabilities *ceteris paribus*. The estimated reduction for the representative strike is about 20%, with the expected duration being correspondingly roughly 20% longer. When these two groups were allowed different effects, the estimated coefficients were very similar in magnitude and insignificantly different from one another.

Conditional settlement probabilities are also lower, *ceteris paribus*, for strikes starting at a time of high strike incidence. The number of workers involved in the strike does not appear to have a significant effect on the conditional settlement probability however.

The heterogeneity parameter, α, is not significantly different from zero. The high degree of flexibility allowed in the form for the duration dependence swamps the omitted heterogeneity. In contrast, when this flexibility is suppressed and no allowance made for duration dependence (column 2), omitted heterogeneity takes over as the explanation of the non-constancy of the aggregate hazard function and α becomes significant. This illustrates the nature of the identification problem discussed earlier. The difference in the coefficient on the cyclical variable is however not that great. The likelihood-ratio test of the joint significance of the polynomial coefficients (i.e. against an alternative of no duration dependence) gives a $\chi^2(4)$-statistic of 15.6.

Columns 3 to 5 of table 1 examine alternative specifications of the cyclical variable. The variable used in column 4, MIPTSQ, is also based on the centred 12-month moving average, but the detrending includes a term in time squared. The difference in the estimated cyclical effect is negligible. Column 3 uses the variable that results from removing seasonals and a linear time trend from the logarithm of the original series (without the smoothing associated with the moving average), and column 5 that when a term in time squared is included in the detrending. The estimated coefficient is somewhat lower in each case, but the effect of a 2-standard-deviation shift is not dissimilar: 28% in the case of IPT as compared with 30% for MIPT. The variables based on the moving average dominate in likelihood terms. The additional variation in the IPT series does not therefore seem to be informative with regard to strike durations.

Table 3 presents the corresponding results for the gamma-proportional-hazards model. There are a number of differences worth commenting on. The estimated coefficients in the two models are of course not directly comparable. The factors to convert them to the implied proportional impacts on the settlement probability for the representative strike are, however, similar for the two models and thus the estimates for this model can be seen to be considerably larger than for the beta-logit model. In the case of the beta-logit model,

$$\partial h(t;x,\Pi(x))/\partial x = \beta \cdot h(t;x,\Pi(x)) \cdot \{1-h(t;x,\Pi(x))\},$$

while in the case of the gamma-proportional-hazards model

$$\partial h(t;x,0)/\partial x = \beta \cdot \{1-h(t;x,0)\} \cdot \log\{1/1-h(t;x,0)\}$$

$$\approx \beta \cdot \{1-h(t;x,0)\} \cdot h(t;x,0).$$

In both models, therefore, the coefficients are roughly equal to the derivatives of the logarithm of the representative hazard. Hence the implied proportional impact of the cycle on the conditional settlement probability of the representative strike is some 60% greater in the gamma-proportional-hazards model than in the beta-logit model.

Summary statistics for assessing the impact of the cyclical variable, comparable with those in table 2, are given in table 4. A strike starting at a time when the index of production is one standard deviation above trend has an estimated conditional settlement probability about 65% higher than that for a comparable strike starting when the index is one standard deviation below trend. This compares with a figure of about 30% for the beta-logit model. A strike starting at the 1974 production peak has an estimated settlement probability over double that of a comparable strike starting at the initial low in 1971.

A cubic in elapsed duration is accepted by the data when this model is used. The score test against a 10th-order polynomial gave a $\chi^2(7)$-statistic of 7.4, and the additional terms were not significant when the model was estimated with 4th- or 5th-order polynomials.[18]

The estimated individual settlement probability for the representative strike is rising throughout the range of interest. The estimated rate of increase is faster than for the beta-logit model: about 12% of itself per week at the

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18. The preferred specification for the beta-logit model included a 4th-order polynomial; in the case of the gamma-proportional-hazards model, however, not only was the 4th-order term insignificant but also the effect on the other coefficients, as compared with column 1 of table 3, was trivially small.

TABLE 3

Maximum-Likelihood Estimates of Gamma-Proportional-Hazards Model

Constant	-2.998 (0.186)	-3.213 (0.113)	-3.080 (0.189)	-3.022 (0.185)	-3.096 (0.189)
MIPT	7.253 (1.361)	4.775 (0.847)			
IPT			5.364 (1.123)		
MIPTSQ				7.584 (1.395)	
IPTSQ					5.976 (1.201)
ISSD	-0.391 (0.102)	-0.239 (0.063)	-0.385 (0.103)	-0.395 (0.102)	-0.397 (0.104)
STFR	-0.976 (0.298)	-0.456 (0.179)	-0.834 (0.300)	-0.918 (0.297)	-0.773 (0.303)
WS	0.355 (0.342)	0.301 (0.264)	0.355 (0.344)	0.353 (0.344)	0.359 (0.351)
t	0.999 (0.314)		0.983 (0.332)	1.030 (0.315)	1.066 (0.341)
t^2	-0.118 (0.032)		-0.116 (0.032)	-0.120 (0.032)	-0.122 (0.033)
t^3	0.00495 (0.00157)		0.00495 (0.00158)	0.00495 (0.00159)	0.00495 (0.00163)
σ^2	0.747 (0.259)	0.126 (0.026)	0.743 (0.277)	0.771 (0.260)	0.810 (0.284)
log L	-6696.4	-6709.1	-6700.7	-6695.1	-6698.3

TABLE 4

Summary Statistics for the Cyclical Effect in the Estimated Gamma-Proportional-Hazards Model

MIPT=	initial 1971 low	1 st.dev. below trend	zero	1 st.dev. above trend	1974 peak
Initial hazard for representative strike=	1.5%	1.8%	2.3%	3.0%	3.7%
Hazard at 65 days=	5.3%	6.4%	8.2%	10.4%	12.7%
Mean duration=	36 days	31 days	27 days	22 days	19 days
Median duration=	32 days	27 days	23 days	19 days	16 days

NOTES

1. For the representative individual strike, defined by u=0,

$$h(t) = 1-\exp\{-\exp(x\beta+\gamma_t)\}.$$

The survivor function is calculated using

$$S(D) = \{1-h(D)\}\cdot S(D-1)$$

and the mean duration is given by

$$E = \sum_{D=0}^{\infty} S(D).$$

2. All other variables in the x-vector are evaluated at their means.

start. The estimated settlement probability is initially 2.3% (compared with 2.5% for the beta-logit model) and at 65 days it has risen to 8.2% (compared with 3.5% for the beta-logit model).

Strikes involving secondary issues in addition to wages or being primarily concerned with "wage adjustments" (ISSD = 1) now have conditional settlement probabilities lower by about a third, and expected durations longer by about a third, than other comparable strikes.

Unlike α in the beta-logit model, the heterogeneity parameter, σ^2, is in this case significant in all specifications. The parameter is not, however, significantly different from unity, and for $\sigma^2 = 1$ the model is approximately equal to the homogeneous logit model (when the hazards are small).

The likelihood-ratio test of the joint significance of the coefficients in the duration-dependence polynomial gives a $\chi^2(3)$-statistic of 25.4. In this model suppression of the duration-dependence polynomial leads to a considerably greater underestimation of the cyclical effect than in the beta-logit model.

We now turn to the score tests on the models with the basic x-vector. The statistics for a number of tests on both models, with MIPT as the cyclical variable throughout, are presented in table 5. We consider first tests for omitted factors. The first row of the table tests for the possibility that initial individual settlement probabilities exhibit a significant seasonal pattern or trend after controlling for the components of the basic x-vector. The z-vector for the score tests contains 11 monthly dummy variables and a linear time trend constructed on the basis of the month in which the strike started. This test was conducted to examine the possibility that it might be inappropriate to remove seasonal components from the cyclical variable, since this seasonality could contribute to fluctuations in the total cost of a strike, and thereby to variation in conditional strike-settlement probabilities. As it turns out, however, both models produce $\chi^2(12)$-statistics less than the corresponding 5% critical point.

The second row of the table considers the possibility that MIPT on its own inadequately captures the cyclical effects relative to the other three variables used in table 1 and 3. Again the null hypothesis is accepted at the 5% level. In the third row, we test whether regionally diversified strikes take longer to settle, a variation on the proposition mentioned earlier that the complexity of the strike may lower the conditional settlement probabilities. The z-vector for the score tests contains dummy variables indicating strikes involving plants in more than one city or more than one province. Again both models produce test-statistics less than the corresponding 5% critical point.

TABLE 5

Score Test Statistics for Both Models with the Basic x-Vector

	k	$\chi^2(k)$ Beta-Logit Model	$\chi^2(k)$ Gamma-PH Model	5% Critical Point
Omitted Variables				
Seasonals and Trend	12	16.3	17.6	21.0
Other Cyclical Vars.	3	3.1	5.8	7.8
>1 City, >1 Province	2	3.8	3.8	6.0
Issue Effects	4	6.6	4.2	9.5
Industry Effects	12	47.2	40.2	21.0
Union Effects	10	25.4	21.9	18.3
2nd-order Misspecification				
α or σ^2 a Function of MIPT	1	2.7	0.3	3.8

NOTE

1. The statistics are based on the estimated models presented in table 1, column 1, and table 3, column 1.

The fourth row of the table tests against an alternative in which there is additional heterogeneity with respect to the issue(s) involved in the strikes. The sample used is restricted to strikes involving wage issues, and a dummy variable (ISSD) indicating strikes involving secondary issues and those primarily concerned with "wage adjustments" was included in the basic x-vector. Here we test for further heterogeneity within the ISSD = 1 category. The z-vector used contains four dummy variables indicating the issues of (i) "wage adjustments", (ii) working conditions, (iii) fringe benefits, and (iv) hours. In both models the additional heterogeneity with respect to issue is rejected.

The fifth row of the table tests against heterogeneity across industries. The sample was grouped into 13 broad industry groups.[19] The z-vector for the score tests accordingly contains a set 12 of dummy variables. In both models the null hypothesis of homogeneity is conclusively rejected. We return to this below when considering appropriate modification of the x-vector. In the sixth row heterogeneity across unions is examined. The z-vector contains dummy variables for each of the 10 unions with the

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19. The data identify 20 sectors within manufacturing. The reduction to 13 resulted from combining to ensure a minimum of 60 observations in each group.

highest frequency in the sample.[20] Again the null hypothesis of homogeneity is rejected in both models. Of course there is likely to be considerable overlap between the heterogeneity being indicated in these two cases, and we return to this point below also.

Finally we consider alternative hypotheses in which the omitted heterogeneity parameter (α in the case of the beta-logit model, σ^2 in the case of the gamma-proportional-hazards model) is a function of MIPT, a possibility discussed in section II. In both cases the score test statistics are less than the corresponding 5% critical point and the null of constancy is accepted.

Turning to the modification of the specification of the x-vector in the light of these results, we focus initially on the paper and printing industries. There are a number of <u>a priori</u> reasons why one might expect the pattern of strike durations to be different in this group of industries. On more pragmatic grounds, during the period under investigation the mean duration of this category of strike

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20. This turned out to be equivalent to including effects for any identifiable union involved in at least 30 strikes in the sample. The 10 unions are: the steelworkers, the autoworkers, the food and commercial workers, the Canadian paperworkers union, the teamsters, the metal trades federation, the machinists union, the carpenters, the woodworkers, and the united electrical workers.

was over 60% higher than in the rest of the manufacturing sector; furthermore the dummy for this industry group makes the largest contribution to the cross-industry-heterogeneity score test statistic.[21] As an initial re-specification we simply introduce a dummy variable for the paper and printing industry into the x-vector, both additively and interactively with MIPT.

Results for this re-formulation are presented in table 6, for both the beta-logit and the gamma-proportional-hazards models. The additive dummy variable has a significantly negative coefficient in both cases, indicating that the representative strike evaluated at the mean of the x-vector has a significantly lower settlement probability in the paper and printing industries, *ceteris paribus*. The interaction with MIPT is positive in both cases but only significantly so in the gamma-proportional-hazards model; settlement probabilities are, therefore, more cyclically-sensitive in the paper and printing industry, *ceteris paribus*. Finally the derivative with respect to PAPD is negative throughout the observed range of MIPT, implying that the representative strike has a lower settlement

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21. The mean duration in the paper and printing industries is 63.7 days compared with 38.8 days in the remainder of manufacturing. In the other industry groups used above the longest mean duration is 48.2 days.

TABLE 6

Maximum-Likelihood Estimates of Both Models with Re-formulated x-Vector

Beta-Logit Model

Sample	All	All
Constant	-3.339 (0.117)	-3.341 (0.118)
MIPT	3.623 (0.847)	4.170 (0.786)
ISSD	-0.212 (0.058)	-0.206 (0.058)
STFR	-0.270 (0.165)	-0.280 (0.165)
WS	0.368 (0.248)	0.337 (0.250)
PAPD	-0.438 (0.082)	-0.433 (0.082)
PAPD·MIPT	3.868 (2.219)	
α	0.0115 (0.0093)	0.0111 (0.0093)
log L	-6687.6	-6689.1
Sample Size	1429	1429

TABLE 6 (contd.)

Gamma-Proportional-Hazards Model

Sample	All	All	Excluding Paper and Printing	Paper and Printing Only
Constant	-3.027 (0.177)	-3.095 (0.166)	-3.164 (0.202)	-4.266 (1.012)
MIPT	5.509 (1.329)	6.466 (1.237)	5.858 (1.488)	21.960 (6.490)
ISSD	-0.408 (0.099)	-0.360 (0.091)	-0.424 (0.126)	-0.791 (0.438)
STFR	-0.775 (0.282)	-0.704 (0.265)	-0.597 (0.325)	-5.399 (1.547)
WS	0.476 (0.334)	0.411 (0.315)	0.676 (0.387)	-1.446 (1.432)
PAPD	-0.753 (0.145)	-0.690 (0.133)		
PAPD·MIPT	10.289 (3.999)			
σ^2	0.732 (0.201)	0.568 (0.169)	0.878 (0.389)	2.838 (0.939)
log L	-6675.9	-6680.0	-5738.9	-922.7
Sample Size	1429	1429	1245	184

NOTES

1. Estimates of the beta-logit model contain a 4th-order polynomial in elapsed duration, and those of the Gamma Proportional-Hazards Model include a 3rd-order polynomial.
2. The variables are as described in the notes to table 1, plus: PAPD = 1 for strikes in the paper and printing industry, = 0 for all other strikes.

probability in the paper and printing industries, _ceteris paribus_, at all points of the cycle.

Score test statistics for industry- and union-heterogeneity are presented in table 7 for both revised models. All four statistics are less than the corresponding critical points; once the effect of the paper and printing industries has been taken into account, the null hypothesis of homogeneity across the remaining industry groups and unions is accepted at the 5% level.

Turning to the estimated impact of the cyclical variable on the settlement probability of the representative strike, summary statistics for paper and printing and the remainder separately are presented in table 8. The estimated settlement probabilities are closest at the 1974 peak. They diverge markedly in the troughs. The mean duration of the representative individual strike starting at the 1971 low in the paper and printing industries is over twice the comparable figure for the remainder of manufacturing (in both models).

The cyclical effects on the settlement probabilities outside paper and printing are similar to those described earlier. In the paper and printing industries they are much more marked. in the estimated beta-logit model the hazard at the 1974 peak is over twice that at the 1971 low and the mean duration is nearly three times as long. These relative

TABLE 7

Score Test Statistics for Both Models with Modified x-Vector

	k	$\chi^2(k)$ Beta-Logit Model	$\chi^2(k)$ Gamma-PH Model	5% Critical Point
Omitted Variables				
Industry Effects	11	18.4	16.3	19.7
Union Effects	9	14.1	12.9	16.9
Other Interactions with PAPD	3	6.8	11.8	7.8
2nd-order Misspecification				
α or σ^2 a Function of MIPT	1	0.5	0.8	3.8

NOTE

1. Statistics are based on estimates with PAPD and PAPD·MIPT included in the x-vector.

figures are even larger in the case of the gamma-proportional-hazards model.

The score test for interactions between PAPD and the remaining elements of the x-vector (see table 7) accepts the null in the case of the beta-logit model but not in the case of the gamma-proportional-hazards model. Accordingly table 6 also presents, for this model, estimates on the two sub-samples. The restriction to the model with PAPD and PAPD·MIPT only is rejected. The likelihood-ratio test gives a $\chi^2(7)$-statistic of 28.6. The estimated cyclical effect in paper and printing is even larger than in the restricted model, but somewhat imprecisely determined.[22] The impacts of ISSD and STFR are also increased in the paper and printing industry. It would appear that, if anything, the restricted model understates both the cyclical sensitivity of settlement probabilities in paper and printing and the differences between this group and the rest of manufacturing.

V. SUMMARY

This paper has presented an empirical investigation of

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22. Because of the sample size, predicted hazards in the paper and printing industry on the basis of the separate equation are also likely to be somewhat unreliable and have not been calculated.

TABLE 8

<u>Summary Statistics for Cyclical Effects on the Representative Strike</u>

Beta-Logit Model (PAPD=0)

MIPT=	initial 1971 low	1 st.dev. below trend	zero	1 st.dev. above trend	1974 peak
Initial hazard=	2.1%	2.3%	2.6%	3.0%	3.3%
Hazard at 65 days=	3.0%	3.4%	3.8%	4.3%	4.7%
Mean duration=	39 days	35 days	32 days	28 days	26 days

Beta-Logit Model (PAPD=1)

MIPT=	initial 1971 low	1 st.dev. below trend	zero	1 st.dev. above trend	1974 peak
Initial hazard=	1.1%	1.3%	1.7%	2.2%	2.7%
Hazard at 65 days=	1.6%	1.9%	2.5%	3.2%	3.9%
Mean duration=	88 days	64 days	47 days	37 days	30 days

TABLE 8 (contd.)

Gamma-PH Model (PAPD=0)

MIPT=	initial 1971 low	1 st.dev. below trend	zero	1 st.dev. above trend	1974 peak
Initial hazard=	1.8%	2.1%	2.5%	3.0%	3.5%
Hazard at 65 days=	6.6%	7.7%	9.3%	11.1%	12.9%
Mean duration=	32 days	28 days	25 days	22 days	19 days

Gamma-PH Model (PAPD=1)

MIPT=	initial 1971 low	1 st.dev. below trend	zero	1 st.dev. above trend	1974 peak
Initial hazard=	0.4%	0.7%	1.2%	2.0%	3.2%
Hazard at 65 days=	1.7%	2.6%	4.5%	7.6%	11.8%
Mean duration=	72 days	56 days	41 days	29 days	21 days

NOTE

1. Statistics are based on estimates with PAPD and PAPD·MIPT included in the x-vector.

the cyclical variation in individual strike-settlement probabilities. Two contrasting flexible hazard models have been used and a number of useful score tests proposed. The main empirical findings are:

(1) there is strong support for the hypothesis that strike duration varies counter-cyclically, with conditional settlement probabilities at the peak of the cycle being up to twice as high, _ceteris paribus_, as at the trough (depending on the model used);

(2) the individual conditional settlement probability rises with elapsed duration;

(3) strikes involving secondary issues in addition to wages or being primarily concerned with "wage adjustments" have significantly lower conditional settlement probabilities, _ceteris paribus_;

(4) the cyclical sensitivity is even more marked in the paper and printing industries and individual conditional settlement probabilities there are much lower, _ceteris paribus_, than elsewhere in manufacturing. The hypothesis of homogeneity across industries and unions in the remainder of manufacturing is accepted by the data.

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