A THEORETICAL AND EMPIRICAL ANALYSIS OF AN ADVERSE SELECTION MODEL OF WAGES AND STRIKES

by

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ABSTRACT

Traditional views are that strikes are the result of mistakes in bargaining [Reder and Neumann (1980), Kennan (1980)] or that they are due to irrational or politically motivated behaviour [Ashenfelter and Johnson (1969)]. However, according to recently developed models of strikes based on the theory of asymmetric information, strikes are not merely mistakes, nor are strikes the outcome of irrational behaviour. The two-type adverse-selection models of Morton (1983) and Hayes (1984) and the sequential bargaining models of Fudenberg, Levine and Ruud (1984) and Tracy (1987) assume that the firm is better-informed than the union about the firm's level of profitability. Since a more profitable firm can afford to pay higher wages, it may be in the interest of the union to elicit information about the level of profitability from the firm. In the asymmetric information models, the union uses strikes to elicit the information from the firm.

One of the contributions of the present dissertation is to propose a variant of Hayes' adverse-selection model which may be generalized to an arbitrary number of firm types. Two key results are derived. First, wages and strikes are predicted to satisfy what is called a monotonicity property. Second, there are two kinds of equilibria in the model: a separating equilibrium, which includes the possibility that a strike is the outcome of negotiations, and a pooling equilibrium, which rules out the possibility of a strike. In addition, comparative statics results are derived for all the exogenous variables in the theoretical model, significantly extending the results of Hayes (1984).

Without exception, existing empirical work examining the asymmetric information theory of strikes estimates separate reduced form equations for strike incidence, strike duration and, to a lesser degree, wage levels. According to the asymmetric information theory, however, wage levels and strike lengths are jointly endogenously determined given values for exogenous variables. A major contribution of the present dissertation is the derivation of a maximum likelihood procedure which allows for joint estimation of strike duration and wage level equations. This econometric model has several attractive features. First, the estimating equations are interpreted as linear approximations to the structural equations of the theoretical model. Second, the likelihood function is derived using the monotonicity property of wages and strikes predicted by the theoretical model. Third, the model allows the data to estimate which observations correspond to separating equilibria and which to pooling equilib-
ria. Fourth, the estimates may be used to calculate the probability of a separating equilibrium at each observation. Finally, it is argued that since the endogenous variables in many models of adverse-selection exhibit a monotonicity property, the econometric model proposed here potentially has a much wider application than simply as a model of wages and strikes.

Another contribution of the present dissertation is the assembly of a data set which merges information from 2,459 contracts in Canadian industries covering the period 1964–86 with data specific to the firm. Explanatory variables implied by the theoretical model are then used to test the comparative statics predictions in the econometric model of wages and strikes. The empirical results confirm all but one of the comparative statics predictions of the theory, and the econometric model fits the data well. Pooling equilibria are estimated to be present in approximately 13% of the contracts. In addition, the results are entirely consistent with the stylized facts concerning the behaviour of strikes over time, the business cycle and relative to other variables.
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CHAPTER I

Introduction, Review of the Literature and Overview

1.1 Introduction

One of the most puzzling aspects about collective bargaining between unions and firms is the occurrence of strikes. Since a strike makes both sides worse off, it is difficult to understand why rational agents would allow them to occur. Until recently, theoretical models concluded either that strikes are mistakes [Reder and Neumann (1980), Kennan (1980)], or that strikes are due to irrational or politically motivated behaviour [Ashenfelter and Johnson (1969)]. Lacking rigid theoretical guidelines, empirical research on the subject of strikes is largely a hotchpotch of intuitive guesses about the relevant explanatory variables. This research has proved very useful at determining several “stylized facts” concerning the behaviour of strikes, but much less useful at isolating the determinants of the economic function of a strike.

At the heart of the matter has been the absence of a satisfactory theoretical model of strikes. Recently, however, such a theoretical model has been proposed based upon the theory of asymmetric information. According to the asymmetric information theory, strikes are not merely mistakes, nor are strikes the outcome of irrational behaviour. In the asymmetric information approach, strikes are the outcome of rational behaviour in a world of private information—strikes are used as an information-revealing device.

The two-type adverse-selection models of Morton (1983) and Hayes (1984) and the sequential bargaining models of Fudenberg, Levine and Ruud (1984) and Tracy (1987) assume that the firm is better-informed than the union about the firm’s level of profitability. Since a more profitable firm can afford to pay higher wages, it may be in the interest of the union to elicit information about the level of profitability from the firm. According to the asymmetric information models, the union uses strikes to elicit the information from the firm.

The Hayes (1984) model is limited to two firm types—high profit and low profit—and the models of Morton (1983), Fudenberg, Levine and Ruud (1984) and Tracy (1987) assume that the level of employment at the firm is fixed. One of the contributions of the present dissertation is to propose a variant of Hayes’ adverse-selection model which may be generalized to an arbitrary number of firm types. In addition,
comparative statics results are derived for all the exogenous variables in the theoretical model proposed below, significantly extending the results of Hayes (1984).

The relevance of the asymmetric information models to the observed behaviour of strikes will, to a large extent, depend on empirical tests of the theory. Owing to the substantial theoretical contributions of the asymmetric information approach, there is a rapidly growing empirical literature concerned with testing the implications of the theory. Without exception, the empirical work in this area which has been accomplished to date concentrates on the separate estimation of reduced form equations for strike incidence, strike duration and, to a lesser degree, wage levels. A major contribution of the present dissertation is the derivation of a procedure which allows joint estimation of (approximate) structural equations for strike duration and the wage level, using restrictions implied by the theoretical model.

It is argued below in chapter 4 that the econometric model of wages and strikes is applicable to the estimation of any model based on the theory of adverse-selection. The claim is based on the recognition that the endogenous variables in all models of adverse-selection exhibit what is called a monotonicity property. It turns out that the monotonicity property is a crucial step in deriving the likelihood function for the joint estimation of the structural equations. Thus the econometric model proposed below is, in principle, applicable to a much wider class of models than simply a model of wages and strikes.

In order to estimate the econometric model and be consistent with the theory, it is desirable to have a large number of observations on collective agreements together with the relevant firm-specific data. Another contribution of the present dissertation is the assembly of a data set which merges information from 2,459 contracts in Canadian industries covering the period 1964–86 with data specific to the firm. Explanatory variables implied by the theoretical model are then used to test the comparative statics predictions.

The remainder of the chapter is concerned with motivating the above claims regarding the contributions of the dissertation. In section 1.2 there is a review of the theoretical and empirical literature on the topic of strikes, with special attention paid to work that is related to the asymmetric information theory of strikes. Section 1.3 provides an overview of the dissertation including a brief discussion of the theoretical model, the empirical implementation of the theory and the estimation results.
1.2 A Survey of the Literature

The theoretical and empirical literature on strikes is extensive and is comprehensively covered in a recent survey article by Kennan (1986). The literature survey presented here is focused on theoretical and empirical contributions to the study of strikes which are directly related to the asymmetric information models.

Section 1.2.1 discusses theoretical models of strikes. It contains a detailed discussion of the asymmetric information models of strikes as well as an analysis of other theoretical strike models. Section 1.2.2 examines the empirical literature on strikes which is related to the asymmetric information models. The section includes a detailed discussion of empirical tests of the asymmetric information theory of strikes carried out to date. The section also contains an overview of other empirical studies of strikes that are related to the present study.

1.2.1 Theoretical Models of Strikes

All of the current theoretical models of strikes have their antecedent in the work of Hicks (1932). The problem posed by Hicks is to provide a rigorous theoretical explanation of the economic function of a strike. It has proved a difficult problem for economic theory to solve. For, given perfect information, it is difficult to understand why a rational firm and a rational union would engage in an event which makes them both worse off.

As a first step, Hicks examined the trade-offs faced by the firm and the union between the wage level and the length of a strike, under the assumption that the costs of continuing the strike increase as the strike progresses. Hicks postulates the existence of a union concession schedule which relates the minimum acceptable wage increase for union members to the duration of a strike. As costs increase with strike duration, Hicks argues that the union becomes more willing to accept successively smaller wage increases as the strike progresses. In other words, the union’s concession schedule is negatively sloped. The firm is hypothesized to have a resistance schedule\(^1\) which is a relationship between the maximum acceptable wage increase the firm is willing to pay and the duration of a strike. Again, because costs increase with strike

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\(^1\) Hicks actually uses the term “concession schedule” to refer wage and strike trade-offs faced by the firm and the term “resistance schedule” to refer to the union’s trade-offs. Subsequently, however, the terms have been switched between the firm and the union. I have chosen to follow the more recent convention.
duration, the firm is willing to grant higher wages as the strike progresses. Thus the firm’s resistance curve is positively sloped. A strike occurs in Hicks’ model when the union’s initial wage demand is greater than the firm’s initial wage proposal. In these circumstances, the equilibrium wage level and strike length outcome is determined at the intersection of the firm’s resistance schedule and the union’s concession schedule.

The most obvious criticism of the Hicks model is that it predicts a particular wage increase and strike duration outcome. Since the outcome is known in advance, the firm and union can simply agree to the wage predicted by the model and, in so doing, avoid the costs of a strike. Furthermore, if the firm and union agree to the wage level in advance, Hicks’ model ceases to be a model of strikes. Hicks eventually concludes that strikes are the result either of the union trying to maintain a reputation for toughness, or of mistakes arising because of the existence of private information on the part of at least one of the bargaining agents.

Despite the Hicks Paradox, the basic idea that contract negotiations involve the firm resisting union demands and the union making concessions is intuitively appealing. Perhaps this explains why work subsequent to that of Hicks focuses on the union’s concession schedule and the firm’s resistance schedule. In Ashenfelter and Johnson (1969), the concession schedule of the union is given a central role. Ashenfelter and Johnson argue that there are three, rather than two, parties involved in contract negotiations: the firm, the union leaders, and the rank and file of the union. The rank and file are assumed to have a negatively sloped concession schedule, reflecting the assumption that wage expectations are revised downward over the course of a strike. The role of the union leaders is to convey the shape of the concession schedule to the firm, and to convey the extent of feasible wage demands to the rank and file. If the rank and file has very high expectations, the union leaders may agree to a strike rather than suffer the politically damaging consequences of accepting a wage increase far below the expectations of the rank and file.

There is no firm resistance curve in the Ashenfelter and Johnson model. The firm simply maximizes its present value taking the union’s concession schedule as given. The equilibrium wage and strike length outcome is determined either, in the case of an interior solution, at a tangency point between a firm isoprofit curve and the union concession schedule, or at a boundary point, in which case the optimal strike length

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2 This is called the “Hicks Paradox” by Kennan (1986).

3 The reputation argument is sometimes called the “rusty weapons hypothesis.”
Eaton (1972) examines a model similar to that of Ashenfelter and Johnson, but instead he concentrates on the role of the firm’s resistance schedule. In Eaton’s model, it is the firm which has expectations of the wage the union will accept, and these expectations are revised upward as the strike progresses. The union takes the firm’s resistance curve as given and maximizes utility, thereby determining a wage and strike outcome. As for the Ashenfelter and Johnson model, an interior solution to the union’s maximization problem results in a positive strike length and a boundary solution results in no strike. It is interesting to note that Eaton finds—comparing the firm’s final pre-strike wage offer to the eventual wage settlement—some strikes result in a present value gain for the union. In some simple sense, therefore, his results indicate that strikes do “pay” the union in the same sense that price wars “pay” the firm in the entry-deterrence model of Shubik (1959).

A well-known criticism of the Ashenfelter and Johnson model is that the concession schedule of the rank and file is not based on any maximizing behaviour. If the rank and file were to behave rationally there would be no concession schedule in the first place; the rank and file could achieve its most preferred wage simply by having a concession schedule independent of the strike length. To be fair, Ashenfelter and Johnson recognize this point, claiming that their model focuses on the political, rather than the economic, functions of the strike. Thus the Ashenfelter and Johnson model fails to address the central issue, namely the economic function of strikes.

Following Hicks, Reder and Neumann (1980) and Kennan (1980) focus on the notion that strikes are mistakes. The presumption of this approach is that strikes impose joint costs on the union and the firm. Since some strikes impose greater joint costs than others, it is in the interest of the firm and union to minimize the possibility of a mistake occurring when joint strike costs are high. The possibility of mistakes is reduced by an implicit long-term agreement or protocol which covers contingencies where joint strike costs are high. Thus the central prediction of the model is that strikes are less likely when they are more costly.

The Reder and Neumann model is not based on rigorous theoretical arguments. For example, the existence of a protocol is merely assumed rather than derived from an underlying model of optimizing behaviour. It is not clear how the protocol deals

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4 A similar criticism applies to Eaton’s model since the firm’s resistance schedule is not based on maximizing behaviour by the firm.
with instances where the joint costs of a strike are high but where the costs fall mostly on one party. Furthermore, if a particular situation is not covered by a protocol, there is no reason why a strike should occur. Simply stating that strikes are mistakes does little to address question of the economic function of strikes. Indeed, in the context of the Reder and Neumann model, the strike has no economic function. Finally, the prediction that strikes are less likely in circumstances where they are more costly would presumably be found in any economic model of strikes.

The theoretical models of Ashenfelter and Johnson and Reder and Neumann provide little insight into the original problem posed by Hicks. Despite this limitation, the Ashenfelter–Johnson model and the joint cost theory have fostered a significant body of empirical work. Yet, given the absence of rigorous theoretical models, it is not clear that the empirical results teach us a great deal about the underlying causes of strikes. As pointed out by Koopmans (1947), measurement without theory serves little more than as a description of the observed data.

Recent work has tackled the theoretical problem of strikes by assuming, as suggested by Hicks, that there is an asymmetry of information between the union and the firm. The asymmetric information approach applies rigorous theoretical techniques to models where both the firm and the union behave rationally. The adverse-selection models of Hayes (1984) and Morton (1983) and the sequential bargaining models of Tracy (1987) and Fudenberg, Levine and Ruud (1984) assume that the firm has more information than the union about the firm’s level of profitability. If the firm is highly profitable it can afford to pay high wages, so it may be in the union’s interest to elicit the information of the firm. According to the asymmetric information approach, the union uses strikes to gain the firm’s information.

The asymmetric information approach potentially represents an important step toward understanding the economic function of strikes. In contrast to the Ashenfelter and Johnson model, both the parties in contract negotiations are assumed to behave rationally in the asymmetric information models. Contrary to the joint cost theory, strikes are not mistakes in the asymmetric information models. Instead, strikes serve as an information revealing device. The more satisfactory theoretical foundations of the asymmetric information approach furnish a rich set of empirically testable hypotheses.

The asymmetric information models predict that wage levels are inversely related
to strike durations.\(^5\) Second, the models predict that the incidence of strikes is inversely related to the level of profitability or demand at the firm. In a model where the level of employment is fixed, Tracy (1987) shows that strikes are positively related to the degree of uncertainty the union has about the firm’s level of profitability, and positively related to the level of wages outside the firm. In a model where the level of employment is variable, it is shown in chapter 2 below that strikes are positively related to the degree of dispersion in the firm’s level of demand, while the relationship of strikes to the level of wages outside the firm is ambiguous.

Without exception, the asymmetric information models of strikes examine the case where the firm possesses private information. As a consequence of the information structure, it is the union which makes the bargaining proposals to the firm, the firm’s role being merely to accept or reject the proposals. Kennan (1986) has criticized the asymmetric information approach to collective bargaining on the grounds that requiring the union to make the proposals to the firm arbitrarily assigns a bargaining advantage to the union. Although aimed at asymmetric information models of collective bargaining, Kennan’s criticism may be applied to a whole class of models with one-sided asymmetric information. In the literature on the principal-agent problem, for example, where it is the agent which possesses the private information, the asymmetry of information becomes irrelevant if the agent is allowed to make the proposals. In other words, the assumption of one-sided private information requires, per force, the assignment of a bargaining advantage to the party without the information. A model in which the asymmetry of information is two-sided may be able to circumvent the problem, but such models are still in their infancy.

Hayes and Morton characterize collective bargaining under asymmetric information as a problem of adverse-selection, while Tracy and Fudenberg et al follow a game-theoretic approach by incorporating asymmetric information into a sequential bargaining game between the union and the firm. In the adverse-selection approach, the firm is potentially one of two “types;” one type having a high level of demand for its output, the other having a low level of output demand. The asymmetry of information arises because the firm knows which of the two types it is, whereas the union only knows the firm is one of two types. The union proposes a combination of wage levels and strike lengths to maximize its expected utility across firm type.

\(^5\) If strikes were mistakes, there would be no reason to expect them to be systematically related to wage levels.
The firm then chooses one of the union's wage-strike proposals to maximize profits. The union is able to induce the firm to reveal which type it is by making two proposals: one consisting of a high wage and a strike of zero length, which the firm will choose if it is the high demand type; the other consisting of a low wage and a strike of positive length, which the firm will choose if it is the low demand type. As the proposals are the result of optimizing behaviour by the union, and since the firm chooses the proposal which maximizes its profit, the chosen proposal is efficient *ex ante*. However, if the firm is the low demand type, the proposal chosen by the firm will include a strike. In this instance, since both the firm and the union are better off foregoing the strike, the low wage-strike proposal is inefficient *ex post*. Thus the validity of the adverse-selection approach depends critically upon whether it is credible for the union to commit to a series of wage and strike proposals some of which may be inefficient *ex post*.

To understand why commitment to potentially inefficient proposals is credible clearly requires a more complex approach than the static adverse-selection models of Hayes and Morton. Consider a model which takes into account the long-term association between the union and the firm by recognizing that bargaining takes place at repeated intervals. At each bargaining interval, suppose the union makes two proposals, one consisting of a low wage and a strike the other consisting of a high wage without a strike. When viewed in isolation, this strategy appears irrational because the firm knows that if it chooses the proposal with a strike that it is not in the union's interest to carry the strike through.

However, suppose there is an additional asymmetry of information present in that the firm is doubtful about the union's options, motivation, and behaviour. In particular, the firm may not be entirely certain that the union is behaving as a rational agent. In these circumstances, Kreps and Wilson (1982) and Milgrom and Roberts (1982) have shown that commitment to an inefficient outcome in the short run is credible because, in so doing, the union establishes a reputation which provides long run gains. Thus the threat of a strike is credible because it allows the union to establish a reputation for "toughness" which yields high wages in the future offsetting the immediate losses from a strike.6

The sequential bargaining models of Tracy and Fudenberg *et al* provide an

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6 Of course, the idea that strikes are used to maintain a reputation for toughness is nothing new. Hicks' so-called "rusty weapons hypothesis" embodies exactly this notion.
alternative approach to the asymmetric information theory of strikes. In these models, the union has a prior distribution summarizing its beliefs about the firm’s profits and bargaining takes place in a sequence of “rounds.” Based on the prior beliefs about profits, the union makes an initial wage proposal which the firm may either accept or reject. As these models assume a fixed time period between proposals, rejection of the first wage proposal results in a strike. The strike continues as subsequent wage proposals are rejected, ending only when the firm accepts one of the proposals. Owing to the assumption of the fixed time period between proposals, strikes occur in these models as much because of the rules of the game as because of the asymmetry of information. Furthermore, Gul, Sonnenschein and Wilson (1986) show that as the time period between proposals becomes arbitrarily small, the probability of a strike occurring goes to zero.

Thus both the adverse-selection and the sequential bargaining models are vulnerable to fairly fundamental criticisms. However, as with the Kennan criticism noted above, these criticisms also apply to a much wider class of models, and indeed these problems are unresolved in the related literature. For example, the ex post inefficiency problem of the adverse-selection models is related to the issue of ex post renegotiation discussed recently by Dewatripont (1988). Similarly, the fixed-time-between-offers problem of the sequential bargaining models is part of the more general topic of delays in bargaining discussed in Gul and Sonnenschien (1985) and Admati and Perry (1987).

A particularly attractive feature of the sequential bargaining models is their inherent dynamic nature. This more closely approximates the intuitive view of the strike as a learning process whereby both the firm and the union acquire knowledge about the other agent’s preferences over the course of time. In contrast, the adverse-selection approach is inherently static. The union makes its offers at a point in time, and the firm immediately selects one of the offers. Thus any learning which does take place is essentially instantaneous. This does not conform with intuition. The dynamic nature of the sequential bargaining models also, it turns out, imposes a limitation. The game theoretic paradigm of the sequential bargaining models restricts union preferences to depend only on the wage level by assuming that the employment level is fixed. Conversely, the adverse-selection model of Hayes (1984) allows employment

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7 In the sequential bargaining models of Tracy and Fudenberg et al, of course, it is only the union doing the learning as the asymmetry of information is one-sided.
at the firm to vary. The approach taken in the present paper is to build on Hayes' model and thus allow for more general union preferences.

1.2.2 Empirical Models of Strikes

Econometric Models of Aggregate Strike Behaviour

The empirical analysis of strike behaviour essentially follows one of two approaches. The first, and until recently the most prevalent, type of analysis is performed at the aggregate or macro level. The earliest systematic multivariate regression analysis of this type is contained in Ashenfelter and Johnson (1969), where U.S. quarterly data on the number of strikes for the period 1952:1 to 1967:2 is examined.

As noted above, the union's concession schedule in the Ashenfelter-Johnson model is not based on any assumption of maximizing behaviour by the rank and file. Hence, the empirical content of the model comes from intuitive guesses about the determinants of the concession schedule. For example, Ashenfelter and Johnson hypothesize that the rank and file concede more rapidly when the unemployment rate is high. This seems reasonable because it is consistent with rational behaviour by the rank and file. But it is assumed in the theoretical model that the rank and file is not behaving rationally. Thus the empirical implementation of the Ashenfelter and Johnson model is inconsistent with the theoretical assumptions.

Ashenfelter and Johnson (1969) estimate a model where the dependent variable is the frequency of strikes in each quarter. The independent variables include the unemployment rate, the lagged rate of change in the nominal wage, the lagged rate of change in consumer prices and seasonal dummy variables. Their results indicate that the frequency of strikes is negatively related to the unemployment rate (i.e., procyclical), negatively related to the lagged rate of nominal wage changes, positively related to the lagged inflation rate and greater in the spring, summer and autumn relative to the winter months.

Despite the inconsistency inherent in the empirical implementation of the Ashenfelter and Johnson model, their approach remains the standard to examining the behaviour of strikes at the aggregate level. Indeed, the Ashenfelter and Johnson empirical results are consistently replicated in studies using data from different countries and time periods. Some examples are Abbott (1984) who uses Canadian data, Kaufman (1982) who examines U.S. data for the period 1900–77, and Pencavel (1970) who examines U.K. strike data.
Micro-econometric Models of Strikes

With the increasing availability of individual contract data, attention has shifted from aggregate studies of strikes to the behaviour of strikes at the firm level. Farber (1978) adapts the Ashenfelter-Johnson model to examine contract data from 10 bargaining pairs in U.S. manufacturing industries, covering 80 contracts and 21 strikes over the period 1954–70. Farber derives maximum likelihood estimates of the parameters of the union's concession schedule, which are modelled as linear functions of exogenous variables. Most of the coefficient estimates are not statistically significantly different from zero based on a one-tailed test. Thus Farber's results appear to be hampered by the relatively small number of observations. He does find that the union's concession rate is significantly reduced by labour's share of total sales and significantly increased by wage guidelines in effect during the period 1962–66. Farber’s study obviously suffers from the same inconsistency noted for the Ashenfelter and Johnson (1969) study. Furthermore, the limited cross-section of firms in his sample may not be representative of the population of firms in the manufacturing industries.

Recently, there have been a number of studies of strike incidence using Canadian contract data. Swidinsky and Vanderkamp (1982) examine data on 1,641 contracts in the manufacturing industries for the period 1967–75; Cousineau and Lacroix (1986) model the incidence of strikes in 1,871 contracts, also for manufacturing industries, and; Gunderson, Kervin and Reid (1986) study strike incidence in 2,437 private sector contracts for the period 1971–83. It is useful to examine these studies for two reasons. First, the empirical analysis contained in chapters 3 and 5 below use a similar data set to those listed in the above three studies. Second, these three papers are representative of the approach used in the majority of empirical studies of strike incidence. The approach is to use a vaguely specified theoretical model to justify the choice of explanatory variables in a strike probability equation. Due to the absence of rigorous theory, the choice of explanatory variables is ad hoc and the empirical results difficult to interpret.

Consider the following example. Gunderson, Kervin and Reid (1986) claim that a principal determinant of strikes is uncertainty, misinformation and divergent expectations. The variable chosen to proxy for this uncertainty is the level of employment

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8 Namely, the intercept, the concession rate and the firm's discount rate.
9 Once again, the influence of Hicks' work is evident.
growth at the firm. According to their "theory," expanding firm employment increases the level of uncertainty etc., causing an increase in the incidence of strikes. Elsewhere in their analysis, the joint cost theory of strikes is invoked. Thus strike activity and the joint costs of a strike to the union and the firm are hypothesized to be inversely related. The problem is that growing firm employment could be interpreted as a signal of rising joint costs—firm costs rise because demand is growing, and union costs rise because employment is growing. However, if joint costs are rising, employment growth would lead to a reduction in strike incidence, i.e., contrary to the effect predicted for the rising level of uncertainty. The contradiction arises in part because there is no theoretical model to provide a consistent way of choosing explanatory variables. Furthermore, the \textit{ad hoc} choice of explanatory variables necessitated by the lack of underlying theory means it is difficult to interpret the estimated coefficients since it is not clear what effects they are measuring.

Cousineau and Lacroix (1986) maintain that strikes are due to imperfect information, and that strikes are less likely when they are more costly. A major drawback of the joint cost theory is that it fails to specify a mechanism which transforms individual costs into joint costs. In particular, the predictions of the theory are unspecified for instances when strike costs are high for one of the bargaining parties and low for the other. This leaves the empirical researcher two options: (i) variables which increase the cost of strikes to one party and decrease them to the other may simply be ignored; (ii) such variables are included in the analysis with the implicit assumption that there exists some unspecified mechanism to allocate the costs among the parties. Cousineau and Lacroix choose the second option. However, as the implicit mechanism in the joint cost theory remains unspecified, it could be argued that Cousineau and Lacroix are not actually testing any theoretical model at all. At the very least, it is difficult to place the estimation results of Cousineau and Lacroix (1986) in a consistent framework. The extreme form of the \textit{ad hoc} approach to modelling strike incidence is represented by Swidinsky and Vanderkamp (1982). Here, the authors make no pretence about the existence of an underlying theoretical model. They are content to explain the estimated coefficient signs after the regressions have been calculated.

Although the \textit{ad hoc} approach does little to further our understanding of the economic function of strikes, it does provide a useful summary description of the data. Swidinsky and Vanderkamp (1982), Cousineau and Lacroix (1986), and Gunderson, Kervin and Reid (1986) all report a significant positive impact of bargaining unit size
on the probability of a strike. Wage controls, as administered by the Anti-Inflation Board during the period 1975–1978, are found to reduce the probability of a strike. Swidinsky and Vanderkamp and Gunderson et al report substantial inter-industry and seasonal variation in the propensity to strike. Both studies find relatively little variation in the propensity to strike across regions in Canada. Mirroring the results of Ashenfelter and Johnson (1969), who use aggregate data, all three of the Canadian studies find that strike incidence is negatively related to the unemployment rate, i.e., that strikes are procyclical.

Studies of strike incidence using micro contract data sets from the U.S. are found in Mauro (1982), Gramm (1986), McConnell (1987a) and Card (1988). Mauro postulates that strikes occur because each of the bargaining agents habitually misspecify its opponent's preferences. He tests the implications of this hypothesis using a logit model of strike incidence and data from 149 contracts covering 14 U.S. bargaining pairs for the period 1952–77. Although Mauro's assumption that the agents consistently misrepresent each others' preferences is untenable, his empirical results indicate procyclical strike incidence and a negative relationship between profits and strike incidence.

Gramm (1986) examines data from 1,050 contracts in U.S. manufacturing during the period 1971–80 using a probit model of strike incidence and a tobit model of strike duration.\(^\text{10}\) Explanatory variables proxy the expected costs to the firm and the expected benefits to the union of a strike. Notably, Gramm finds that neither the unemployment rate nor the inflation rate over the previous contract have any significant effect on strike incidence or duration.

Card (1988) is purely an econometric study\(^\text{11}\) of the probability of a strike. Card uses panel data from contracts for 253 bargaining pairs in U.S. industries for the period 1955–79. The use of panel data allows Card to employ cross-sectional and longitudinal estimation techniques to examine the effect of endogenous contract characteristics on the probability of a strike, and the variation of strike activity over time within bargaining pairs. His results indicate that strike incidence is higher in the summer and autumn relative to the winter and spring. In addition, strike incidence is increased by the length of time between negotiations and is significantly affected

\(^{10}\) Since strikes do not occur in the majority of contract negotiations, i.e., most strikes have a zero duration, the distribution of observed strike lengths is truncated at zero.

\(^{11}\) That is, as opposed to a model based on a specific "theory."
by lagged strike outcomes.

Finally, there are two noteworthy econometric studies of strike duration contained in Kennan (1986) and Harrison and Stewart (1987). Kennan estimates the conditional settlement probability of a strike (the hazard function) using data from 576 contract strikes in U.S. manufacturing for the period 1968–76. The level of industrial production is estimated to have a significant positive effect on the conditional settlement probability of a strike. That is, the results indicate that strike duration is countercyclical. Harrison and Stewart (1987) estimate a hazard function with data from 4,531 strikes during the period 1946–83 in the Canadian manufacturing sector. Using several different measures of the cycle in industrial production, the authors find strong support for the hypothesis that conditional strike settlement probabilities are procyclical, and hence that strike duration is countercyclical.

Micro-econometric Models of Strikes Based on Asymmetric Information Theory

The first empirical study of strikes based on the theory of asymmetric information is contained in Fudenberg, Levine and Ruud (1984). They examine U.S. data on 159 contracts in manufacturing industries for the period 1955–79. The authors attempt to estimate reduced form equations for the endogenous variables in their model; the wage level, strike length and firm sales. Contrary to a theoretical prediction, they find that the level of sales is significantly lower in firms for which there was no strike during contract negotiations. A major drawback with the Fudenberg et al study is the small sample size.

Studies by Tracy (1986, 1987) examine data from 1,319 contracts in U.S. industries for the period 1973–77. Both the Tracy papers estimate a logit model of strike incidence and a conditional strike settlement probability equation. Tracy (1986) uses the rate of return on the firm’s stock as a proxy for the level of profitability at the firm. The standard deviation of the stock return for the year ending 6 months prior to contract negotiations is used as a proxy for the degree of uncertainty facing the union. Tracy finds that the rate of return on the stock does not significantly affect strike incidence or duration. In support of the asymmetric information model, he finds that the variability of stock returns significantly increases strike incidence and duration.

Tracy (1987) uses three variables to proxy for the firm’s level of profitability: worker experience, the concentration ratio in the firm’s industry, and the trend

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12 The data are an extended version of the Farber (1978) data set.
in industry demand. The union's degree of uncertainty about the firm's level of profitability is proxied by the standard deviation of excess returns from a capital-asset pricing model securities equation. In support of the theory, he finds that the standard deviation of excess returns significantly increases strike incidence, while worker experience and the trend in industry demand significantly decrease strike incidence. Only the worker experience variable significantly affects strike duration, and the effect is negative as predicted by the theory.

The two Tracy studies suffer from the limited time-series of the contract data. In fact, McConnell (1987b, p.9) reports that Hirtle (1986) is not able to replicate the result reported in Tracy (1987) that the standard deviation of excess returns increases strike incidence, and Hirtle's sample covers a longer time period, 1957-80. Another problem with the Tracy studies is the large number of competing independent variables. This makes it difficult to sort out precisely what effects the variables measuring. For example, Tracy (1986) employs measures of both the industry and regional unemployment rate to proxy labour market conditions facing union members.

The studies by McConnell (1987a, 1987b) touch on several empirical aspects of the asymmetric information theory of strikes. McConnell (1987a) assumes that the firm's private information enables it to predict future industry demand conditions better than the union. Thus the difference between realized industry prices and price forecasts are used to proxy for the firm's private information. She finds that neither the level nor the variance of the residual from the price forecasting equation has any significant effect on strike incidence or duration. The major problem with these results is that the firm's private information is proxied by industry level data. McConnell (1987b) concentrates on the negative relationship between wages and strikes predicted by the asymmetric information theory. The contract data contain 1,986 observations taken from U.S. industries during the period 1970-81. McConnell essentially estimates the union's concession schedule, finding that the schedule has a negative slope but that the slope is relatively small: the real wage decreases by only 3% after a strike lasting 100 days. McConnell does not include any firm-specific independent variables in the estimation.

Card (1987) also examines the contemporaneous relationship between wages and strike using 2,258 contracts on the Canadian manufacturing sector for the period 1964-85. He finds no significant correlation between the real contract wage and either strike incidence or strike duration, contrary to the predictions of the theory. Since
Card makes no attempt to control for firm-specific information, it is not regarded as a complete test of the asymmetric information models.

It is useful to summarize the empirical results so far obtained for the asymmetric information theory of strikes. With the exception of the study by Fudenberg, Levine and Ruud (1984), empirical models which have explicitly included firm-specific variables to proxy for the firm’s level of profitability and the union’s uncertainty about the firm’s profitability are, by and large, in agreement with the predictions of the theory. Studies which examine the relationship between the wage level and strike duration present much less evidence in favour of the asymmetric information model of strikes. However, the studies which examine this relationship fail to control for firm-specific information despite the fact that the essence of the asymmetric information approach is the hypothesis that firm-specific information plays a crucial role in the determination of wage and strike outcomes.

Existing tests of the asymmetric information theory of strikes represent what shall be called the reduced form approach. The reduced form approach is simply concerned with the specification and estimation of separate strike incidence, strike duration and wage equations. In this regard, the empirical work based on the asymmetric information theory of strikes follows precisely the same methodology as earlier empirical studies of strikes. The reduced form approach is useful for measuring the impact of key exogenous variables, such as the firm’s level of profitability, on wage levels and strike incidence. In this approach few (if any) priors about the structure of the wage and strike determination process are imposed on the data.

In contrast, the empirical approach presented in the present dissertation recognizes that, according to the asymmetric information approach, wage levels, strike incidence and strike lengths are jointly endogenously determined in the model, given values for the exogenous variables. The (approximated) structural equations for the level of wages and strike lengths derived from the theoretical model are jointly estimated by the method of maximum likelihood. Furthermore, the likelihood function is derived using a restriction implied by the theoretical model. Thus, in contrast to the reduced form approach, the full theoretical model is brought to bear on the data in the sense that the estimated wage and strike equations satisfy the restrictions suggested by the theory.

\footnote{Another recent example of this is found in Herrington (1988).}
1.3 Overview

The next chapter formally introduces the adverse-selection model of firm and union collective bargaining. It is assumed that there are two firm types; one facing a high level of output demand, the other facing a low level of demand. The union's problem is to make a series of wage and strike offers to the firm which maximizes expected utility across firm type. It is shown that there are two sorts of equilibria in the model: (i) the union makes two wage-strike offers—a high wage offer with no strike and a low wage offer with a positive strike—such that the union elicits the firm’s information (a separating equilibrium), or; (ii) the union makes a single wage offer with no strike, in which case the union does not elicit the firm’s information (a pooling equilibrium). It is shown that strikes occur only in low demand firm types. Given the various properties of the utility and profit functions, comparative statics results are derived for all the exogenous variables in the model. Many of these results have not previously appeared in the literature. For example, it is shown that strikes are more likely in firms with a greater degree of variability in output demand.\(^\text{14}\) Chapter 2 concludes with a summary of the theoretical results and a brief discussion about empirical implementation of the model.

Chapter 3 discusses the assembly of the collective agreement and firm-specific data set used to test the model and the choice of the variables used to proxy the exogenous variables in the adverse-selection model. Firm revenue is used as the proxy for the level of demand at the firm, and the coefficient of variation of firm revenue is used to proxy for the variability in the firm’s demand. To judge the performance of the adverse-selection explanatory variables in a conventional empirical setting, a reduced form logit model of strike incidence is estimated. The results confirm the predictions of the theory; strike incidence is negatively correlated with the level of firm revenue, and positively correlated with the variability in revenue.

Chapter 4 introduces the main econometric model of wages and strikes. The model has three particularly attractive features. First, the estimating equations are linear approximations to the first-order conditions derived in the theoretical model. Second, following the theoretical model, the econometric specification allows the occurrence of separating and pooling equilibria to be endogenously determined.

\(^{14}\) Furthermore, it turns out that the results from the two-type model may be generalized to the case where the number of firm types is arbitrary.
Third, the likelihood function used to derive coefficient estimates is obtained using a property of the endogenous variables implied by the theoretical model.

Chapter 5 contains the results from the estimation procedure proposed in chapter 4. The model is found to fit the data relatively well, especially for the wage equations. In all but one instance, the coefficient signs on the explanatory variables are as expected. The estimates indicate that the high wage offer is more sensitive to the bargaining environment than the low wage offer. In this way, the coefficient estimates are in accordance with the "stylized facts" that strikes are procyclical and that strike activity is reduced during the period the Anti-Inflation Board was in effect.
CHAPTER II

An Adverse-Selection Model of Collective Bargaining

2.1 Introduction

Before the theoretical model is introduced, a remark is warranted on the relationship of the adverse-selection model to "real-world" collective bargaining. Typically, contract negotiations are thought of as beginning with the union making a series of "demands" to the firm. These demands may include guarantees for employment levels, standards for working conditions, increased or reduced hours of work and a host of other conditions. In the majority of negotiations, however, it is surely the level of wages and employment that is central to the union's demands. In the model presented below, both the wage level and the level of employment are constituents of negotiations. This restriction is consistent with the observation that employment and wages are the most important part of the union's demands and it simplifies the model considerably.

Suppose, for the sake of argument, that the union is demanding higher wages. Typically, the firm's response is to claim that it cannot afford to pay higher wages. However, the firm may give in to the union rather than, for example, be faced with a strike. Thus, from the union's point of view, the issue is whether the firm's claim of inability to pay higher wages is credible. If not, the union may decide to call the firm's bluff, or in other words, the union may attempt to induce the firm to reveal whether it is willing to pay the higher wages. This is precisely the type of behaviour that the adverse-selection model addresses.

The next section of the chapter formally introduces the adverse-selection model of collective bargaining. During the course of the discussion the model is related to the work of Hayes (1984) and Morton (1983). The section derives the complete set of comparative statics properties of the model and concludes with an arithmetic example. The last section of the chapter discusses empirical implementation of the model, thereby linking the theoretical model with the estimation to follow.

2.2 An Adverse-Selection Model of Collective Bargaining

Collective bargaining takes place between a union, which represents a fixed homogeneous pool of workers, and a firm which hires some of the union workers
in order to produce output. Labour is assumed to be the only factor of production. The asymmetry of information between the union and the firm takes the following form. The firm is one of two types indexed by \( i = 1, 2 \). The difference between types is captured by a single parameter, \( \theta \), which reflects the level of demand for the firm’s output. The union does not know the actual firm type, \( \theta_i \), but it does know that \( \theta_i \in (\theta_1, \theta_2) \). Without loss of generality, firm types are indexed in terms of their level of demand, so \( \theta_2 > \theta_1 \). The union has a subjective probability, \( q \), that the firm is the type represented by \( \theta_2 \). Thus, \( (1 - q) \) is the union’s subjective probability that the firm is type \( \theta_1 \).

Collective bargaining is characterized in the following way. At time \( t = 0 \), the union presents the firm with a schedule of offers. Each offer in the schedule of offers comprises a wage level, \( w \), and a strike duration, \( s \). The firm is free to choose one wage-strike offer from the schedule of offers. For the sake of argument, let \( (\hat{w}, \hat{s}) \) be the wage-strike offer chosen by the firm. The outcome of the firm’s choice is a strike of duration \( \hat{s} \) after which the firm chooses the level of employment, \( L \), subject to \( \hat{w} \). Thus the union unilaterally sets the wage at which the workers are hired and the firm unilaterally sets the level of employment given the wage. This corresponds to the so-called monopoly model of collective bargaining. At \( t = \hat{s} \), the profit function of firm type \( \theta_i \) is defined

\[
\pi(\hat{w}, \theta_i) \equiv \max_L \{ \theta_i p[f(L)]f(L) - \hat{w}L \},
\]

(2.1)

where \( \theta_i p \) is the inverse demand function for the type \( \theta_i \) firm, and \( f \) is an increasing, concave, twice continuously differentiable production function. It is assumed that \( p \) and \( f \) possess the requisite properties to ensure that \( \pi \) is three-times continuously differentiable. The demand curve in (2.1) is assumed to be downward-sloping. In the industrial relations literature, the firm is usually assumed to have some degree of market power in order that there exist rents for the firm and union to bargain over. Of course, a downward sloping demand curve does not guarantee the existence of rents; a firm in monopolistic competition earns zero profit in the long-run, for example. In such cases, it is assumed that there exist rents available to each specific bargaining

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1 The results for the theoretical model presented below generalize to the case where the number of firm types is arbitrary. However, as the empirical analysis is restricted to two types (see the discussion in section 2.3 below) and in the interest of brevity, only the two-type model is discussed here.

2 To this point the set up of the model is the same as Hayes (1984).
pair, i.e. quasi-rents, for the firm and union to bargain over. For further discussion, consult Tracy (1987), Morton (1983) and the paper by Hayes.

**Lemma 1. Properties of the firm profit function \( \pi \):**

1. \( \pi \) is non-decreasing in \( \theta \), non-increasing in \( w \) and homogeneous of degree 1 in \( w \) and \( \theta \).
2. \( \pi \) is convex.
3. Labour demand, \( L(w, \theta) \), has the following properties:
   
   (i) \( L(w, \theta) = \frac{-\partial \pi(w, \theta)}{\partial w} = -\pi_w(w, \theta) \);
   
   (ii) \( L_w(w, \theta) < 0 \);
   
   (iii) \( L_{\theta}(w, \theta) > 0 \).

**PROOF.** Property 1 is obvious from the definition of \( \pi \). Convexity of \( \pi \) is shown as follows. Let \( (w, \theta) \) and \( (w', \theta') \) be two pairs of wage and demand level combinations, and let \( L \) and \( L' \) be the respective profit maximizing input levels. Define \( w'' = tw + (1-t)w' \) and \( \theta'' = t\theta + (1-t)\theta' \) for any \( 0 < t < 1 \). Now

\[
\pi(w'', \theta'') \equiv \max_L \{ [t\theta + (1-t)\theta']p(f(L)) f(L) - [tw + (1-t)w']L \} \\
= [t\theta + (1-t)\theta']p(f(L'')) f(L'') - [tw + (1-t)w']L'' \\
= t\theta p(f(L'')) f(L'') - twL'' + (1-t)\theta' p(f(L'')) f(L'') - (1-t)w'L'' \\
\leq \pi(tw, t\theta) + \pi((1-t)w', (1-t)\theta')) \\
= t\pi(w, \theta) + (1-t)\pi(w', \theta')
\]

since \( L'' \) is not necessarily the profit maximizing input level at \((w, \theta)\) or \((w', \theta')\). Property 3 (i) follows from Hotelling's lemma. Property 3 (ii) follows since \( L_w(w, \theta) = -\pi_{ww}(w, \theta) < 0 \) since \( \pi \) is a convex function. Property 3 (iii) is demonstrated thus. The first-order condition for (2.1) implies \( w/\theta = pf'(1 + \eta) \), where \( \eta \) is the elasticity of demand. To maintain this equality when \( \theta \) increases, \( f' \) must decrease. This in turn implies that, since \( f \) is concave, \( L \) must increase. The differentiability of \( L(w, \theta) \) follows from the assumption that \( \pi \) is twice continuously differentiable. 

Consider firm profits from the wage-strike offer \((\bar{w}, \bar{s})\) at time \( t = 0 \). During the strike, i.e., from \( t = 0 \) to \( t = \bar{s} \), no output is produced at the firm so profits are zero. After the strike, i.e., from \( t = \bar{s} \) onwards, firm profits are given by (2.1). Therefore,
the discounted present value of firm profits from the wage–strike offer \((\hat{w}, \hat{s})\) at \(t = 0\) is given by
\[
\Pi(\hat{w}, \hat{s}, \theta_i) \equiv \int_0^{\hat{s}} 0 \cdot e^{-rt} \, dt + \int_{\hat{s}}^{\infty} \pi(\hat{w}, \theta_i) \cdot e^{-rt} \, dt
\]
\[
= \frac{\pi(\hat{w}, \theta_i) e^{-r\hat{s}}}{r},
\]
where \(r, 0 < r < 1\), is a discount rate. The discount rate is assumed to be fixed throughout the analysis. Thus, taking the natural logarithm of (2.2) and ignoring the constant term \(\ln r\), the discounted present value of profits at \(t = 0\) is written
\[
\Pi(\hat{w}, \hat{s}, \theta_i) \equiv \ln \pi(\hat{w}, \theta_i) - r\hat{s},
\]
It is more convenient to work with II rather than \(\Pi\) since the former is a quasilinear function of \(w\) and \(s\). From (2.3), the marginal rate of substitution (measured as a positive number) for firm type \(\theta_i\), given the wage offer \(\hat{w}\) and the strike offer \(\hat{s}\), is written
\[
\text{MRS}(\hat{w}, \hat{s}, \theta_i) = \frac{-r \pi(\hat{w}, \theta_i)}{\pi_w(\hat{w}, \theta_i)}
\]
where the second equality follows from property 3 of the firm profit function. Hence isoprofit curves for the type \(\theta_i\) firm in \((w, s)\) space are negatively sloped. In addition, the slope of the isoprofit curves is independent of \(s\). A family of isoprofit curves for firm type \(\theta_i\) is shown in Figure 2.1, where the arrow indicates the direction of increasing profits.

As in the Hayes model, union utility is modelled as the utility of a representative union member. Employment is allowed to vary in this model, so union utility from the offer \((\hat{w}, \hat{s})\) at time \(t = \hat{s}\) and thereafter depends on the wage and the level of employment. Since the firm unilaterally sets the level of employment, union utility from the offer \(\hat{w}\) and firm type \(\theta_i\) is defined
\[
u(\hat{w}, \theta_i) \equiv \tilde{u}(\hat{w}, L(\hat{w}, \theta_i)).
\]
Thus \(u(\hat{w}, \theta_i)\) is the level of utility at the wage level \(\hat{w}\) and labour demand \(L(\hat{w}, \theta_i)\). Following Hayes (1984), it is assumed that \(\tilde{u}\) is concave, \(\tilde{u}_w > 0, \tilde{u}_L > 0\) and \(\tilde{u}_{wL} > 0\). Two examples of union utility functions that have been proposed in the literature
Figure 2.1
A Family of Isoprofit Curves for Firm Type $\theta_i$. 

\[ \Pi (\hat{w}, \hat{s}, \theta_i) \]
and that meet Hayes' restrictions are; the Stone-Geary form [Dertouzos and Pencavel (1981)] defined as

\[ u(w, \theta_i) = (w - \bar{w})^\alpha [L(w, \theta_i) - \bar{L}]^{1-\alpha}, \quad 0 \leq \alpha \leq 1, \quad (2.6) \]

and the form proposed by MacDonald and Solow (1981)

\[ u(w, \theta_i) = L(w, \theta_i)[V(w) - V(\bar{w})], \quad (2.7) \]

where \( \bar{w} \) is the wage available to union members outside the firm, \( \bar{L} \) is a "subsistence" level of employment, and where \( V \) is an increasing, concave and twice differentiable function. The wage-bill and rent-maximization specifications of union preferences are both included as special cases of (2.6) and (2.7).

**Lemma 2.** Union utility, \( u \), is an increasing function of \( \theta \).

**Proof.** This follows from Property 3 (iii) of the firm's profit function, and the assumption that \( \bar{u}_L > 0 \). Thus, from (2.5), \( u_\theta(w, \theta) = \bar{u}_L L_\theta > 0 \). The differentiability of \( u \) follows from the differentiability of \( \bar{u} \) and \( L(w, \theta) \).

In fact, \( u \) is concave in \( w \) provided that \( L(w, \theta) \) is concave or not "too" convex, since from (2.5)

\[ u_{ww}(w, \theta) = \bar{u}_{ww} + \bar{u}_w L_w + \bar{u}_L L_{ww}. \]

Since \( \bar{u}_{ww} < 0 \) and \( \bar{u}_w L > 0 \) by assumption and \( L_w < 0 \) by Property 3 (iii) of Lemma 1, the first two terms are unambiguously negative. Therefore, \( u_{ww}(w, \theta) < 0 \) provided \( L_{ww} \) is not "too" positive. For example, if labour demand is linear then \( L_{ww} = 0 \) and the concavity of \( u \) is guaranteed. Figure 2.2 illustrates the derivation of the utility function \( u \) for the case where labour demand is linear.

The level sets of the function \( \bar{u}(w, L) \) are shown in the upper panel of Figure 2.2 along with labour demand for the two firm types. The utility functions given by (2.5) are illustrated in the lower panel of Figure 2.2 for the two firm types. For labour demand \( L(w, \theta_2) \) in the upper panel, union utility is maximized at \( w_2^* \) as illustrated in the lower panel. Similarly, union utility is maximized at \( w_1^* \) for labour demand \( L(w, \theta_1) \). Since \( L(w, \theta_2) > L(w, \theta_1) \) for all \( w > 0 \), \( u(w, \theta_2) > u(w, \theta_1) \) for all \( w > 0 \).

Note that \( \bar{w} \) is not the wage available to union members during a strike; wages during a strike are zero. Instead, \( \bar{w} \) is the wage available if workers quit the firm.
Figure 2.2
Union Utility Functions
That is, at every wage, utility is greater if the firm is type $\theta_2$ because employment is larger if the firm is the high demand type. Finally, it is reasonable to assume that $L_{u\theta} > 0$, or in other words, that labour demand is more steeply sloped for the high demand firm, i.e., as illustrated in Figure 2.2. Hence, labour demand for the high demand firm type is assumed to be less responsive to the wage than labour demand for the low demand firm type. In this case, $u_{w\theta} = \bar{u}_{u\theta} L_{\theta} + \bar{u}_L L_{u\theta} > 0$ since all the terms on the right-hand side are positive. This result will be important for determining a comparative static property of the model.\footnote{See Proposition 3 below.}

During the period $t = 0$ to $t \neq \hat{s}$, wages and employment at the firm are zero. Thus it is assumed that union utility is zero during the strike. From $t = \hat{s}$ onwards, union utility is given by (2.5). Therefore, the discounted present value of union utility for the wage–strike offer $(\hat{w}, \hat{s})$ and firm type $\theta_i$ at time $t = 0$ is simply

$$U(\hat{w}, \hat{s}, \theta_i) = \int_0^{\hat{s}} 0 \cdot e^{-rt} dt + \int_{\hat{s}}^{\infty} u(\hat{w}, \theta_i) e^{-rt} dt$$

which is analogous to equation (2.2) for the firm. Again the denominator, $r$, in (2.8) is ignored, so union utility is simply

$$U(\hat{w}, \hat{s}, \theta_i) = u(\hat{w}, \theta_i) e^{-r\hat{s}}. \tag{2.9}$$

The union knows that the firm is either type $\theta_1$ or type $\theta_2$, it just does not know which type. Basically, the union has two options. Either the union can attempt to get the firm to reveal its type, or the union can ignore firm type altogether. It turns out that both these options may be written down in the same problem. Suppose the union decides to induce the firm to reveal its type. Thus the union designs the schedule of offers so that if the firm is type $\theta_1$ it will choose one particular offer, and if the firm is type $\theta_2$ it will pick the other. Without loss of generality, the offer intended for the type $\theta_1$ firm may be written $(w_1, s_1)$ and the offer intended for the type $\theta_2$ firm may be called $(w_2, s_2)$. In order for firm type $\theta_1$ to pick offer $(w_1, s_1)$ it must be that profits for firm type $\theta_1$ are at least as great from the offer $(w_1, s_1)$ as from the offer $(w_2, s_2)$. Similarly, to induce firm type $\theta_2$ to pick the offer $(w_2, s_2)$ it must be
that profits for firm type $\theta_2$ from the offer $(w_2, s_2)$ are at least as great as profits from the offer $(w_1, s_1)$. In other words, to induce the firm to reveal its type the schedule of offers $(w, s) = \{(w_1, s_1), (w_2, s_2)\}$ must satisfy the following constraints

$$\Pi(w_1, s_1, \theta_1) \geq \Pi(w_2, s_2, \theta_1),$$
$$\Pi(w_2, s_2, \theta_2) \geq \Pi(w_1, s_1, \theta_2).$$

(2.10)

Of course, the union may not wish to induce the firm to reveal its type. In this case, the union may set $w_1 = w_2$ and $s_1 = s_2$, and the constraints in (2.10) are satisfied as equalities. Equations (2.10) are called self-selection or incentive-compatibility constraints.

There may be any number of offers which satisfy (2.10), but the union is interested only in the schedule of offers that maximizes expected utility. The expected utility to the union of a schedule of offers $(w, s) = \{(w_1, s_1), (w_2, s_2)\}$ is given by

$$(1 - q) U(w_1, s_1, \theta_1) + q U(w_2, s_2, \theta_2).$$

(2.11)

Substituting for the utility functions from (2.9), expected utility is written

$$(1 - q) u(w_1, \theta_1) e^{-\tau s_1} + q u(w_2, \theta_2) e^{-\tau s_2}.$$  

(2.12)

The union's maximization problem may now be stated formally.

**Problem 1.** Choose a schedule of offers $(w, s)$ to maximize expected utility (2.12) subject to the self-selection constraints (2.10).

Notice in the statement of Problem 1 that there are no constraints which guarantee the firm a non-negative level of profits. Such individual-rationality or participation constraints, as they are referred to in the literature, are common in other models of adverse-selection. Since the firm unilaterally sets the level of employment in the present model, it can always guarantee that profits are non-negative. Thus the participation constraints are always satisfied, and may be safely ignored. In contrast, Morton (1983) finds that the participation constraint is binding for the low firm type in a model where employment is fixed.

---

5 Remember that $q$ is the union's subjective probability the firm is type $\theta_2$.

The model presented here is essentially the same as Hayes (1984), with one or two exceptions. Hayes assumes that the strike and the period following the strike cover a fixed time interval. Hence the longer a strike lasts, the shorter the production period. In the present model, there is no restriction on the amount of time the strike and the production period may take up. Also, there is no discounting of utility and profits in the Hayes model. Most important, care has been taken in the model here to derive the underlying properties of the utility and profit functions. This allows the derivation of comparative statics results not found in Hayes' paper.

Problem 1 is simplified with the help of an additional assumption. From (2.4) it is clear that the marginal rate of substitution for a given wage–strike offer is simply the discount rate times profit per worker. Consider the following result.

Lemma 3 If the elasticity of labour demand, \( L(w, \theta) \), is less than or equal to 1, then profit per worker, \( \pi(w, \theta)/L(w, \theta) \), is an increasing function of \( \theta \).

PROOF. It is required that

\[
\frac{\partial}{\partial \theta} \left[ \frac{\pi(w, \theta)}{L(w, \theta)} \right] = \frac{\pi_\theta(w, \theta)}{L(w, \theta)} - \frac{\pi(w, \theta) L_\theta(w, \theta)}{L(w, \theta)^2} > 0.
\]

Multiplying both sides by \( L(w, \theta)^2 > 0 \) it is required that

\[
\pi_\theta(w, \theta)L(w, \theta) - \pi(w, \theta) L_\theta(w, \theta) > 0.
\]

Using the envelope theorem it is straightforward to show that \( \pi_\theta(w, \theta) = pf \). By substitution it follows that

\[
pf L - (\theta pf - w L) L_\theta > 0
\]

is the sufficient condition. Rearranging terms and dividing through by \( L > 0 \)

\[
pf \left[ 1 - \frac{\theta L_\theta}{L} \right] + wL_\theta > 0.
\]

Since \( L_\theta > 0 \), the inequality follows if

\[
\frac{\theta L_\theta}{L} \leq 1,
\]

where the expression on the left-hand side is the elasticity of demand for labour with respect to a change in \( \theta \).
Clearly, the elasticity condition to ensure that profits per worker are increasing in \( \theta \) is much stronger than required. However, because it is simple to formulate and seems intuitively plausible it is a maintained assumption. The assumption plays an important role in the analysis. Under the elasticity assumption it is guaranteed that the marginal rate of substitution is an increasing function of \( \theta \). This implies that isoprofit curves for the high demand firm type are more steeply sloped than isoprofit curves for the low demand firm type given any wage–strike offer. In the idiom of the adverse-selection literature, firm isoprofit curves are said to exhibit the *single-crossing property*. The interpretation of the property is that the opportunity cost of a strike, measured in terms of forgone revenue, is greater for the high demand firm type. The property may also be interpreted as stating that the high demand firm type is willing to pay larger wage increments for a given reduction in strike length. Representative isoprofit curves for the two firm types at the wage–strike offer \((w, s)\) are illustrated in Figure 2.3.

The single-crossing property of firm isoprofit curves implies that schedules of offers which satisfy the self-selection constraints (2.10) possess important properties. These properties are summarized in Proposition 1.

**Proposition 1.** If a schedule of offers satisfies the self-selection constraints then:

(i) \( w_1 \leq w_2 \) and \( s_1 \geq s_2 = 0 \)

with \( w_1 < w_2 \) and \( s_1 > 0 \) if \( (w_1, s_1) \neq (w_2, s_2) \),

and;

(ii) \( \Pi(w_2, s_2, \theta_2) = \Pi(w_1, s_1, \theta_2) \).

Thus wage offers are non-decreasing in type, strike offers are non-increasing in type and no strike is offered to the high demand firm type. As this result plays an important role immediately below and in the empirical implementation of the model, it is called the *monotonicity property*. Secondly, if the two firm types are offered different wages, the low demand firm type must be offered a positive strike duration. Finally, only the self-selection constraint on the high demand firm type is binding. In the terminology of the adverse-selection literature, this is known as a binding *adjacent downward* self-selection constraint.

The monotonicity property is a general result in models of adverse-selection, see Cooper (1984) or Matthews and Moore (1987) for example, so it is stated here
Figure 2.3
Single-Crossing Property of Firm Isoprotit Curves
without proof. Hayes (1984) finds that the inequalities in the monotonicity property are reversed for the case where firm isoprofit curves are more steeply sloped for the low demand firm type and where wages behave like a Giffen good in the union's utility function. This result is not considered further in the present paper for two reasons. First, it seems counter-intuitive to suppose that firm types facing a high level of demand have a lower opportunity cost of a strike, which must be the case for the type \( \theta_1 \) firm to have more steeply sloped isoprofit curves. Second, the requirement the wages behave like a Giffen good in the union's utility function is unlikely to be empirically verified. In this context, it should be noted that the single-crossing property is derived from a restriction on the firm's profit function in the present model,\(^7\) whereas the property is a maintained assumption in Hayes' (1984).

The results contained in Proposition 1 allow Problem 1 to be simplified. From Proposition 1 it is known that \( \Pi(w_2, s_2, \theta_2) = \Pi(w_1, s_1, \theta_2) \). Also from Proposition 1 it is known that \( s_2 = 0 \). Thus, substituting for \( s_2 \) and using the definition of the profit function in (2.3), the binding self-selection constraint may be written

\[
\ln \pi(w_2, \theta_2) = \ln \pi(w_1, \theta_2) - rs_1,
\]

which is rearranged to yield

\[
e^{-rs_1} = \frac{\pi(w_2, \theta_2)}{\pi(w_1, \theta_2)}.
\]

Given \( s_2 = 0 \), it follows that \( e^{-rs_2} = 1 \). Substituting for \( e^{-rs_1} \) from (2.14) and 1 for \( e^{-rs_2} \), the objective function for Problem 1 becomes

\[
F(w_1, w_2) = (1 - q)u(w_1, \theta_1) \frac{\pi(w_2, \theta_2)}{\pi(w_1, \theta_2)} + qu(w_2, \theta_2).
\]

Given \( \theta_1, \theta_2 \) and \( q \), the objective function is a function only of the wage offers \( w_1 \) and \( w_2 \). Hence, optimizing (2.15) with respect to \( w_1 \) and \( w_2 \) is the same as optimizing the objective function for Problem 1. However, simply optimizing (2.15) is not enough to guarantee that the self-selection constraints are satisfied. Matthews and Moore (1987) show that the monotonicity property together with a binding self-selection constraint imply that all the self-selection constraints are satisfied. Therefore, optimizing (2.15) subject to the constraint that \( w_2 \geq w_1 \) yields the solution to Problem 1.

\(^7\) This suggests that an empirical test of the single-crossing property could be derived given a functional form for the firm profit function and suitable data.
Problem 2. Choose \( w_1 \) and \( w_2 \) to maximize \((2.15)\) subject to \( w_2 \geq w_1 \).

Let the solution to Problem 2 be \( \bar{w}_1 \) and \( \bar{w}_2 \). Thus Problem 2 yields the solution to Problem 1 for the two wage offers and the two strike offers are given by \( \bar{s}_2 = 0 \) and, from \((2.14)\),

\[
\bar{s}_1 = -\frac{1}{r} \ln \left[ \frac{\pi(w_2, \theta_2)}{\pi(w_1, \theta_2)} \right].
\] (2.16)

The technique whereby the original optimization problem, Problem 1, is simplified by using the monotonicity property and the pattern of binding self-selection constraints is found in Weymark (1986) and Guesnerie and Laffont (1984). Problem 2 is called the union’s surrogate or reduced-form problem because the strike offers have been eliminated from the union’s objective function.

It is clear from the statement of Problem 2 that there are two possible solutions to the union’s problem, depending on whether the constraint \( w_2 \geq w_1 \) is binding. If the constraint is binding at the solution, i.e. \( \bar{w}_2 = \bar{w}_1 \), then from Proposition 1, \( \bar{s}_1 = \bar{s}_2 = 0 \). That is, the union’s optimal schedule of offers collapses to a single wage offer with no strike. A solution of this nature is known as a pooling or bunching equilibrium in the adverse-selection literature. Alternatively, the constraint in Problem 2 may not bind at the solution, in which case \( \bar{w}_2 > \bar{w}_1 \) and \( \bar{s}_1 > 0 \). In this case, the union’s optimal schedule of offers consists of two wage–strike combinations; a high wage offer without a strike, \((\bar{w}_2,0)\), and a low wage offer with a strike, \((\bar{w}_1,\bar{s}_1)\). A solution of this kind is technically called a separating equilibrium. The possible solutions to Problem 2 are summarized in Proposition 2.

Proposition 2. The optimal schedule of offers takes one of two forms:

(i) \((\bar{w}, \bar{s}) = \{ (\bar{w}_1, \bar{s}_1), (\bar{w}_2, 0) \} \) where \( \bar{w}_2 > \bar{w}_1 \) and \( \bar{s}_1 > 0 \) (a separating equilibrium), or;

(ii) \((\bar{w}, \bar{s}) = \{ (\bar{w}_P, 0) \} \) (a pooling equilibrium).

The next Proposition states the comparative statics properties of the solution to Problem 2 when it is assumed that the outcome is a separating equilibrium.

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9 Weymark (1986)
Proposition 3. Let \( \bar{w}_1(\theta_1, \theta_2, q) \) and \( \bar{w}_2(\theta_1, \theta_2, q) \) be an interior (separating equilibrium) solution to Problem 2. Then,

\[
\frac{\partial \bar{w}_1(\theta_1, \theta_2, q)}{\partial \theta_1} > 0, \quad \frac{\partial \bar{w}_2(\theta_1, \theta_2, q)}{\partial \theta_1} < 0, \\
\frac{\partial \bar{w}_1(\theta_1, \theta_2, q)}{\partial \theta_2} < 0, \quad \frac{\partial \bar{w}_2(\theta_1, \theta_2, q)}{\partial \theta_2} > 0, \\
\frac{\partial \bar{w}_1(\theta_1, \theta_2, q)}{\partial q} = 0, \quad \frac{\partial \bar{w}_2(\theta_1, \theta_2, q)}{\partial q} > 0.
\]

PROOF. Given an interior solution to Problem 2, the first-order conditions are

\[
F_1 = (1 - q) \frac{\pi(w_2, \theta_2)}{\pi(w_1, \theta_2)} \left[ u_w(w_1, \theta_1) - \frac{u(w_1, \theta_1) \pi_w(w_1, \theta_2)}{\pi(w_1, \theta_2)} \right] = 0,
\]

\[
F_2 = (1 - q) u(w_1, \theta_1) \frac{\pi_w(w_2, \theta_2)}{\pi(w_1, \theta_2)} + q u_w(w_2, \theta_2) = 0.
\]

From the first-order condition \( F_1 \) it follows that \( u_w(w_1, \theta_1) < 0 \), and from the first-order condition \( F_2 \) it follows that \( u_w(w_2, \theta_2) > 0 \), since \( \pi_w < 0 \). The second-order cross derivative evaluated at the optimum is

\[
F_{12} = F_{21} = F_1 \frac{\pi_w(w_2, \theta_2)}{\pi(w_2, \theta_2)} = 0. \quad (2.17)
\]

Let \( H \) be the Hessian of the objective function \( F \) evaluated at the optimum \( (\bar{w}_1, \bar{w}_2) \). Given an interior solution to Problem 2, it follows that \( H \) is negative definite. Hence, \( F_{11} < 0 \) and \(|H| = F_{11} F_{22} - F_{12}^2 > 0\), and it must be that \( F_{22} < 0 \).

Totally differentiating the first-order conditions with respect to the exogenous variables, putting the equations into matrix form and substituting for \( F_{12} \), it follows that

\[
\begin{bmatrix} F_{11} & 0 \\ 0 & F_{22} \end{bmatrix} \begin{bmatrix} dw_1 \\ dw_2 \end{bmatrix} = \begin{bmatrix} -F_{1\theta_1} d\theta_1 - F_{1\theta_2} d\theta_2 - F_{1q} dq \\ -F_{2\theta_1} d\theta_1 - F_{2\theta_2} d\theta_2 - F_{2q} dq \end{bmatrix} \quad (2.18)
\]

Now,

\[
F_{1\theta_1} = (1 - q) \frac{\pi(w_2, \theta_2)}{\pi(w_1, \theta_2)} \left[ u_{w\theta}(w_1, \theta_1) - \frac{u_{\theta}(w_1, \theta_1) \pi_w(w_1, \theta_2)}{\pi(w_1, \theta_2)} \right] > 0,
\]

since \( u_{w\theta} > 0 \) by assumption, \( u_{\theta} > 0 \) and \( \pi_w < 0 \).

\[
F_{2\theta_1} = (1 - q) u_{\theta}(w_1, \theta_1) \frac{\pi_w(w_2, \theta_2)}{\pi(w_1, \theta_2)} < 0,
\]

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since \( u_\theta > 0 \) and \( \pi_w < 0 \).

\[
F_{1\theta_2} = -(1 - q) \frac{\pi(w_2, \theta_2)u(w_1, \theta_1)}{\pi(w_1, \theta_2)} \frac{\partial}{\partial \theta_2} \left[ \frac{\pi_w(w_1, \theta_2)}{\pi(w_1, \theta_2)} \right] < 0,
\]
since the term in brackets is positive by the single-crossing property.

\[
F_{2\theta_2} = (1 - q) u(w_1, \theta_1) \frac{\partial}{\partial \theta_2} \left[ \frac{\pi_w(w_2, \theta_2)}{\pi(w_1, \theta_2)} \right] + q u_w(w_2, \theta_2) > 0,
\]
since the term in brackets is positive given the assumption that the elasticity of labour demand with respect to a change in \( \theta \) is less than or equal to 1, and \( u_{w_\theta} > 0 \) by assumption. Finally,

\[
F_{1q} = \frac{-F_1}{(1 - q)} = 0,
\]

\[
F_{2q} = -u(w_1, \theta_1) \frac{\pi_w(w_2, \theta_2)}{\pi(w_1, \theta_2)} + u_w(w_2, \theta_2) > 0,
\]
since \( u_w(w_2, \theta_2) > 0 \). Using Cramer's rule it is straightforward to show that, since \( F_{12} = F_{21} = 0 \),

\[
\frac{\partial w_1}{\partial \theta_1} = -\frac{F_{22}F_{1\theta_1}}{|H|} > 0, \quad \frac{\partial w_2}{\partial \theta_1} = -\frac{F_{11}F_{2\theta_1}}{|H|} < 0,
\]

\[
\frac{\partial w_1}{\partial \theta_2} = -\frac{F_{22}F_{1\theta_2}}{|H|} < 0, \quad \frac{\partial w_2}{\partial \theta_2} = -\frac{F_{11}F_{2\theta_2}}{|H|} > 0,
\]

\[
\frac{\partial w_1}{\partial q} = -\frac{F_{22}F_{1q}}{|H|} = 0, \quad \frac{\partial w_2}{\partial q} = -\frac{F_{11}F_{2q}}{|H|} > 0.
\]

Before summarizing the results to Proposition 3, it is worthwhile discussing the role the various assumptions play in the derivation of the comparative statics results. In order to determine the signs of the \( F_{1\theta_1} \) and \( F_{2\theta_2} \) derivatives, the property that \( u_{w_\theta} > 0 \) is required. Recall that \( u_{w_\theta} > 0 \) if it is assumed that labour demand is more steeply sloped for the high demand firm type. The single-crossing property is invoked in order to determine the sign of the \( F_{2\theta_1} \) derivative, while a stronger assumption—that the elasticity of labour demand with respect to a change in \( \theta \) is less than or equal to 1—is required to determine the sign of \( F_{2\theta_2} \). It is interesting to note that the cross-type derivatives, i.e., \( F_{1\theta_2} \) and \( F_{2\theta_1} \), can be determined under less restrictive assumptions.
Consider the results in Proposition 3. First, the low wage offer depends positively on the level of demand at the low firm type and negatively on the level of demand at the high firm type. Conversely, the high wage offer depends negatively on the level of demand at the low firm type, and positively on the level of demand at the high firm type. Thus, the wage offers depend positively on own-firm type and negatively on cross-firm type. This result has not appeared in the literature, to the best of the author's knowledge. As can be seen from Figure 2.4 the result comes about because of the self-selection constraints. Consider an increase in the level of demand at the low demand firm type, i.e., an increase in the value of $\theta_1$, and an unchanged level of demand at the high firm type. For a small increase in $\theta_1$, the isoprofit curves for the low demand firm become slightly steeper, as illustrated in Figure 2.4, and the low wage offer increases from $w_1$ to $w'_1$. To maintain the binding self-selection constraint on the high demand firm type, the high wage offer has to decrease from $w_2$ to $w'_2$.\(^\text{10}\) Notice that in changing only one of the exogenous variables in this example, namely $\theta_1$, all the offers—the high wage offer, the low wage offer and the strike offer—are affected.

Second, the low wage offer is independent of the union's subjective probability the firm is the high demand type, whereas the high wage offer depends positively on the probability the firm type $\theta_2$. This result is found in Hayes (1984, Proposition 5, p.69). As the union is more certain the firm is the high demand type, the wage offer to the high firm type increases while the wage offered to the low demand type remains unchanged.

The results from Proposition 3 may be used to derive the comparative static properties of the strike offer to the low demand firm, $\tilde{s}_1$.

**Corollary 1.** Let $\tilde{s}_1(\theta_1, \theta_2, q)$ be the strike offer to the low demand firm given a separating equilibrium solution to Problem 2. Then,

$$\frac{\partial \tilde{s}_1(\theta_1, \theta_2, q)}{\partial \theta_1} < 0,$$

$$\frac{\partial \tilde{s}_1(\theta_1, \theta_2, q)}{\partial \theta_2} > 0,$$

$$\frac{\partial \tilde{s}_1(\theta_1, \theta_2, q)}{\partial q} > 0.$$

\(^\text{10}\) It also appears from the diagram that the strike offer to the low demand firm decreases. This result is confirmed below in the Corollary to Proposition 3.
Figure 2.4
Illustration of the Comparative Statics for an Increase in $\theta_1$
PROOF. The Corollary follows directly from Proposition 3 and equation (2.14). For example, from Proposition 3, increasing \( \theta_1 \) increases \( \bar{w}_1 \) and decreases \( \bar{w}_2 \), so from (2.14) \( \bar{s}_1 \) must increase.

Thus the strike offer increases the more certain the union is that the firm is the high demand type. Therefore, as the union becomes more certain the firm is the high demand type, the wage offer to the high demand type increases and the strike offer to the low demand type also increases. The first two results in the Corollary may be interpreted as indicating that the strike offer increases the greater the difference between the high and the low demand firm types. For, as \( \theta_1 \) decreases and \( \theta_2 \) increases—i.e., the difference between firm types widens—both changes effect \( \bar{s}_1 \) in the same way, namely the strike offer increases.

Since \( \bar{w}_1 \) is independent of \( q \) and since \( \bar{w}_2 \) depends positively on \( q \), there exists a unique value of \( q \), called \( \bar{q} \), which depends on \( \theta_1 \) and \( \theta_2 \), such that if \( q > \bar{q} \) then \( \bar{w}_2 > \bar{w}_1 \), and if \( q \leq \bar{q} \) then \( \bar{w}_1 = \bar{w}_2 = \bar{w}_P \). It is natural to call \( \bar{q} \) the union's strike threshold probability; if \( q > \bar{q} \), the solution to Problem 2 is a separating equilibrium and therefore one of the union’s offers includes a strike. If \( q \leq \bar{q} \) the solution to Problem 2 is a pooling equilibrium in which case there is no strike offered. Suppose \( q > \bar{q} \), so \( \bar{w}_2 > \bar{w}_1 \), and then decrease \( q \) until \( \bar{w}_2 = \bar{w}_1 \), at which point \( q = \bar{q} \) by definition. Since \( \bar{w}_1 \) has not changed, \( \bar{q} \) is determined as the value of \( q \) for which \( \bar{w}_1 \) is the solution to Problem 2. Setting \( w_2 = \bar{w}_1 \), the objective function (2.15) for Problem 2 is written:

\[
F(\bar{w}_1) = (1 - q)u(\bar{w}_1, \theta_1) + q u(\bar{w}_1, \theta_2).
\] (2.19)

Given the objective function (2.19), the first-order condition to Problem 2 is

\[
F_P = (1 - q) u_w(\bar{w}_1, \theta_1) + q u_w(\bar{w}_1, \theta_2) = 0.
\] (2.20)

Hence, \( \bar{q} \) is determined by solving (2.20) for \( q \):

\[
\bar{q}(\theta_1, \theta_2) = \frac{-u_w(\bar{w}_1, \theta_1)}{u_w(\bar{w}_1, \theta_2) - u_w(\bar{w}_1, \theta_1)}.
\] (2.21)

It is convenient to assume that \( \bar{q}(\theta_1, \theta_2) \) is differentiable. From the proof to Proposition 3, it is known that \( u_w(\bar{w}_1, \theta_1) < 0 \) and \( u_w(\bar{w}_1, \theta_2) > 0 \).
Proposition 4. The union’s strike threshold probability is an decreasing function of $\theta_2$ and a increasing function of $\theta_1$.

PROOF. Differentiating (2.21) it follows that
\[
\frac{\partial \tilde{q}(\theta_1, \theta_2)}{\partial \theta_1} = -\frac{u_w(\bar{w}_1, \theta_1)}{u_w(\bar{w}_1, \theta_2) - u_w(\bar{w}_1, \theta_1)} + \frac{u_w(\bar{w}_1, \theta_1)u_{w\theta}(\bar{w}_1, \theta_2)}{[u_w(\bar{w}_1, \theta_2) - u_w(\bar{w}_1, \theta_1)]^2} < 0,
\]
since $u_w(\bar{w}_1, \theta_1) < 0$, $u_w(\bar{w}_1, \theta_2) > 0$ and $u_{w\theta} > 0$. In addition
\[
\frac{\partial \tilde{q}(\theta_1, \theta_2)}{\partial \theta_2} = \frac{u_w(\bar{w}_1, \theta_1)u_{w\theta}(\bar{w}_1, \theta_2)}{[u_w(\bar{w}_1, \theta_2) - u_w(\bar{w}_1, \theta_1)]^2} > 0,
\]
since $u_w(\bar{w}_1, \theta_1) < 0$ and $u_{w\theta} > 0$.

Thus increasing the value of $\theta_2$ reduces the union’s strike threshold probability. A corollary to this result is that, for given values of $q$, a separating solution to Problem 2 is more likely the higher the value of $\theta_2$. Therefore, the greater the difference between the high and the low demand firm types, the more likely that a strike will be the outcome of bargaining, ceteris paribus.

The intuition for this important result is straightforward. From the union’s point of view the high demand firm type is always more attractive than the low demand type, since the high demand type can afford to pay higher wages and hire more workers. Increasing $\theta_2$ relative to $\theta_1$—making the high demand type a “higher” demand type—only makes the type $\theta_2$ firm more attractive. In other words, as $\theta_2$ increases, there is more incentive for the union to induce the high demand firm to reveal its type. To induce the high demand firm to reveal its type, the union must offer a separating equilibrium which, of course, includes the offer of a strike to the low demand type. The result contained in Proposition 4 has not appeared elsewhere in the adverse-selection literature, to the best of my knowledge.

The next proposition states the comparative statics properties for Problem 2 given that a pooling solution obtains.

Proposition 5. Let $\bar{w}_P(\theta_1, \theta_2, q)$ be a boundary (pooling equilibrium) solution to Problem 2. Then,
\[
\frac{\partial \bar{w}_P(\theta_1, \theta_2, q)}{\partial \theta_1} > 0,
\]
\[
\frac{\partial \bar{w}_P(\theta_1, \theta_2, q)}{\partial \theta_2} > 0,
\]
\[
\frac{\partial \bar{w}_P(\theta_1, \theta_2, q)}{\partial q} > 0.
\]
PROOF. Given \(wp = w_1 = w_2\), the objective function (2.15) for Problem 2 is given in (2.19) and the first-order condition is given in (2.20) where \(wp\) is substituted for \(\bar{w}_1\). The second-order derivative is

\[
F_{pp} = (1 - q)u_{ww}(wp, \theta_1) + q u_{ww}(wp, \theta_2) < 0,
\]

at a solution. Totally differentiating the first-order condition with respect to the exogenous variables yields

\[
F_{pp} dw_p + (1 - q)u_{w\theta}(wp, \theta_1)d\theta_1 +
q u_{w\theta}(wp, \theta_2)d\theta_2 + [u_w(wp, \theta_2) - u_w(wp, \theta_1)]dq = 0.
\]

Therefore

\[
\frac{\partial \bar{w}_p(\theta_1, \theta_2, q)}{\partial \theta_1} = -\frac{(1 - q)u_{w\theta}(wp, \theta_1)}{F_{pp}} > 0,
\]

\[
\frac{\partial \bar{w}_p(\theta_1, \theta_2, q)}{\partial \theta_2} = -\frac{q u_{w\theta}(wp, \theta_2)}{F_{pp}} > 0,
\]

\[
\frac{\partial \bar{w}_p(\theta_1, \theta_2, q)}{\partial q} = \frac{u_w(wp, \theta_1) - u_w(wp, \theta_2)}{F_{pp}} > 0,
\]

where the first two results follow from the assumption that \(u_{w\theta} > 0\), and the last result follows from the fact that \(u_w(wp, \theta_1) < 0\) and \(u_w(wp, \theta_2) > 0\) as discussed preceding Proposition 5.

2.2.1 An Example

The model presented above is quite difficult to solve for analytic solutions, even when the functional forms chosen for the utility and profit functions are relatively simple. This is illustrated with the following example. Let the firm's profit function be defined thus

\[
\pi(w, \theta_i) = \frac{\bar{L}(w - \theta_i)^2}{2\theta_i}, \quad i = 1, 2,
\]

where \(\bar{L} > 0\). Applying Hotelling's Lemma, labour demand is given by

\[
L(w, \theta_i) = -\pi_w(w, \theta_i) = \frac{\bar{L}(\theta_i - w)}{\theta_i}, \quad i = 1, 2.
\]

Hence, the profit function is such that labour demand is linear for each firm type. Since \(\theta_i\) is the vertical intercept for the labour demand curve for each firm type in \(w\)
and $L$ space, the union will never make a wage offer such that $w_i \geq \theta_i, i = 1, 2$. It is easy to show that the profit function also satisfies the single-crossing property. Let union utility be simply the total wage bill, thus

$$u(w, \theta_i) = w L(w, \theta_i) = \frac{w L_i(\theta_i - w)}{\theta_i}, \quad i = 1, 2.$$ 

The union's surrogate objective function is found by substituting the above expressions for utility and profits into (2.15). Maximizing the resulting function with respect to $w_1$ and $w_2$, and assuming a separating solution, it follows from the first-order conditions that

$$\frac{\partial u}{\partial w_1} = \frac{\theta_1 \theta_2}{2\theta_2 - \theta_1},$$

$$\frac{\partial u}{\partial w_2} = \frac{q[\theta_2(2\theta_2 - \theta_1)]}{q(4\theta_2 - 3\theta_1) - \theta_1}.$$

Substituting for $w_1$ and $w_2$ in (2.16) yields

$$s_1 = -\frac{1}{r} \ln \left( \frac{q^2(2\theta_2 - \theta_1)^2}{[q(4\theta_2 - 3\theta_1) - \theta_1]^2} \right).$$

Assuming a pooling solution it can be shown that the optimal value of the pooling wage is given

$$\bar{w}_p = \frac{\theta_1 \theta_2}{2[(1 - q)\theta_2 + q\theta_1]}.$$

If a more complicated utility function is chosen, for example the Stone-Geary form given in (2.6), it is no longer possible to derive analytic solutions for the optimal wage offers in the case of a separating solution.

### 2.3 Empirical Implementation of the Model

Empirical studies of strikes have tended to focus primarily on the estimation of strike duration or strike probability equations. These approaches are certainly valid for an empirical study based on the present model. An econometric model of strike probabilities is estimated in Chapter 3. However, in the adverse-selection model proposed above, the wage and strike offers are jointly determined given values for the exogenous variables. Ideally, then, an empirical evaluation of the adverse-selection model would include joint estimation of wage and strike equations. Furthermore, the monotonicity property imposes restrictions on the wage and strike offers that may be
used to identify the econometric model. Chapters 4 and 5 are concerned with building an econometric model of wage and strike offers using the monotonicity property, and then using the model to test for the comparative statics predictions listed above.

As stated above,\textsuperscript{11} the monotonicity property is common to many models of asymmetric information based on the adverse-selection approach. For this reason, it is argued that the econometric model proposed in chapter 4 is applicable not only to an empirical model of wage and strike offers, but also to any empirical application of an adverse-selection model.

For the purposes of empirical implementation, one of the most important aspects of limiting the theoretical model to the case of two potential firm types is that the occurrence of a strike identifies the presence of a low demand firm type and, hence, a low wage offer. Thus, the occurrence of a strike enables the observed wage to be identified as a low wage offer. In addition, the absence of a strike indicates either the presence of a high demand firm and, therefore, a high wage offer, or the presence of a pooling wage offer. These facts will prove extremely useful in formulating the econometric model proposed in chapter 4. In the case where the number of firm types is greater than two, the analogue to the two-type monotonicity property in Proposition 1—for the case of a separating solution—is that a zero strike is offered only to the highest demand firm type. Consequently, the occurrence of a strike indicates only the absence of the highest firm type, rather than the presence of a specific firm type. Therefore, when a strike occurs, the observed wage cannot be linked to a particular wage offer.

For empirical implementation it is useful to enrich the set of comparative static predictions in order to yield a greater number of testable hypotheses, and to enhance the realism of the model. Introducing extra variables into the profit and utility functions must be done carefully so the structure of the model remains intact. Suppose the union utility function (2.5) is written

\[ u(w, x, \theta_i) = \tilde{u}(w, x, L(w, \theta_i)). \]  

(2.22)

The variable \( x \) may be thought of as something which affects union utility but not firm profits and which is determined outside the model. For example, \( x \) could be chosen to be \( \tilde{w} \), the alternative wage available to union workers, in the Stone-Geary and MacDonald and Solow (1981) functional forms given in (2.6) and (2.7). The

\textsuperscript{11} See the discussion after Proposition 1.
utility function (2.22) may be substituted into the union’s surrogate objective function (2.15), and the resulting first-order conditions are

\[ F_1 = (1 - q)\frac{\pi(w_2, \theta_2)}{\pi(w_1, \theta_1)} \left[ u_w(w, x, \theta_1) - \frac{u(w, x, \theta_1)\pi_w(w_1, \theta_2)}{\pi(w_1, \theta_2)} \right] = 0, \]

\[ F_2 = (1 - q)u(w_1, x, \theta_1)\frac{\pi_w(w_2, \theta_2)}{\pi(w_1, \theta_2)} + q u_w(w_2, x, \theta_2) = 0, \]

which are the same as before except for the presence of the variable \( x \) in the utility function. Clearly, the second-order derivatives with respect to \( w_1 \) and \( w_2 \) do not change with the new specification for \( u \). Therefore, in order to derive the comparative statics properties with respect to changes in \( x \), only the following derivatives need to be calculated

\[ F_{1x} = (1 - q)\frac{\pi(w_2, \theta_2)}{\pi(w_1, \theta_1)} \left[ \frac{u_{wx}(w_1, x, \theta_1)}{\pi(w_1, \theta_2)} - \frac{u_x(w_1, x, \theta_1)\pi_w(w_1, \theta_2)}{\pi(w_1, \theta_2)^2} \right], \]

\[ F_{2x} = (1 - q)u_x(w_1, x, \theta_1)\frac{\pi_w(w_2, \theta_2)}{\pi(w_1, \theta_2)} + q u_{wx}(w_2, x, \theta_2). \]

The signs of the derivatives \( F_{1x} \) and \( F_{2x} \) clearly depend on the signs of the derivatives \( u_x \) and \( u_{wx} \). Sufficient conditions to determine the signs of the derivatives \( F_{1x} \) and \( F_{2x} \) are provided in the following lemma.

**Lemma 4.** For the union utility function defined by (2.22) and the second derivatives \( F_{1x} \) and \( F_{2x} \) given in (2.23):

(i) \( F_{1x} < 0 \) if \( u_x < 0 \) and \( u_{wx} < 0 \), while \( F_{1x} > 0 \) if \( u_x > 0 \) and \( u_{wx} > 0 \), and;

(ii) \( F_{2x} < 0 \) if \( u_x > 0 \) and \( u_{wx} < 0 \), while \( F_{2x} > 0 \) if \( u_x < 0 \) and \( u_{wx} > 0 \).

**PROOF.** By inspection of equations (2.23) given \( \pi_w < 0 \),

Thus if \( u_x \) and \( u_{wx} \) have the same sign, the sign of \( F_{1x} \) may be unambiguously predicted while the sign of \( F_{2x} \) is indeterminate. On the other hand, if \( u_x \) and \( u_{wx} \) have opposite signs, the sign of \( F_{2x} \) may be unambiguously predicted while the sign of \( F_{1x} \) is indeterminate. Following the proof of Proposition 3, Cramer’s rule may be used to solve for the comparative statics properties of \( \bar{w}_1 \) and \( \bar{w}_2 \) with respect to a change in \( x \). Let \( \bar{w}_1(\theta_1, \theta_2, q, x) \) be the optimal low wage offer and \( \bar{w}_2(\theta_1, \theta_2, q, x) \) be the optimal high wage offer in the case of a separating solution for the union utility function defined in (2.22), then

\[ \frac{\partial \bar{w}_1(\theta_1, \theta_2, q, x)}{\partial x} = -\frac{F_{22}F_{1x}}{|H|}, \quad \frac{\partial \bar{w}_2(\theta_1, \theta_2, q, x)}{\partial x} = -\frac{F_{11}F_{2x}}{|H|}, \]
where $H$ is the Hessian of the surrogate objective function and $F_{11}$ and $F_{22}$ are defined in the proof of Proposition 3. The comparative statics of the model with respect to a change in $x$, are defined in the following proposition.

**Proposition 6.** Let $\bar{w}_1(\theta_1, \theta_2, q, x)$ and $\bar{w}_2(\theta_1, \theta_2, q, x)$ be a separating solution to Problem 2 given the utility function defined in (2.22). Then,

$$\frac{\partial \bar{w}_1(\theta_1, \theta_2, q, x)}{\partial x} \leq 0 \quad \text{if} \quad F_{1x} \leq 0,$$

and

$$\frac{\partial \bar{w}_2(\theta_1, \theta_2, q, x)}{\partial x} \geq 0 \quad \text{if} \quad F_{2x} \geq 0.$$

**PROOF.** The result follows immediately from equations (2.24)

Given the assumptions on $u_x$ and $u_{wx}$ it is known from Lemma 4 that only one of the derivatives $F_{1x}$ and $F_{2x}$ can be signed. Thus if the response of one of the wage offers to a change in $x$ can be predicted, it follows that the response of the other wage offer to a change in $x$ is ambiguous. The Stone-Geary and the MacDonald and Solow (1981) utility functions given in (2.6) and (2.7) respectively, both have the property that $u_w < 0$ and $u_{wL} > 0$. Hence, it follows from Lemma 4 that $F_{2w} > 0$ and so from Proposition 6 it must be that $\partial \bar{w}_2(\theta_1, \theta_2, q, \bar{w})/\partial \bar{w} > 0$. If the alternative wage available to union members increases, the wage offer to the high demand firm type increases. It can also be shown for the Stone-Geary utility function that $u_L < 0$ and $u_{wL} < 0$. Thus from Lemma 4 and Proposition 6 it follows $\partial \bar{w}_1(\theta_1, \theta_2, q, \bar{L})/\partial \bar{L} < 0$. If the “subsistence” level of employment increases, the wage offer to the low demand firm type decreases.

As in the proof for Proposition 5, the comparative statics predictions for $\bar{w}_P$ given a change in $x$ may be derived. Let $\bar{w}_P(\theta_1, \theta_2, q, x)$ be the optimal pooling wage offer for a pooling solution to Problem 2 given the utility function defined in (2.22). The first-order condition is given by

$$F_P = (1 - q)u_w(w_P, x, \theta_1) + qu_w(w_P, x, \theta_2), \quad (2.25)$$

so it follows that

$$F_{Px} = (1 - q)u_{wx}(w_P, x, \theta_1) + qu_{wx}(w_P, x, \theta_2).$$

Hence, if $u_{wx} > 0$ then $F_{1x} > 0$, and if $u_{wx} < 0$ then $F_{1x} < 0$.

12 That is, after the appropriate substitutions for $x$ in the utility functions.
Proposition 7. Let $\hat{w}_P(\theta_1, \theta_2, q, x)$ be a pooling solution to Problem 2 given the utility function defined in (2.22). Then,

$$\frac{\partial \hat{w}_P(\theta_1, \theta_2, q, x)}{\partial x} \geq 0 \quad \text{if} \quad F_{P_x} \geq 0.$$  

PROOF. Totally differentiating the first-order condition (2.25) and setting $d\theta_1 = 0$, $d\theta_2 = 0$ and $dq = 0$, it follows that

$$\frac{\partial \hat{w}_P(\theta_1, \theta_2, q, x)}{\partial x} = -\frac{F_{P_x}}{F_{PP}},$$

and the result follows since $F_{PP} < 0$ at a solution.

The Stone-Geary utility function is such that $u_{w\hat{w}} > 0$ and $u_{wL} < 0$. Hence, increasing the alternative wage $\hat{w}$ increases $\hat{w}_P$, and increasing $\hat{L}$ decreases $\hat{w}_P$.

For reference, the complete set of comparative statics properties of the model contained in Propositions 3, 5, 6 and 7 are summarized in Table 2.1. The endogenous variables in the adverse-selection model are arranged in a row across the top of the table, and the exogenous variables are listed in the left-hand column. Each entry in the table depicts the predicted relationship between a specific exogenous variable and a specific endogenous variable. For example, the "+" in the upper left-hand corner of the table indicates that the low wage offer, $w_1$, is positively related to the level of demand at the low demand firm type, $\theta_1$. A zero entry in the table indicates that the two variables are not related, while a "?" indicates that an unambiguous relationship is not predicted by the model.

The last row in Table 2.1 depicts the relationship between the endogenous variables—assuming a separating equilibrium—and the ratio of the high to the low levels of demand, $\frac{\theta_2}{\theta_1}$. Entries in this row are derived using the fact that $\theta_1$ and $\theta_2$ have opposite effects on each of the endogenous variables. For example, since $\theta_1$ and $w_1$ are positively related and $\theta_2$ and $w_1$ negatively related, it follows that an increase in the ratio $\theta_2/\theta_1$ results in a decrease in $w_1$. However, since $\theta_1$ and $\theta_2$ both have a positive effect on the pooling wage, the effect of a change in the ratio on $w_P$ cannot be determined. For example, increasing $\theta_2$ increases the ratio $\theta_2/\theta_1$ and has a positive effect on the pooling wage; but decreasing $\theta_1$, which also increases the ratio $\theta_2/\theta_1$, has a negative effect on the pooling wage. Hence the appearance of "N.A." (not applicable) in the lower right hand corner of the table.
Table 2.1
Summary of Comparative Statics Predictions from the Model

<table>
<thead>
<tr>
<th></th>
<th>$w_1$</th>
<th>$w_2$</th>
<th>$s_1$</th>
<th>$w_P$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\theta_1$</td>
<td>+</td>
<td>-</td>
<td>-</td>
<td>+</td>
</tr>
<tr>
<td>$\theta_2$</td>
<td>-</td>
<td>+</td>
<td>+</td>
<td>+</td>
</tr>
<tr>
<td>$q$</td>
<td>0</td>
<td>+</td>
<td>+</td>
<td>+</td>
</tr>
<tr>
<td>$\bar{w}$</td>
<td>?</td>
<td>+</td>
<td>?</td>
<td>+</td>
</tr>
<tr>
<td>$\bar{L}$</td>
<td>-</td>
<td>?</td>
<td>?</td>
<td>-</td>
</tr>
</tbody>
</table>

$\theta_2/\theta_1$ | - | + | + | N.A.
CHAPTER III

Description of the Data and a Conventional Model of Strike Incidence

3.1 Introduction

The endogenous variables in the asymmetric-information models of strikes are the wage level and the length of the strike. However, empirical tests of the asymmetric information theory have tended to focus on the incidence and duration of strikes rather than on the level of wages. An econometric model in which the wage level and the strike length are endogenous variables is the subject of chapters 4 and 5. For the present chapter, however, a natural departure point for an empirical analysis of the adverse-selection model is the existing empirical literature. This chapter is concerned with the specification and estimation of strike probability equations, and the construction of the data set used in the estimation.

There are several reasons for following the conventional empirical approach to modelling strikes. First, the existing empirical knowledge provides useful standards of comparison for the present model. For example, it will be instructive to see how variables suggested by the adverse-selection model perform in a conventional econometric model of strike incidence. Second, the contract data used in this study are a more recent and expanded version of the contract data used by Swidinsky and Vanderkamp (1982) and Gunderson, Kervin and Reid (1986). As these authors estimate strike incidence equations, their work provides a useful check of the firm and contract data set assembled and analyzed here. Finally, the analysis in the present chapter represents the first time firm-level data have been used to evaluate adverse-selection models of collective bargaining.

An essential feature of the theoretical model of strikes proposed in the previous chapter is the notion that strikes are a microeconomic phenomenon: the strike serves to resolve an asymmetry of information between the union and the firm. This being the case, an empirical evaluation of the theory is compelled to use data which are specific to individual firm-union bargaining pairs. The most useful source of information about individual bargaining pairs in Canada is the Major Collective Agreements (MCA) computer tape assembled by Labour Canada.

1 Two exceptions are Card (1987) and McConnell (1987b).
2 The computer tape is encoded with contract data from the monthly publication "Collective Bargaining Review," Ottawa: Labour Canada.
a rich source of data pertaining to collective agreements between firms and unions in all sectors of the economy and from all regions of the country. However, apart from the name of the firm co-signing the collective agreement, the tape contains no information about the firm. According to the theoretical model, firm-specific conditions, e.g., the level of product demand, play a crucial role in the outcome of bargaining. Therefore, it is necessary to add information about the firm to the contract data from the MCA tape in order to derive a data set to test the model. One of the major contributions of this study is the construction of a data set which merges contract data with relevant firm-specific information.

Briefly, the chapter is organized as follows. In the next section, the estimating model is formally presented and the choice of explanatory variables is discussed. Section 3.3 details the assembly of the firm and contract data set and discusses the calculation of the independent variables. In section 3.4, the regression results from OLS and logit estimation of the strike probability equations are presented and evaluated. The final section of the chapter contains some concluding remarks.

3.2 A Conventional Model of Strike Incidence

The key exogenous variables in the adverse-selection model are: \( \theta \), the level of firm output demand, and; \( q \), the union's subjective probability the firm is the high demand type. Proposition 1 states that strikes occur only in low demand firm types. Thus, according to the theory, there should be a negative correlation between the incidence of strikes and firm-level output demand. Proposition 4 states that strikes are more likely to occur the greater the difference between the levels of demand for the two potential firm types. Given individual firm and contract data, the conventional way to test these predictions is to use an econometric specification of the form:

\[
DS_i = g(\theta_i, (\theta_2/\theta_1)_i, q_i, X_i) + e_i \quad i = 1, ..., N
\]

(3.1)

where \( DS_i \) is a dummy variable with a value of 1 if a strike occurs during negotiations for the \( i \)th contract and 0 otherwise, \( g \) is the functional form of the estimating equation (e.g., linear or logitistic), \( N \) is the number of contracts in the sample, and \( e_i \) is a random disturbance term with a zero mean. The vector \( X \) contains variables known to affect strike incidence but ignored by the adverse-selection model. The reasons for including \( X \) are explained below.
To estimate the model, $g$ must be specified and suitable proxies found for the independent variables. In the adverse-selection model, $\theta_i$ represents the level of demand at the firm that will prevail over the life of the contract. Obviously, when the contract is being negotiated, $\theta_i$ is not observed by the union or, for that matter, by the researching economist. Thus the proxy chosen for the level of demand at the firm must be something that is not observed until after the contract is signed. A suitable proxy for $\theta_i$ which satisfies this criterion is the revenue of the firm for the fiscal year which ends some time after the starting date of the contract.

Morton (1983) and McConnell (1987a) point out the proxy for the firm's private information must not be something that is somehow observed by a researching economist but not by the union. "Private information" means that only the firm knows the information. When the information is revealed—becomes public knowledge—everyone knows it. Clearly, what is important is the moment at which the private information becomes public, or in other words, when the variable that is to proxy the private information is observable. As long as the appropriate variable is observed after the union and the firm have agreed to a contract, we need not be concerned about pretending economists are better informed than unions. Choosing to proxy $\theta$ with a revenue figure that is unknown at the time of bargaining but becomes known at a future date, is entirely consistent with this approach.

Assuming that the firm's type is observable some time after the contract is agreed upon requires that the possibility of contingent contracts is ruled out. If the union were able to make offers contingent on a future realization of the private information variable, the asymmetry of information would be irrelevant. The union could achieve incentive compatibility by, for example, imposing a penalty if it turned out the firm had incorrectly revealed its type. Ruling out contingent contracts, however, maintains the sanctity of the asymmetry of information between the union and the firm. In fact, contracts which are contingent on some characteristic of the firm are not observed in collective bargaining data. Perhaps periodic renegotiation, which we do observe, is a substitute for such contracts.

An empirical proxy for $(\theta_2/\theta_1)$, the ratio of the high to the low state of firm product demand, is difficult to implement because only one of the two levels is observed for each contract. Suppose, however, that $\theta_1$ and $\theta_2$ are determined by the historical pattern of the firm's product demand. Thus the more unstable the firm's product demand, the larger the value of $(\theta_2/\theta_1)$. A suitable proxy for the stability of
firm product demand is the coefficient of variation (defined as the standard deviation divided by the mean) of firm revenue over an appropriately chosen period.

An empirical proxy for \( q \), the union's subjective probability the firm is the high demand type, is difficult to implement. Indeed, the variable is, by definition, subjective. However, \( q \) is certain to be correlated with the business cycle (i.e., the aggregate level of demand in the economy), the time of year (e.g., in seasonal industries, or at Christmas), and possibly the industry in which the bargaining pair are located. To the extent that these factors are important determinants of \( q \), the effect of the union's subjective probability on strike incidence may be captured by measures of the aggregate business cycle, seasonal and industry dummy variables. In addition, to the extent that differences in \( q \) across unions remain constant over time, the cross-sectional effect of the variable may be captured by a union dummy variable.

The variables in \( X \) serve two purposes. First, the variables suggested by the theoretical model are not, by themselves, expected to provide a fully satisfactory empirical model of strike incidence. The nature of collective bargaining is such that a variety of institutional, seasonal, and demographic factors may play an important role. One reason for including \( X \), therefore, is to control for these factors. Second, elements of \( X \) are chosen on the basis of proven statistical performance in previous empirical studies of strikes. In this way, (3.1) may be thought of as a representative econometric model of strike incidence with extra variables suggested by the adverse-selection model.

3.3 Description of the Data

The data used in the empirical analysis are derived from two principal sources. Information on collective agreements between unions and firms is obtained from a computer tape constructed by Labour Canada called the Major Collective Agreements (MCA) tape. Individual firm revenue data are collected from the Financial Post Corporation Card Service. Other data used in the study are culled from various sources such as the CANSIM University Data Base, the CANSIM Main Base and Statistics Canada publications.

The MCA tape was made available by the Labour Data Branch of Labour Canada and is dated August 29, 1986. The MCA tape contains 211,960 records—one record per month for each month of the duration of 9,423 collective agreements or contracts covering unionized bargaining units with 500 or more workers located in
Canada for the period 1964 to 1986. The first step is to reduce the massive amount of data to a more manageable unit by choosing just one of the monthly records per contract. Thus the last monthly record from each series of monthly contract observations is chosen to represent each contract. In this way, each record in the data set corresponds to one collective agreement. The following information is available for each contract: a file number specific to each firm-union bargaining pair which is held over from one agreement to the next; the settlement date, effective date, and termination date of the agreement; the duration of the agreement; geographic location; the name of the union; the jurisdiction of the relevant labour legislation; the settlement stage, including whether a strike occurred during negotiations; the duration of negotiations; the number of employees; the three-digit S.I.C. number of the employer; a commercial-non-commercial sector indicator; a public-private sector indicator; the wage of the base rate occupation; and, the employer name.

Since the theoretical model applies to contracts between unions and profit-maximizing firms, public sector contracts are deleted from the sample using the public-private sector indicator. This reduces the size of the sample by 4,723 observations, i.e. from 9,423 to 4,700 contracts. For the same reason, non-commercial sector contracts are removed from the sample, reducing the number of contracts by another 4 to 4,696. The settlement stage of the contract is coded into one of 15 categories. If the settlement stage is "unknown," "legislated," or "other," the observation is deleted. This reduces the sample by a further 103 contracts, resulting in a sample size of 4,593 private sector contracts.

Having constructed a sample of private sector contracts, the next step is to match each contract with the relevant firm revenue data. This is done using "employer name" from the MCA tape. The source of firm-specific data is the Financial Post Corporation Card Service which comprises one card for each of 600 or so major corporations in Canada. The Card Service is updated periodically as new information on each corporation becomes available, which is usually after the company releases its annual report. Each card contains a list of wholly- and partially-owned subsidiaries, the company's balance sheet at the most recent fiscal year end, and a complete historical record of annual sales or revenue, dividends, annual share price highs

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3 Thus the average duration of agreements on the MCA tape is $211,960/9,423 = 22.5$ months.

4 The name of this particular statistic differs between firms due to variations in accounting procedures. However, Statistics Canada claims these differences are negligible for industrial
and lows, net income, and assets. For the purposes of this study, the main variable of interest is the annual revenue of the firm. The nature of the contract data is such that most firm-union bargaining pairs are present for more than one contract. A revenue series corresponding to each sequence of collective agreements for the same firm and bargaining unit is constructed for the period starting with the effective date of the first contract and ending with the termination date of the last contract in the sequence. Some contracts last longer than others, and some bargaining pairs are present in the data for longer periods than others. As a result, the revenue series' vary in length; the shortest revenue series has 2 observations and the longest revenue series has 23 observations, with the median being 11.

In this manner, 2,459 contracts representing 497 bargaining pairs are successfully matched with firm revenue data. Some of the corporations report revenue in U.S. dollars. To make sure all the revenue figures are in the same units, the U.S. dollar figures are converted to Canadian dollars using the annual average of the noon spot exchange rate. The source of the exchange rate data is the Bank of Canada Review, February, 1975, p.63 and March, 1986, p.II. The revenue figures are converted to constant dollars (with 1971 as the base year) using the Consumer Price Index taken from Statistics Canada, Catalogue 62-001, December, 1979, p.5 and June, 1987, p.11.

To proxy the level of demand at the firm for each contract, a variable REVENUE is defined as the revenue figure for the fiscal year ending at least 6 months and no more than 18 months after the effective date of the contract. Since the revenue data are annual and since contracts last, on average, two years, it was decided to calculate the coefficient of variation of firm revenue, C.V.REVENUE, for the entire revenue series corresponding to each sequence of contracts for the same firm-union pair. Thus the value of C.V.REVENUE is the same for each contract in a particular sequence of contracts for the same bargaining pair. Calculating the coefficient of variation in this way conserves degrees of freedom and gives a more robust estimate of the variability of firm product demand than, for example, calculating a separate coefficient of variation for each contract using the revenue data to that point in time.

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5 In some cases, the firm data are taken from the Financial Post Survey of Industrials which gathers information from the same source—company reports to shareholders—as the Card Service.
Several caveats apply to the revenue data. First, for those companies which are subsidiaries and for which the firm's own revenue figure is not available, the revenue figure of the parent company is used. In addition, some companies which produce more than one product provide disaggregated revenue figures. Where feasible, the revenue figure of the appropriate division is used to ensure that the revenue figure chosen corresponds as closely as possible to the product produced by the bargaining unit on the contract.

Second, due to acquisitions of other firms, new plants coming on stream, and other similar infrequent events, the revenue figures for a firm may change substantially from one year to the next. In such cases, C.V.REVENUE as calculated will overstate the variability in the firm's level of demand. To control for discrepancies of this nature, a dummy variable DRC is assigned a value of 1 for contracts after an infrequent event known to have affected the firm's revenue, and 0 otherwise. Occasionally, firm name changes between successive contracts for the same bargaining unit. This may occur for a number of reasons. For example, the original company may have been taken over, or the plant where the union is located may have been sold to another company, or the company may simply have changed its name. If suitable revenue figures can be found for the "new" firm, they are added to the revenue series of the "old" firm. To remove any discrepancies this may cause, DRC is assigned a value of 1 if the revenue figures come from different companies during a sequence of contracts with a particular union, and 0 otherwise. In special cases where a firm is taken over, a dummy variable DTO is assigned a value of 1 for contracts after the take over, and 0 otherwise. Another source of discrepancy in the revenue data occurs when companies change fiscal year ending dates. To ensure that all the revenue figures cover a 12 month period, revenue figures after fiscal year end changes are converted to annual figures by calculating the monthly average and multiplying the resulting figure by 12. Finally, in a few cases there is a substantial time lapse between the termination date of a contract and the effective date of the subsequent contract from the same bargaining pair. This may be because the union and the firm were working without a contract for a period of time, or simply because Labour Canada omitted a contract from the data for some (unknown) reason. The result of a break in the contract series is that the revenue data will skip a few years. Therefore C.V.REVENUE may be distorted,

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6 These events are identified with the help of a historical synopsis of each company's development which is contained in the Card Service.
especially if the gap in the contract series is for several years or more. To control for such an event, a dummy variable, DCM, is assigned a value of 1 if the break in the revenue series is more than two years, and 0 otherwise.

So far, only variables suggested by the adverse-selection model of strikes have been discussed. Now attention is turned to the components of X. There is some evidence that the propensity to strike is positively related to the size of the bargaining unit. It is also likely that the larger the bargaining unit, the larger the revenue of the firm. Therefore there is some danger that the variable REVENUE will capture the effect of bargaining unit size rather than the effect of an asymmetry of information. To try to avoid the problem, a variable measuring the number of employees covered by the contract (NEMP), is included in the estimation of (3.1). A similar variable is used by Swidinsky and Vanderkamp (1982), Gunderson et al (1986), Cousineau and Lacroix (1986), and Gramm (1986), all of whom report a positive and significant effect of the variable on strike incidence.

Swidinsky and Vanderkamp (1982) report a significant negative relationship between strike incidence and the length of contract negotiations. In the present study, a variable measuring the duration of negotiations in months for the contract (LNEG), is included in estimation. Gunderson et al (1986) and Swidinsky and Vanderkamp include regional and seasonal dummy variables in their strike incidence equations. In the present study, regional dummy variables are constructed for the Maritimes (Newfoundland, Prince Edward Island, Nova Scotia, New Brunswick), Quebec, Ontario, the Prairies (Manitoba, Saskatchewan, Alberta), British Columbia, the North (Yukon, Northwest Territories), and for contracts covering bargaining units in more than one province. Seasonal dummy variables are constructed for the winter (January, February, March), spring (April, May, June), summer (July, August, September), and autumn (October, November, December) according to the settlement date of the contract. The reason settlement date is used rather than effective date or termination date is that the latter two dates are often fixed by institutional factors such as the fiscal year end of the firm. The purpose of these dummy variables is to capture the variability in the propensity to strike over the seasons and the settlement date is the only one that is truly variable. Gunderson et al (1986) also use settlement date as the relevant contract date for the seasonal dummy variables. The variables LNEG, NEMP, and the regional and seasonal dummy variables are constructed from information on the MCA tape.
Gunderson et al (1986) also include in their incidence equation dummy variables for the five largest private-sector unions by total membership in Canada. In the present study, the same five union dummy variables are used. The five unions are: the United Autoworkers, Carpenters and Joiners, Food and Commercial Workers, United Steelworkers, and Teamsters. Swidinsky and Vanderkamp, and Gunderson et al (1986), also use industry dummy variables to control for possible variations in the propensity to strike across industries. Following their work, dummy variables for the 32 two-digit S.I.C. industries in the sample are constructed here. Gunderson et al (1986) find a significant reduction in strike activity during the period when the Anti-Inflation Board was in effect. Thus a variable (DAIB) takes a value of 1 if the effective date of the contract is during the period October 14, 1975 to April 14, 1978, and 0 otherwise.

Lastly, to control for cyclical factors, the unemployment rate is usually included as an independent variable in the estimation of strike incidence equations. Gunderson et al (1986) and Swidinsky and Vanderkamp find a negative and significant relationship between the rate of unemployment and strike incidence. Thus a variable (URM), defined as the seasonally adjusted unemployment rate for males aged 25 years and older in the month the contract is settled, is included in estimation of (3.1). Owing to changes by Statistics Canada in the method used to calculate the unemployment rate, the unemployment rate data are composed from two series. The first series, covering the period prior to 1966, is assembled from the Statistics Canada publication Historical Labour Force Statistics, (1973 edition), Catalogue 71-201, p.245. The second series, covering the period from 1966 onwards, is from the CANSIM University Base (4th Quarter, 1987) where the series number is D767657. Following Card (1987, p.33), the data prior to 1966 are adjusted to be consistent with the more recent series by multiplying each monthly unemployment rate by 0.8052.

For the econometric model estimated in chapter 5, information on the occurrence of a strike is insufficient; it is necessary to have strike duration information as well. However, the MCA tape does not contain information on strike length, it only provides information on strike incidence. Fortunately, David Card has matched strike duration information with the MCA tape contracts\(^7\) and he has, very graciously, made these data available for the present study. However, several caveats apply to the use of the Card data here.

\(^7\) See Card (1987) for details.
First, Card (1987) examines only manufacturing sector contracts, whereas the contract data discussed above have over 900 contracts from outside the manufacturing sector. There are 1,518 manufacturing sector contracts in the present data. These are matched with the Card data using the contract file number and the effective date of the contract, resulting in 1,516 successful matches, i.e., all but 2 of the manufacturing sector contracts in the present data are successfully allocated a strike duration from the Card data. Thus there are two data sets under consideration: the Full Sample with 2,459 contracts and a subsample of these data covering 1,516 contracts in the manufacturing sector.

Second, Card (1987) finds discrepancies between the MCA tape strike indicator and the publication used as the source of strike duration information. There are two sorts of discrepancy: either Strikes and Lockouts indicates a strike where, according to the MCA tape, no strike occurred, or vice versa, with the former being more frequent. Of the 1,516 contracts successfully merged with duration data, strikes occurred in 372 contracts according to the MCA tape, while strikes occurred in 409 contracts according to the Card data. Furthermore, Card is able to find strike duration information for only 393 of the 409 manufacturing contracts known to have had a strike. There are 349 contracts (93.8%) which indicate a strike in the Full Sample data and for which Card is able to find a strike duration. Therefore, the full 2,459 contract data set and the manufacturing subsample with 1,516 contracts are not directly comparable with respect to strike information.

Table 3.1 contains a summary of the variable definitions for the independent variables used in the estimation of the strike probability equation in (3.1). The table also lists the mean values for all the variables and the standard deviations of the continuous variables for the Full Sample, i.e., the data set with 2,459 contracts, and for the Manufacturing Subsample, i.e., the data set including 1,516 contracts from the manufacturing sector. Table 3.2 indicates the contract coverage, the number of strikes and the strike incidence across major S.I.C. industry groups. It should be noted that the gross strike incidence varies substantially across industries, and between the various sectors in the economy.

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8 The manufacturing sector is defined as all industries with a three-digit (1970) S.I.C. code from 101 to 399 inclusive.

9 The publication is Strikes and Lockouts, Labour Canada, Monthly, (Ottawa: Supply and Services Canada).
Table 3.1: Variable Names and Descriptive Statistics

<table>
<thead>
<tr>
<th>Variable Name</th>
<th>Variable Description</th>
<th>Full Sample Mean</th>
<th>Std. Dev.</th>
<th>Manufacturing Subsample Mean</th>
<th>Std. Dev.</th>
</tr>
</thead>
<tbody>
<tr>
<td>NEMP</td>
<td>Number of employees in the bargaining unit (1000's)</td>
<td>1.6255</td>
<td>2.4355</td>
<td>1.3489</td>
<td>1.7997</td>
</tr>
<tr>
<td>LNEG</td>
<td>Length of contract negotiations (months)</td>
<td>5.9711</td>
<td>3.0501</td>
<td>5.8840</td>
<td>2.7291</td>
</tr>
<tr>
<td>MANW</td>
<td>Average manufacturing wage rate (1971 dollars per hour)</td>
<td>3.5209</td>
<td>.45220</td>
<td>3.5030</td>
<td>.45358</td>
</tr>
<tr>
<td>URM</td>
<td>Monthly unemployment rate for males aged 25 and over</td>
<td>4.8402</td>
<td>1.9506</td>
<td>4.7884</td>
<td>1.9440</td>
</tr>
<tr>
<td>YEAR</td>
<td>Trend term equal to year of contract effective date</td>
<td>75.801</td>
<td>5.7671</td>
<td>75.369</td>
<td>5.6876</td>
</tr>
<tr>
<td>REVENUE</td>
<td>Firm revenue (billions of 1971 dollars)</td>
<td>.55307</td>
<td>.76402</td>
<td>.50603</td>
<td>.80202</td>
</tr>
<tr>
<td>C.V. REVENUE</td>
<td>Coefficient of variation of firm revenue</td>
<td>.30068</td>
<td>.21851</td>
<td>.28045</td>
<td>.1894</td>
</tr>
</tbody>
</table>

Dummy variables:

- DTO: Contracts after firm taken over | .087914 | .081135 |
- DRC: Contracts after an identified change in revenue series | .20008 | .19657 |
- DCM: Contracts after a gap in sequence of more than two years | .15535 | .12895 |
- Winter: January, February, or March | .19398 | .19459 |
- Spring: April, May, or June | .28426 | .28815 |
- Autumn: October, November, or December | .24909 | .22825 |
- Maritime provinces: Newfoundland, P.E.I., Nova Scotia, or New Brunswick | .044327 | .038259 |
- Quebec: Quebec | .3217 | .35752 |
- Ontario: Ontario | .37373 | .45778 |
- Prairie provinces: Manitoba, Saskatchewan, or Alberta | .095097 | .046334 |
- British Columbia | .091084 | .080896 |
- Autoworkers: United Automobile Workers of America (U.A.W.) | .086894 | .10618 |
- Carpenters: United Brotherhood of Carpenters and Joiners of America | .023180 | .000 |
- Food and Commercial: United Food and Commercial Workers | .051847 | .031682 |
- Steelworkers: United Steelworkers of America (U.S.W.) | .14559 | .14249 |
- Teamsters: International Brotherhood of Teamsters | .0089487 | .0052770 |

Number of contracts with a strike (2) | 501 | 393 |
Number of contracts without a strike | 1958 | 1123 |
Total number of contracts | 2459 | 1516 |
Strike Incidence | .20374 | .25923 |

Notes: (1) The mean of the dummy variables is equal to the sample proportion.
(2) The number of observations with a strike for the manufacturing subsample is the number for which Card (1987) is able to find a positive strike duration. This differs from the number of strikes originally reported in the data (372), and the number for which Card (1987) reports the occurrence of a strike (409). See section 3.3 for the relevant details.
Table 3.2: Number of Contracts, Strikes and Strike Incidence by Industry

<table>
<thead>
<tr>
<th>S.I.C. Code</th>
<th>Industry</th>
<th>Contracts</th>
<th>Strikes</th>
<th>Incidence</th>
</tr>
</thead>
<tbody>
<tr>
<td>Primary Industries:</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>031</td>
<td>Logging</td>
<td>92</td>
<td>11</td>
<td>.120</td>
</tr>
<tr>
<td>051-059</td>
<td>Metal Mines</td>
<td>152</td>
<td>31</td>
<td>.204</td>
</tr>
<tr>
<td>061,064</td>
<td>Mineral Fuels</td>
<td>29</td>
<td>0</td>
<td>.276</td>
</tr>
<tr>
<td>071-079</td>
<td>Non-Metal Mines</td>
<td>27</td>
<td>7</td>
<td>.259</td>
</tr>
<tr>
<td>TOTAL</td>
<td></td>
<td>300</td>
<td>57</td>
<td>.190</td>
</tr>
<tr>
<td>Manufacturing Industries:</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>101-109</td>
<td>Food and Beverage</td>
<td>181</td>
<td>24</td>
<td>.133</td>
</tr>
<tr>
<td>151,153</td>
<td>Tobacco Products</td>
<td>30</td>
<td>3</td>
<td>.100</td>
</tr>
<tr>
<td>162,165</td>
<td>Rubber and Plastics</td>
<td>14</td>
<td>4</td>
<td>.286</td>
</tr>
<tr>
<td>181-189</td>
<td>Textile Industries</td>
<td>114</td>
<td>22</td>
<td>.193</td>
</tr>
<tr>
<td>231,239</td>
<td>Knitting Mills</td>
<td>4</td>
<td>0</td>
<td>0.0</td>
</tr>
<tr>
<td>243-249</td>
<td>Clothing Industries</td>
<td>9</td>
<td>0</td>
<td>0.0</td>
</tr>
<tr>
<td>251-259</td>
<td>Wood Industries</td>
<td>18</td>
<td>6</td>
<td>.333</td>
</tr>
<tr>
<td>261-268</td>
<td>Furniture Industries</td>
<td>11</td>
<td>4</td>
<td>.364</td>
</tr>
<tr>
<td>271-274</td>
<td>Pulp and Paper Industries</td>
<td>266</td>
<td>75</td>
<td>.282</td>
</tr>
<tr>
<td>286-289</td>
<td>Printing and Publishing</td>
<td>10</td>
<td>1</td>
<td>.100</td>
</tr>
<tr>
<td>291-298</td>
<td>Primary Metals Industries</td>
<td>213</td>
<td>48</td>
<td>.225</td>
</tr>
<tr>
<td>301-309</td>
<td>Metal Fabricating</td>
<td>30</td>
<td>6</td>
<td>.200</td>
</tr>
<tr>
<td>311-318</td>
<td>Machinery Industries</td>
<td>41</td>
<td>13</td>
<td>.317</td>
</tr>
<tr>
<td>321-329</td>
<td>Transportation Equipment</td>
<td>212</td>
<td>91</td>
<td>.429</td>
</tr>
<tr>
<td>331-339</td>
<td>Electrical Products</td>
<td>217</td>
<td>47</td>
<td>.217</td>
</tr>
<tr>
<td>351-359</td>
<td>Non-Metallic Mineral Products</td>
<td>74</td>
<td>16</td>
<td>.216</td>
</tr>
<tr>
<td>365,369</td>
<td>Petroleum and Coal Products</td>
<td>6</td>
<td>1</td>
<td>.167</td>
</tr>
<tr>
<td>372-379</td>
<td>Chemical Products</td>
<td>55</td>
<td>11</td>
<td>.200</td>
</tr>
<tr>
<td>391-399</td>
<td>Miscellaneous Manufacturing</td>
<td>13</td>
<td>0</td>
<td>0.0</td>
</tr>
<tr>
<td>TOTAL</td>
<td></td>
<td>1518</td>
<td>372</td>
<td>.245</td>
</tr>
<tr>
<td>Transportation, Communications and Other Utilities:</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>501-519</td>
<td>Transportation</td>
<td>149</td>
<td>12</td>
<td>.081</td>
</tr>
<tr>
<td>543-548</td>
<td>Communications</td>
<td>72</td>
<td>8</td>
<td>.111</td>
</tr>
<tr>
<td>572-579</td>
<td>Electric; Gas and Water</td>
<td>84</td>
<td>16</td>
<td>.190</td>
</tr>
<tr>
<td>TOTAL</td>
<td></td>
<td>305</td>
<td>36</td>
<td>.118</td>
</tr>
<tr>
<td>Trade:</td>
<td></td>
<td>305</td>
<td>36</td>
<td>.118</td>
</tr>
<tr>
<td>611-619</td>
<td>Trade, Wholesale</td>
<td>52</td>
<td>7</td>
<td>.135</td>
</tr>
<tr>
<td>631</td>
<td>Trade, Retail</td>
<td>206</td>
<td>20</td>
<td>.097</td>
</tr>
<tr>
<td>701-707</td>
<td>Finance Industries</td>
<td>10</td>
<td>2</td>
<td>.200</td>
</tr>
<tr>
<td>TOTAL</td>
<td></td>
<td>268</td>
<td>29</td>
<td>.108</td>
</tr>
<tr>
<td>Services:</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>861-869</td>
<td>Business Services</td>
<td>8</td>
<td>1</td>
<td>.125</td>
</tr>
<tr>
<td>881-886</td>
<td>Accomodation and Food Services</td>
<td>56</td>
<td>4</td>
<td>.071</td>
</tr>
<tr>
<td>891-899</td>
<td>Miscellaneous Services</td>
<td>4</td>
<td>2</td>
<td>.500</td>
</tr>
<tr>
<td>TOTAL</td>
<td></td>
<td>68</td>
<td>7</td>
<td>.103</td>
</tr>
<tr>
<td>TOTAL, ALL INDUSTRIES</td>
<td>2459</td>
<td>501</td>
<td>.204</td>
<td></td>
</tr>
</tbody>
</table>

3.4 Results From Estimation of the Strike Incidence Equation

Equation (3.1) is first estimated by OLS.\textsuperscript{10} Since the dependent variable takes the value of 1 only when a strike occurs, the fitted value of $DS_i$ is interpreted as the probability of a strike given the values of the independent variables. In addition, estimated coefficients are interpreted as the increase in the probability of a strike occurring, given a unit increase in the corresponding independent variable. For this reason, OLS estimation of (3.1) is also called the linear probability model. The disturbance terms, $e_i$, in (3.1) are assumed to have a zero mean and to be pairwise uncorrelated. However, as Kmenta (1971) points out, the dichotomous dependent variable means that the variance of $e_i$ depends on $i$, that is the $e_i$ are heteroskedastic. Thus it is assumed that $E(e_i) = 0$ and $E(e_i^2) = \sigma_i^2$, for all $i$, and $E(e_i e_j) = 0$ for $i \neq j$. The OLS coefficient estimates from (3.1) are corrected for an unknown form of heteroskedasticity following the procedure in White (1980). The omitted dummy variables are: the North and multi-province region dummy variables, the summer seasonal dummy variable, and the miscellaneous manufacturing, knitting and clothing industries dummy variables (S.I.C. numbers 23, 24 and 39 respectively).

Table 3.3 contains the coefficient estimates of the linear probability model of strike incidence. The first column of Table 3.3 contains the means of the data, with the variables suggested by the adverse-selection model of strikes listed first. Generally speaking, the coefficient estimates in Table 3.3 confirm the predictions of the theoretical model of strikes. The coefficient on REVENUE indicates that, as predicted by the theory, strike incidence is negatively related to the level of demand at the firm as measured by the firm's revenue, and the estimate is statistically significant at the .10 level. The coefficient on C.V.REVENUE is positive and statistically significant at the .05 level, indicating that greater variability in firm revenue increases the likelihood of a strike. Again this confirms a prediction of the model. Though the effects of the adverse-selection variables REVENUE and C.V.REVENUE are statistically significant, the effects of the variables on the probability of a strike are quite small. The coefficient on REVENUE indicates that a $1$ billion increase in firm revenue reduces the propensity to strike by 2%. Therefore, given average firm revenue is slightly over $0.5$ billion, a 20% increase in revenue will reduce the average propensity to strike by 0.2%, \textit{ceteris paribus}. Similarly, the coefficient

\textsuperscript{10} The SHAZAM computer econometric package [(White (1978, 1987))] is used for all the regressions.
Table 3.3: OLS Estimates of Strike Incidence(1)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Gross strike rate(3)</th>
<th>OLS coefficient</th>
<th>Standard error(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Adverse-selection variables:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>REVENUE</td>
<td>.553</td>
<td>N.A.</td>
<td>-.0191 *</td>
</tr>
<tr>
<td>C.V.REVENUE</td>
<td>.301</td>
<td>N.A.</td>
<td>-.0899 **</td>
</tr>
<tr>
<td>DTO</td>
<td>.068</td>
<td>.18</td>
<td>-.0001</td>
</tr>
<tr>
<td>DRC</td>
<td>.200</td>
<td>.20</td>
<td>-.0423</td>
</tr>
<tr>
<td>DCM</td>
<td>.155</td>
<td>.24</td>
<td>.0122</td>
</tr>
<tr>
<td>Other contract-specific variables:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>NEMP</td>
<td>1.626</td>
<td>N.A.</td>
<td>.0075 *</td>
</tr>
<tr>
<td>LNEG</td>
<td>5.971</td>
<td>N.A.</td>
<td>.0323 **</td>
</tr>
<tr>
<td>YEAR</td>
<td>75.601</td>
<td>N.A.</td>
<td>.0195 **</td>
</tr>
<tr>
<td>DAIB</td>
<td>.141</td>
<td>-.1126 **</td>
<td>.007</td>
</tr>
<tr>
<td>URM</td>
<td>4.840</td>
<td>N.A.</td>
<td>-.0567 **</td>
</tr>
<tr>
<td>Region:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Multi-province,</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Yukon, N.W.T.</td>
<td>.105</td>
<td>.09</td>
<td>--</td>
</tr>
<tr>
<td>Maritimes</td>
<td>.044</td>
<td>.15</td>
<td>.0029</td>
</tr>
<tr>
<td>Quebec</td>
<td>.321</td>
<td>.22</td>
<td>.0621</td>
</tr>
<tr>
<td>Ontario</td>
<td>.374</td>
<td>.25</td>
<td>.0450</td>
</tr>
<tr>
<td>Prairies</td>
<td>.065</td>
<td>.11</td>
<td>-.0099</td>
</tr>
<tr>
<td>British Columbia</td>
<td>.091</td>
<td>.21</td>
<td>.0135</td>
</tr>
<tr>
<td>Season:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Winter</td>
<td>.194</td>
<td>.20</td>
<td>-.0350</td>
</tr>
<tr>
<td>Spring</td>
<td>.284</td>
<td>.17</td>
<td>-.0464 **</td>
</tr>
<tr>
<td>Summer</td>
<td>.272</td>
<td>.24</td>
<td>--</td>
</tr>
<tr>
<td>Autumn</td>
<td>.250</td>
<td>.21</td>
<td>-.0358</td>
</tr>
<tr>
<td>Specific unions:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Autoworkers</td>
<td>.067</td>
<td>.57</td>
<td>.3587 **</td>
</tr>
<tr>
<td>Carpenters and joiners</td>
<td>.023</td>
<td>.11</td>
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<td>.24</td>
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<td>.0635</td>
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<td>.17</td>
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<td>.00</td>
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<td>Tobacco</td>
<td>.012</td>
<td>.10</td>
<td>.1140 *</td>
</tr>
</tbody>
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... continued overleaf

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### Table 3.3: Continued

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<th>Variable</th>
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<td>Textiles</td>
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<td>.1871 **</td>
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<td>Furniture</td>
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<td>.4242 **</td>
</tr>
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<td>.2300 **</td>
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<td>.0336</td>
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<td>.2109 **</td>
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<td>Machinery</td>
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<td>.2377 **</td>
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<td>.086</td>
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<tr>
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<td>.0566</td>
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<td>Retail trade</td>
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<td>.0874 **</td>
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<td>.2057</td>
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<tr>
<td>Personal services</td>
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<td>.2202 *</td>
</tr>
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<td>Accomodation</td>
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<td>.0425</td>
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<tr>
<td>Miscellaneous services</td>
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<td>.3171 *</td>
</tr>
<tr>
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<td>N.A.</td>
</tr>
</tbody>
</table>

---

Notes: N.A. indicates not applicable.

- (1) There are 2,459 observations (contracts).
- (2) For the dummy variable categories the mean is the proportion of observations in each category.
- (3) Proportion of contracts signed after a strike. Overall mean is 501/2459 = .204
- (4) Significance is denoted by ** at the .05 and * at the .10 level, where the critical values, respectively, are 1.65 and 1.28 for the one-tailed test (when the expected sign is unambiguous) and 1.96 and 1.65 for the two-tailed test (when the expected sign is ambiguous).
on C.V.REVENUE implies that a unit increase in the coefficient of variation of firm revenue (which is unit-independent) increases the propensity to strike by 9%. It is perhaps surprising, given the magnitude of these effects, that the effects are statistically significant. However, these results are quite robust to changes in sample size, estimation technique, and explanatory variables, as will become apparent below.

The estimated coefficient on DTO is negative but it is not statistically significant, indicating some evidence that strikes are less likely for contracts which follow a change in firm ownership. The remaining estimates in Table 3.3 are, with two exceptions, similar to the results obtained by Gunderson et al (1986) and Swidinsky and Vanderkamp (1982). Larger bargaining units have a significantly greater propensity to strike; for every additional thousand workers in the bargaining unit the probability of a strike increases by 0.75%. The coefficient on YEAR indicates roughly a 2% increase per year in the probability of a strike over the period 1964–1986. The propensity to strike during the period when the Anti-Inflation Board was setting contract wage guidelines is about 11% below the average. An increase of 1% in the unemployment rate of males aged 25 and over decreases the probability of a strike by 5.67%, as indicated by the coefficient on URM.\footnote{Estimates with an alternative unemployment rate—the unemployment rate for both sexes aged 15 and older—are similar but less precise.}

The two exceptional results are the coefficient on LNEG, which Swidinsky and Vanderkamp (1982) report to be negative and significant, and the coefficients on the seasonal dummy variables, which have opposite signs to those estimated by Gunderson et al (1986). The discrepancies are almost certainly due to differences between the data used in the present study and the data used in the earlier studies. For example, Swidinsky and Vanderkamp (1982) examine the time period 1965–75, limit the contract sample to the manufacturing sector and include some bargaining units with less than 500 employees. The coefficient signs on the seasonal dummy variables reported by Gunderson et al (1986) are replicated using the present data if the seasonally unadjusted unemployment rate is used in place of the seasonally adjusted data. To control for seasonal variation in the unemployment rate, which is substantial in Canada,\footnote{In 1986, for example, the seasonally unadjusted unemployment rate for males and females 15 years and older varies from 8.7% in October to 10.9% in March.} it is desirable to use seasonally adjusted data.

To summarize, the estimates in Table 3.3 indicate there is no statistically significant regional variation in the propensity to strike. Relative to the summer, a
significantly lower incidence of strikes is found for contracts settled in the spring at the .05 level and, at the .05 level, the United Autoworkers exhibit a significantly greater propensity to strike relative to other unions, while the United Steelworkers exhibit the same tendency at the .10 level. There are 18 industries with a significantly greater propensity to strike at the .05 level, and 5 industries with a significantly greater propensity to strike at the .10 level, relative to the omitted industries (miscellaneous manufacturing, knitting and clothing). All of the 12 industries found by Gunderson et al to have statistically significantly different strike behaviour (at the .10 level) relative to the control group are found here to have a significant effect on strike incidence. In addition, both the United Autoworkers and the United Steelworkers are found to have a positive effect on strike incidence in this study, duplicating the result of Gunderson et al.

The estimates in Table 3.3 represent the basic model of strike incidence. However, there are two important sources of variation in firm revenue which have not been controlled for in the estimation. First, real firm revenue tends to grow over time as the economy expands. Second, firm revenue is subject to fluctuations due to the business cycle. It is reasonable to assume that the union is fully aware of both sources of variation in firm revenue, or in other words to assume that the union has rational expectations about firm revenue. In this way, the asymmetry of information between the firm and the union is captured by the component of firm revenue that is not explained by long-term trends and cyclical fluctuations. Hence, the next step in the empirical analysis of the model is to remove the systematic variations due to time and the business cycle from the revenue data.

The unsystematic component of firm revenue may be specified as the residual from a regression of the natural logarithm of revenue on a trend term and the trend term squared. The business cycle is modelled by a first-order autoregressive process. To control for situations where the revenue series associated with bargaining units is subject to an infrequent event, DRC is also included as an independent variable in the trend regressions. To ensure adequate degrees of freedom, all revenue series with the same number of observations are pooled. The model is described in the following two equations

\[ \ln R E V_{it} = \gamma_1 + \gamma_2 T_{it} + \gamma_3 T_{it}^2 + \gamma_4 DRC_{it} + u_{it} \]  

(3.2)

where \( t = 1, \ldots, NR_j \) is the number of observations in the \( j \)th cross-section of firms, \( i \) indicates the firm in the \( j \)th cross-section, \( i = 1, \ldots, NF_j \), and where \( NF_j \) is the
number of firms with \( NR_j \) revenue observations. The disturbance term is assumed to follow a first-order autoregressive process thus

\[
_u \text{it} = \rho u_{i, t-1} + \epsilon_{it}
\]  

(3.3)

where \( E(\epsilon_{it}) = 0 \), \( E(\epsilon_{it}\epsilon_{jt}) = 0 \), and \( E(\epsilon_{it}\epsilon_{js}) = 0 \) for \( t \neq s \). In other words, disturbances are independent across firms and the autocorrelation coefficient is constrained to be the same for all firms.\(^{13}\) The estimated residual \( \hat{\epsilon}_{it} \), that is the residual from the transformed data, is the desired measure of *adjusted* revenue. The coefficient of variation of revenue is calculated using the adjusted revenue series for each firm. The pooled regressions for the firms with 4 and 5 revenue observations were estimated to have unstable coefficients of autocorrelation. These firms are deleted from the sample, reducing the number of contract observations to 2,356.

Results from estimation of the strike incidence equation where the variables \( \text{REVENUE} \) and \( \text{C.V.REVENUE} \) are calculated from the residuals of the model in (3.2) and (3.3) are shown in Table 3.4.\(^{14}\) Comparing the coefficient estimates of the asymmetric information variables from the adjusted data in Table 3.4 to the estimates with the unadjusted revenue data in Table 3.3, notice that the effect of revenue and the variability in revenue on strike incidence has increased. In addition, both variables remain statistically significant in Table 3.4; the coefficient on \( \text{REVENUE} \) is significant at the .05 level and the coefficient on \( \text{C.V.REVENUE} \) is significant at the .10 level.

Remember that the strike incidence data for the Full Sample, which has been used in the estimation so far, are reported by Card (1987) to have some errors in the strike information. To determine if these errors seriously affect the present results, the strike incidence model is estimated over the Manufacturing Subsample of the data using the information on the occurrence of a strike made available by David Card. For comparison, the strike incidence equation is also estimated over the same subsample of the data using the information on the occurrence of a strike from the MCA tape. The results are reported in Appendix A. Estimates of the strike incidence equation using the strike information available from the MCA tape are in Table A.1, and estimates using the Card (1987) strike information are shown in Table A.2.

---

\(^{13}\) This is consistent with the assumption that the business cycle affects all firms' revenues in the same manner.

\(^{14}\) The variable DRC is not included in this version of the incidence equation because its effect on revenue has already been taken into account, as can be seen from (3.2).
Table 3.4: OLS Estimates of Strike Incidence, Firm Revenue Adjusted for Trend and Cyclical Effects(1)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean</th>
<th>Gross strike rate</th>
<th>OLS coefficient</th>
<th>Standard error</th>
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<td>-0.0373</td>
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<td>0.0106</td>
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<td>N.A.</td>
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<td>0.004</td>
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<td>--</td>
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<td>0.15</td>
<td>0.0107</td>
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<td>0.00</td>
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<td>0.0822 *</td>
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<td>0.10</td>
<td>0.0622</td>
<td>0.065</td>
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... continued overleaf
Table 3.4: Continued

<table>
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<tr>
<th>Variable</th>
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<th>Gross strike rate</th>
<th>OLS coefficient</th>
<th>Standard error</th>
</tr>
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</tr>
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<td>Rubber and plastics</td>
<td>0.006</td>
<td>0.29</td>
<td>0.2297 **</td>
<td>0.115</td>
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<td>0.1414 **</td>
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<td>0.008</td>
<td>0.33</td>
<td>0.2260 *</td>
<td>0.118</td>
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<td>0.36</td>
<td>0.3586 **</td>
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<td>0.1461 **</td>
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</tr>
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<td>0.0855</td>
<td>0.082</td>
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<td>0.1704 **</td>
<td>0.075</td>
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<td>0.2068 **</td>
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<td>0.115</td>
</tr>
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<td>Chemicals</td>
<td>0.022</td>
<td>0.20</td>
<td>0.1773 **</td>
<td>0.064</td>
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<tr>
<td>Transportation</td>
<td>0.059</td>
<td>0.07</td>
<td>-0.0222</td>
<td>0.055</td>
</tr>
<tr>
<td>Communications</td>
<td>0.031</td>
<td>0.11</td>
<td>0.0621</td>
<td>0.068</td>
</tr>
<tr>
<td>Utilities</td>
<td>0.036</td>
<td>0.19</td>
<td>0.0343</td>
<td>0.058</td>
</tr>
<tr>
<td>Wholesale trade</td>
<td>0.022</td>
<td>0.13</td>
<td>0.0641</td>
<td>0.069</td>
</tr>
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<td>Retail trade</td>
<td>0.081</td>
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<td>0.1309</td>
<td>0.131</td>
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<td>0.003</td>
<td>0.13</td>
<td>0.1721</td>
<td>0.118</td>
</tr>
<tr>
<td>Accomodation</td>
<td>0.020</td>
<td>0.06</td>
<td>-0.0475 *</td>
<td>0.056</td>
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<td>N.A.</td>
<td>-1.369</td>
<td>0.168</td>
</tr>
</tbody>
</table>

Notes: N.A. indicates not applicable.

(1) There are 2,356 observations (contracts). Overall proportion of strikes is 485/2356 = .206. For an explanation of the terms and symbols used in this table see the notes after Table 3.3.
In both sets of estimates, the coefficient signs on REVENUE and C.V.REVENUE have the signs predicted by the theory, however neither of the two coefficients is statistically significant in Table A.1 and Table A.2. The coefficient estimates on the remaining explanatory variables using the original and corrected strike data are quite similar. Thus it appears that the missclassification of some strikes in the MCA tape is not too serious a problem. In addition, it appears that the adverse-selection model variables do not play a significant role in the occurrence of strikes in the manufacturing subsample. Nevertheless, as the missclassification problem does not appear to be serious, it is reasonable to expect that, if the correct strike incidence data were available for the Full Sample, the results reported in Tables 3.3 and 3.4 would be replicated.

Although the results so far support the adverse-selection model, it is possible that the negative effect of revenue on strike incidence reported in Tables 3.3 and 3.4 is simply reflecting a fact that strikes cause a reduction in firm revenue, i.e., the direction of causality is the reverse of that predicted by the adverse-selection model. Furthermore, if strikes do decrease revenue then, *ceteris paribus*, firms which have strikes will have more variable revenue than firms which do not have strikes. This could also explain a positive coefficient on C.V.REVENUE. Therefore, the next step in the analysis is to examine the hypothesis that strikes cause a reduction in revenue.

The most obvious way to examine the hypothesis is to estimate a model with firm revenue as the dependent variable and include a dummy variable, which takes a value of 1 for years in which there is a strike and 0 otherwise, as an independent variable.\(^\text{15}\) The model of firm revenue considered is written

\[
\text{REV}_{it} = \beta_1 + \beta_2 \text{YEAR}_{it} + \beta_3 \text{DSTRIKE}_{it} + \beta_4 NEMP_{it} + \beta_5 UR_{it} + \beta_6 \text{DRDC}_{it} + \epsilon_{it}
\]

(3.4)

where \(\text{REV}_{it}\) is the revenue (in 1971 dollars) of the \(i\)th firm in year \(t\); \(\text{YEAR}\) is a trend; \(\text{DSTRIKE}\) is the strike dummy variable; \(\text{NEMP}\) is the number of employees at the \(i\)th firm in year \(t\);\(^\text{16}\) \(\text{UR}\) is the unemployment rate for both sexes aged 15 years and older in year \(t\); \(\text{DRDC}\) is a dummy variable with a value of 1 for years after

---

\(^{15}\) According to the theoretical model, the occurrence of a strike indicates the presence of a low revenue firm. Thus there is possibly some degree of simultaneity present in this specification. One way around the problem would be to use an instrument for the revenue variable, e.g. lagged revenue.

\(^{16}\) The proxy for the number of employees is simply the size of the bargaining unit in the most recent contract.
a change in the firm’s revenue series, and 0 otherwise. Initially, it is assumed that $E(u_{it}) = 0$, $E(u_{it}^2) = \sigma_i^2$, and $E(u_{it}u_{js}) = 0$ where $t \neq s$. Thus $u_{it}$ is a serially and contemporaneously uncorrected disturbance term with a zero mean and a constant variance for each firm. To ensure adequate degrees of freedom, (3.4) is estimated using OLS for every revenue series with 10 or more observations during which at least one strike occurred.

The results of the individual firm regressions are summarized in Table 3.5. Given the estimation procedure, there is scant evidence of a significant negative effect of strikes on firm revenue. In slightly more than 60% of the regressions, the coefficients on the variable DSTRIKE are negative, but only 15 of the 122 coefficients are significantly different from zero. In contrast, there is, not surprisingly, overwhelming evidence that firm revenue has grown over time; 95% of the coefficients on YEAR are positive, and 135 of the 190 coefficients are significantly different from zero. Firm revenue appears to be procyclical; 166 coefficients on UR are negative, of which 78 are significantly different from zero, compared to 35 positive coefficients, of which only 4 are statistically significant. In 42 of the 201 regressions the revenue series was found to be subject to an infrequent event. The coefficient signs on DRDC indicate that, in 29 of the 42 cases, firm revenue jumps discontinuously up, and in 13 cases it jumps down.

The major deficiency with using individual firm regressions to test the hypothesis that strikes have a negative effect on firm revenue is the fact that there are only a small number of observations with a strike. To overcome this problem it is possible to pool all the firms with the same number of revenue observations, and use the cross-sectionally heteroskedastic and time-wise autoregressive model proposed in Kmenta (1971). The model is described in the following two equations

\[ REV_{it} = \delta_1 + \delta_2 YEAR_{it} + \delta_3 DSTRIKE_{it} + \delta_4 NEMP_{it} + \delta_5 UR_{it} + \delta_6 DRDC_{it} + u_{it} \]

where $t = 1, \ldots, NR_j$ is the number of observations in a particular cross-section of firms, and $i = 1, \ldots, NF_j$ indicates the $i$th firm in the same cross-section, and where $NF_j$ is the number of firms with $NR_j$ revenue observations, as above.\(^{17}\) The

\(^{17}\) It should be clear that this model is different from (3.2) and (3.3): this model is aimed at testing a hypothesis about firm revenue; the earlier model is designed to remove the effects of trends and cycles.
Table 3.5: Summary of Estimated Coefficient Signs from Individual Regressions of Firm Revenue

<table>
<thead>
<tr>
<th>Variable</th>
<th>Positive</th>
<th>Negative</th>
</tr>
</thead>
<tbody>
<tr>
<td>YEAR</td>
<td>190 (135)</td>
<td>11 (1)</td>
</tr>
<tr>
<td>DSTRIKE</td>
<td>79 (4)</td>
<td>122 (15)</td>
</tr>
<tr>
<td>NEMP</td>
<td>118 (39)</td>
<td>83 (23)</td>
</tr>
<tr>
<td>UR</td>
<td>35 (4)</td>
<td>166 (78)</td>
</tr>
<tr>
<td>DRDC(2)</td>
<td>29 (23)</td>
<td>13 (4)</td>
</tr>
</tbody>
</table>

Total Number of Regressions 201

Notes: (1) The number of coefficients which are significant at the .05 level is given in parentheses.

(2) Applies only to the 42 regressions where the revenue figure associated with the firm is subject to an infrequent event (see text for details).
disturbance term is assumed to follow a first-order autoregressive process thus

\[ u_{it} = \rho u_{i,t-1} + \epsilon_{it} \] (3.6)

where \( E(\epsilon_{it}) = 0 \), \( E(\epsilon_{it}\epsilon_{jt}) = 0 \), and \( E(\epsilon_{it}\epsilon_{js}) = 0 \) for \( t \neq s \). In other words, disturbances are independent across firms and the autocorrelation coefficient is constrained to be the same for all firms.\(^{18}\)

Pooling all firms with the same number of observations, and estimating the above model for each pool yields the results in Table 3.6 for firms with more than 10 and fewer than 23 revenue observations.\(^{19}\) As expected, the estimates from the pooled regressions echo the results from the individual OLS regressions, though the pooled estimates have lower standard errors. As most of the estimates reported in Table 3.6 have a coefficient of autocorrelation which is close to 1, the same pooled regressions are calculated using data which are first-differenced. The results are very similar to those reported in Table 3.6. YEAR is found to have a positive and significant effect on revenue, and UR is found to have a negative and significant effect, mirroring the majority of the results in Table 3.5. The main variable of interest is the coefficient on DSTRIKE. The pooled regressions for firms with 12 and 14 revenue observations, yield a positive but not statistically significant coefficient estimate for DSTRIKE. The coefficient estimate on DSTRIKE is estimated to be negative and significant at the .05 level in the regressions with 19 and 20 revenue observations. Even so, the effect of strikes on revenue is small. The largest coefficient on DSTRIKE is from the regression with 15 revenue observations per firm, and it indicates that strikes decrease revenue by $34.6 million, which is just over 6% of average firm revenue.

To summarize, the results from the original strike incidence equation estimated are subject to the criticism that the coefficients on the key asymmetric information variables REVENUE and C.V.REVENUE are not measuring the effects for which they are intended. The signs on the estimated coefficients of these variables could simply be due to the fact that strikes reduce firm revenue. Estimates from individual and pooled regressions of firm revenue provide little evidence that strikes reduce revenue. Firms are probably able to use inventories to sell some output during a strike. This

\(^{18}\) In theory, allowance can be made for contemporaneous correlation and/or a different \( \rho \) for each firm. However, individual firm autocorrelation coefficients were often estimated to be greater than 1.

\(^{19}\) There is only one firm with 23 revenue observations.
Table 3.6: Coefficient Estimates from Pooled Regressions of Firm Revenue(1)

<table>
<thead>
<tr>
<th>Variable</th>
<th>11</th>
<th>12</th>
<th>13</th>
<th>14</th>
<th>15</th>
<th>16</th>
</tr>
</thead>
<tbody>
<tr>
<td>YEAR</td>
<td>0.0123**</td>
<td>0.0119**</td>
<td>0.0153**</td>
<td>0.0067**</td>
<td>0.0283**</td>
<td>0.0164**</td>
</tr>
<tr>
<td>DSTRIKE</td>
<td>-0.0061</td>
<td>0.0066</td>
<td>-0.0190</td>
<td>0.0014</td>
<td>-0.0346</td>
<td>-0.0103</td>
</tr>
<tr>
<td>NEMP</td>
<td>0.0130**</td>
<td>-0.0071</td>
<td>-0.0101**</td>
<td>-0.0041</td>
<td>-0.0342**</td>
<td>-0.0140**</td>
</tr>
<tr>
<td>DRDC</td>
<td>0.1680</td>
<td>-0.0005</td>
<td>0.0171</td>
<td>0.0005</td>
<td>0.0100</td>
<td>0.0006</td>
</tr>
<tr>
<td>Constant</td>
<td>-0.6098*</td>
<td>-0.6526**</td>
<td>-0.7507**</td>
<td>-0.3552**</td>
<td>-1.5413**</td>
<td>-0.9224**</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th>17</th>
<th>18</th>
<th>19</th>
<th>20</th>
<th>21</th>
</tr>
</thead>
<tbody>
<tr>
<td>YEAR</td>
<td>0.0135**</td>
<td>0.0160**</td>
<td>0.0043**</td>
<td>0.0217**</td>
<td>0.0160**</td>
</tr>
<tr>
<td>DSTRIKE</td>
<td>-0.0104</td>
<td>-0.0084</td>
<td>-0.0066**</td>
<td>-0.0366**</td>
<td>-0.0031</td>
</tr>
<tr>
<td>NEMP</td>
<td>0.0418**</td>
<td>0.0325*</td>
<td>0.0098</td>
<td>0.0087**</td>
<td>0.0280**</td>
</tr>
<tr>
<td>DRDC</td>
<td>0.0341</td>
<td>0.0140</td>
<td>0.3058**</td>
<td>0.3474**</td>
<td>0.0693**</td>
</tr>
<tr>
<td>Constant</td>
<td>-0.7847**</td>
<td>-0.8483**</td>
<td>-0.2315**</td>
<td>-0.9874**</td>
<td>-0.8443**</td>
</tr>
</tbody>
</table>

Notes: (1) Significance is denoted by ** at the .05 and * at the .10 level for a two-tailed test with degrees of freedom equal to (NR NF-6) for each respective regression. Standard errors are given in parentheses.

(2) As defined by Buse (1973).

(3) NF refers to the number of firms in each regression.

(4) The coefficient of autocorrelation is estimated by the correlation coefficient and is constrained to be the same across all firms in each regression.
would tend to reduce the effect of strikes on revenue, and thus reduce the severity of the reverse causality problem. For a discussion of the role of inventories in offsetting the costs of strikes to firms see Christenson (1953). Paarsch (1988) finds evidence that inventories in the lumber industry of British Columbia are reduced prior to strikes, which may be interpreted as support for the offset theory.

The final phase of the econometric analysis of strike incidence takes into account the fact that the dependent variable in the incidence equation, $DS_i$, is not continuous. The dependent variable in the linear probability model is interpreted as the predicted probability of a strike and, therefore, it is constrained by theory to lie in the unit interval. Unfortunately, nothing in the OLS estimation procedure ensures that this will be the case. To avoid this problem, the right-hand side of (3.1) may be transformed, leading to the logit and the probit models. The logit model is used here because of its numerical simplicity. With $DS_i$ defined the same way as in (3.1), the logistic model of strike incidence is

$$DS_i = \frac{1}{1 + \exp(-z_i'b)}$$

(3.7)

where $z_i$ is shorthand for the independent variables in (3.1), and $b$ is a vector of coefficients to be estimated.

In the linear probability model, estimated coefficients are interpreted as the increase in the probability of a strike occurring, given a unit increase in the corresponding independent variable. In the logit model, estimated coefficients do not have the same interpretation because the right-hand side of (3.7) is a non-linear function of $z_i$. Instead, the change in the probability of a strike occurring given a unit increase in a continuous independent variable is

$$\frac{\partial DS_i}{\partial z_{ij}} = \frac{b_j \exp(-z_i'b)}{[1 + \exp(-z_i'b)]^2}$$

(3.8)

For discontinuous independent variables, like the dummy variables, the above expression is not valid. To calculate the correct expression, it is necessary to incorporate the discontinuous change in the variable into the logistic function. Thus, using (3.7)

$$DS_i + \Delta DS_i = \frac{1}{1 + \exp(-z_i'b - b_j\Delta z_{ij})}$$

(3.9)

Since $\Delta z_{ij} = 1$ for a dummy variable

$$\Delta DS_i = \frac{1}{1 + \exp(-z_i'b - b_j)} - DS_i$$

(3.10)
Table 3.7: Logit Estimates of Strike Incidence Without Industry Effects, Firm Revenue Adjusted for Trend and Cyclic Effects (1)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Logit coefficient</th>
<th>Change in probability</th>
<th>Standard error</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Adverse-selection variables:</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>REVENUE</td>
<td>.1765</td>
<td>.0289 **</td>
<td>.064</td>
</tr>
<tr>
<td>C.V.REVENUE</td>
<td>.7215</td>
<td>1.179 **</td>
<td>.435</td>
</tr>
<tr>
<td>DTO</td>
<td>-.1721</td>
<td>.0267</td>
<td>.226</td>
</tr>
<tr>
<td>DCM</td>
<td>.0439</td>
<td>.0073</td>
<td>.151</td>
</tr>
<tr>
<td><strong>Other contract-specific variables:</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>NEMP</td>
<td>.0295</td>
<td>.0048</td>
<td>.023</td>
</tr>
<tr>
<td>LNEG</td>
<td>.2083</td>
<td>.0340 **</td>
<td>.019</td>
</tr>
<tr>
<td>YEAR</td>
<td>.1249</td>
<td>.0204 **</td>
<td>.017</td>
</tr>
<tr>
<td>DAIB</td>
<td>-.8294</td>
<td>1.1042 **</td>
<td>.181</td>
</tr>
<tr>
<td>URM</td>
<td>-.3870</td>
<td>1.0633 **</td>
<td>.051</td>
</tr>
<tr>
<td><strong>Season:</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Multi-province, Yukon, Northwest Territories</td>
<td>--</td>
<td>--</td>
<td>--</td>
</tr>
<tr>
<td>Maritime</td>
<td>.8015</td>
<td>1.1603 **</td>
<td>.394</td>
</tr>
<tr>
<td>Quebec</td>
<td>.2895</td>
<td>1.2790 **</td>
<td>.267</td>
</tr>
<tr>
<td>Ontario</td>
<td>.2430</td>
<td>1.2674 **</td>
<td>.270</td>
</tr>
<tr>
<td>Prairie</td>
<td>.8306</td>
<td>1.1671 **</td>
<td>.372</td>
</tr>
<tr>
<td>British Columbia</td>
<td>.0605</td>
<td>1.2222 **</td>
<td>.305</td>
</tr>
<tr>
<td><strong>Quarter:</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Winter</td>
<td>-.2333</td>
<td>- .0355</td>
<td>.164</td>
</tr>
<tr>
<td>Spring</td>
<td>-.3702</td>
<td>- .0540 **</td>
<td>.152</td>
</tr>
<tr>
<td>Summer</td>
<td>--</td>
<td>--</td>
<td></td>
</tr>
<tr>
<td>Autumn</td>
<td>-.3188</td>
<td>1.0473 **</td>
<td>.154</td>
</tr>
<tr>
<td><strong>Specific unions:</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Autoworkers</td>
<td>2.2282</td>
<td>1.5006 **</td>
<td>.211</td>
</tr>
<tr>
<td>Carpenters and joiners</td>
<td>-.5624</td>
<td>- .0772</td>
<td>.500</td>
</tr>
<tr>
<td>Food and commercial workers</td>
<td>-.4191</td>
<td>- .0602</td>
<td>.304</td>
</tr>
<tr>
<td>Steelworkers</td>
<td>.5602</td>
<td>1.1063 **</td>
<td>.155</td>
</tr>
<tr>
<td>Teamsters</td>
<td>.2368</td>
<td>1.0414</td>
<td>.550</td>
</tr>
<tr>
<td>All others</td>
<td>--</td>
<td>--</td>
<td></td>
</tr>
<tr>
<td><strong>Constant</strong></td>
<td>- 12.032 **</td>
<td>--</td>
<td>1.232</td>
</tr>
</tbody>
</table>

Notes: (1) For an explanation of the terms and symbols used in this table see the notes for Table 3.3. For summary statistics of the data refer to Table 3.4.

(2) The change in the probability of a strike occurring given a unit change in each respective variable. Equation (3.11) is used to calculate the probability change for dummy variables, otherwise equation (3.9) is used.
The results from estimation of the logit model are given in Table 3.7. The second column in the table gives the change in the probability of a strike occurring associated with a unit change in each independent variable evaluated at the means of the data. Equation (3.8) is used to calculate this value for all the variables except the dummy variables, for which the expression in (3.10) is used. The variables REVENUE and C.V.REVENUE are the de-trended and de-cycled data used earlier.

The mean values of the independent variables and the gross strike rates are not repeated in Table 3.7 as the logit model is estimated over the same sample of contracts as the linear probability model with the adjusted revenue data. The only difference between the model in Table 3.7 and the linear probability model, apart from the logit specification, is the lack of industry dummy variables. The logit model would not converge to a set of coefficient estimates when the industry dummy variables were included. Lack of convergence is not, necessarily, due to a fault in the specification of the model. When a large number of independent variables are included in logit estimation, there is a possibility that nearly perfect sample separation can occur with the result that the model will not converge. For comparison, Table 3.8 includes the OLS estimation results for the model without the industry dummy variables.

First, it is instructive to compare the original OLS strike incidence estimates in Table 3.4 with the estimates in Table 3.8, which is the same model without the industry variables. Notice how the exclusion of the industry variables increases the magnitude and the statistical significance of the regional and union dummy variables. The seasonal dummy variables retain the same impact on strike incidence in the two models. The variables NEMP, LNEG, DAIB, and URM all exhibit roughly the same effect on strike incidence in the two models. Finally, the asymmetric information variables keep the same signs and remain statistically significant, although the effect of C.V.REVENUE is somewhat diminished in the model without industry dummy variables. Failure to control for industry effects reduces the impact of variability in demand at the firm level on strike incidence.

When comparing the estimates from the logit model to the OLS model in Table 3.8, remember that it is the change in probability associated with the independent variables from the logit estimates (the second column in Table 3.7) which should be compared to the OLS coefficient estimates. Thus the effect of revenue on strike incidence is very nearly the same in the two models; a unit increase in REVENUE reduces the probability of a strike by about 2.8%. In fact, this is almost the same
Table 3.8: OLS Estimates of Strike Incidence Without Industry Effects, Firm Revenue Adjusted for Trend and Cyclical Effects

<table>
<thead>
<tr>
<th>Variable</th>
<th>OLS coefficient</th>
<th>Standard error</th>
</tr>
</thead>
<tbody>
<tr>
<td>Adverse-selection variables:</td>
<td></td>
<td></td>
</tr>
<tr>
<td>REVENUE</td>
<td>- .0280 **</td>
<td>.010</td>
</tr>
<tr>
<td>C.V.REVENUE</td>
<td>.0932 *</td>
<td>.059</td>
</tr>
<tr>
<td>DTO</td>
<td>- .0228</td>
<td>.029</td>
</tr>
<tr>
<td>DCM</td>
<td>.0090</td>
<td>.021</td>
</tr>
<tr>
<td>Other contract-specific variables:</td>
<td></td>
<td></td>
</tr>
<tr>
<td>NEMP</td>
<td>.0056 *</td>
<td>.003</td>
</tr>
<tr>
<td>LNEG</td>
<td>.0310 **</td>
<td>.003</td>
</tr>
<tr>
<td>YEAR</td>
<td>.0180 **</td>
<td>.003</td>
</tr>
<tr>
<td>DAIB</td>
<td>- .1110 **</td>
<td>.020</td>
</tr>
<tr>
<td>URM</td>
<td>- .0563 **</td>
<td>.007</td>
</tr>
<tr>
<td>Region:</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Multi-province, Yukon, Northwest Territories</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Maritimes</td>
<td>.0722 *</td>
<td>.040</td>
</tr>
<tr>
<td>Quebec</td>
<td>.1331 **</td>
<td>.024</td>
</tr>
<tr>
<td>Ontario</td>
<td>.1242 **</td>
<td>.024</td>
</tr>
<tr>
<td>Prairies</td>
<td>.0815 **</td>
<td>.031</td>
</tr>
<tr>
<td>British Columbia</td>
<td>.0988 **</td>
<td>.032</td>
</tr>
<tr>
<td>Season:</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Winter</td>
<td>- .0283</td>
<td>.023</td>
</tr>
<tr>
<td>Spring</td>
<td>- .0453 **</td>
<td>.021</td>
</tr>
<tr>
<td>Summer</td>
<td>--</td>
<td>--</td>
</tr>
<tr>
<td>Autumn</td>
<td>- .0373 *</td>
<td>.022</td>
</tr>
<tr>
<td>Specific unions:</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Autoworkers</td>
<td>.4231 **</td>
<td>.041</td>
</tr>
<tr>
<td>Carpenters and joiners</td>
<td>- .0576</td>
<td>.044</td>
</tr>
<tr>
<td>Food and commercial workers</td>
<td>- .0476</td>
<td>.031</td>
</tr>
<tr>
<td>Steelworkers</td>
<td>.0811 **</td>
<td>.023</td>
</tr>
<tr>
<td>Teamsters</td>
<td>.0409</td>
<td>.100</td>
</tr>
<tr>
<td>All others</td>
<td>--</td>
<td>--</td>
</tr>
<tr>
<td>Constant</td>
<td>- 1.2596 **</td>
<td>.171</td>
</tr>
</tbody>
</table>

Notes: (1) For an explanation of the terms and symbols used in this table see the notes for Table 3.3. For summary statistics of the data refer to Table 3.4.
result as OLS estimation of the full model in Table 3.4. The effect of the coefficient of variation of firm revenue on strike incidence is greater in the logit model than in the OLS model, and the logit estimate is statistically significant at the .05 level. Also, the effect of REVENUE and of C.V.REVENUE are increased going from the OLS to the logit estimates. With regard to the remaining independent variables, the OLS and logit results indicate a reasonable degree of similarity between the two models of strike incidence. In most cases, the effects of the variables on strike incidence are estimated to be greater with the logit procedure. This suggests that using OLS in place of logit estimation does not seriously affect the results. Furthermore, given the similarity in the estimates between the OLS models with and without industry dummy variables, it is likely that logit estimation of the full model, were it possible to get convergence, would not contradict earlier results.

3.5 Summary and Conclusions

The variables suggested by the adverse-selection theory of strikes are the revenue of the firm for the contract period, and the variability of firm revenue over the period the firm is present in the contract data. In a simple model of strike incidence, firm revenue is found to have a negative effect on the likelihood of a strike, as predicted by the theory, and the effect is statistically significant. The variability in firm revenue is found to have a positive and statistically significant effect on strike incidence, also as predicted by the theoretical model of strikes. Both results are replicated using either the raw revenue data or revenue data that has been adjusted for the effect of time and the business cycle.

Since the contract data set used here is similar to that used by Gunderson et al (1986), the regression results presented in this chapter are viewed as providing support, in a conventional empirical setting, for the hypothesis that strikes serve to resolve an asymmetry of information between the firm and the union. The important difference between the data used here and the data used by Gunderson et al (1986), of course, lies in the addition of relevant firm-specific information to the contract data. Another purpose of the OLS estimation is to derive a set of explanatory variables to be used in the more complex econometric models of chapters 4 and 5.
CHAPTER IV

An Econometric Model of Wage and Strike Offers

4.1 Introduction

The endogenous variables in the adverse-selection model of collective bargaining are the wage and strike offers made by the union. However, in actual contract data the wage and strike outcomes observed depend on which of the union’s offers is chosen by the firm. The aim of this chapter is to build an econometric model of wage and strike offers recognizing that only wage and strike outcomes are observed.

In the process of constructing the econometric model, the monotonicity property (see chapter 2, Proposition 1) of the wage and strike offers plays a crucial role. In fact, as the monotonicity property in one form or another is a feature of all adverse-selection models, it is argued that the econometric model proposed here has much wider application than simply as a model of wages and strikes. Another interesting feature of the econometric model is that the occurrence of separating and pooling equilibria is endogenous to the model. Thus, for the most general specification of the model, the data are allowed to estimate which observations correspond to separating equilibria, and which to pooling equilibria.

If the occurrence of pooling equilibria is ruled out, the econometric model of wage and strike offers simplifies considerably to a Tobit-type model. Although it may be unreasonable to rule out pooling equilibria a priori in the context of collective bargaining, there may be other empirical applications of adverse-selection models where the assumption is perfectly reasonable. In such cases, the Tobit-type model proposed here applies. The advantages of this approach are that the Tobit model is familiar, the properties of the estimators are well known, and the Tobit procedure is incorporated in currently available computer programs.

To illustrate the claims made for the econometric model of wage and strike offers, the first step is to understand the link between wage and strike offers, on the one hand, and observed wage and strike outcomes, on the other. The next section of the chapter deals with the link between offers and outcomes. Section 4.3 derives the likelihood functions for the general SP Model (“SP” for Separating–Pooling) and for the S Model (“S” for Separating) where the occurrence of pooling equilibria is ruled out. Section 4.4 discusses the properties of the maximum likelihood estimators, the maximizing algorithms and testing procedures used, and the choice of explanatory variables.
4.2 Empirical Implementation of the Theory

Suppose data are available on a number of collective agreements or contracts between various firms and unions. Let \( w \) be the wage level observed in one of these contracts. Thus, \( w \) is the wage \textit{outcome} for the contract. Similarly, let \( s \) be the length of the strike observed in the same contract. Thus, \( s \) is the strike \textit{outcome} for that contract. Obviously, \( w > 0 \) and \( s \geq 0 \) for all the contracts in the data, with \( s = 0 \) in the majority of contracts as strikes occur relatively infrequently.\(^1\)

Now, consider the adverse-selection model of collective bargaining. At this point it is useful to re-introduce the union’s maximization problem as it appears in chapter 2. Problem 2 may be written thus

\[
\max_{w_1, w_2} \quad (1-q) u(w_1, \theta_1) \frac{\pi(w_2, \theta_2)}{\pi(w_1, \theta_2)} + q u(w_2, \theta_2) \quad \text{subject to } w_2 \geq w_1. \quad (4.1)
\]

Let \( \bar{w}_1 \) and \( \bar{w}_2 \) be the solution to this problem. The values of \( \bar{w}_1 \) and \( \bar{w}_2 \) are used to determine \( \bar{s}_1 \) using (2.16). Thus, \( \bar{w}_1 \) and \( \bar{w}_2 \) represent the union’s wage offers and \( \bar{s}_1 \) represents the union’s strike offer to the firm. The solution to the problem takes one of two forms depending on whether the constraint \( w_2 \geq w_1 \) is binding.

If the constraint is not binding at the solution, the union makes two wage-strike offers to the firm: a high wage offer with no strike \((\bar{w}_2, 0)\), and a low wage offer with a strike of duration \( \bar{s}_1 \), \((\bar{w}_1, \bar{s}_1)\). The two offers are said to constitute a \textit{separating equilibrium} solution to the union’s maximization problem. These offers have the property that \( \bar{w}_2 \geq \bar{w}_1 \) and \( \bar{s}_1 > 0 \), that is, the offers exhibit the monotonicity property (see chapter 2, Proposition 1). In this case, the adverse-selection model predicts that the offer \((\bar{w}_2, 0)\) will be chosen if the firm is the high demand \((\theta_2)\) type. Equivalently, if the union’s offers constitute a separating equilibrium and if the firm is type \( \theta_2 \), the offer \((\bar{w}_2, 0)\) will be the observed wage and strike \textit{outcome}. Similarly, the adverse-selection model predicts that the offer \((\bar{w}_1, \bar{s}_1)\) will be the observed \textit{outcome} if the firm is the low demand \((\theta_1)\) type.

The second solution to (4.1) occurs when the constraint is binding at the solution. When \( \bar{w}_1 = \bar{w}_2 \), the union makes a single offer to the firm comprising a wage with no strike. An offer of this nature is said to constitute a pooling equilibrium solution to the union’s maximization problem, and may be written \((\bar{w}_P, 0)\). If the solution to

\(^1\) For example, in the contract data with 2,459 observations examined in chapter 3, the overall strike frequency is .204
(4.1) is a pooling equilibrium, firm type is irrelevant since only one wage-strike offer is proposed by the union. Thus $(\bar{w}_p, 0)$ is the observed outcome.

Clearly, the separating and pooling solutions to (4.1) are mutually exclusive. Therefore, the relationship between wage and strike offers and wage and strike outcomes may be written as follows,

$$ (w, s) = \begin{cases} 
(\bar{w}_1, \bar{s}_1), & \text{if the firm is type } \theta_1; \\
(\bar{w}_2, 0), & \text{if the firm is type } \theta_2; \\
(\bar{w}_p, 0), & \text{if the union makes a pooling offer},
\end{cases} \quad (4.2) $$

where $\bar{w}_2 > \bar{w}_1$ and $\bar{s}_1 > 0$. Thus the observed wage and strike outcome for the contract $(w, s)$ is one of three possible wage and strike offers given on the right-hand side of (4.2). The wage and strike offers exhibit the monotonicity property: $\bar{w}_2 > \bar{w}_1$ and $\bar{s}_1 > 0$.

Although (4.2) highlights the role separating and pooling solutions have in determining wage and strike outcomes, it is not particularly illuminating when it comes to building an econometric model of wage and strike offers. The econometric model is based on the realization that information on the occurrence of a strike in a contract may be used to determine which of the offers on the right-hand side of (4.2) obtains. Examine (4.2) closely, and consider what information is revealed if a strike is observed in the contract under consideration. If $s > 0$, only one of the right-hand side offers in (4.2) applies, $(\bar{w}_1, \bar{s}_1)$, because neither of the other two possible offers includes a positive strike offer. Thus, according to the theoretical model, if a strike is observed in the contract, it must be that the union's offers constituted a separating equilibrium and it must be that the firm is the low demand type. Therefore, it may be concluded that $\bar{w}_1 = w$ and $\bar{s}_1 = s$. In addition, it is known from the monotonicity property that $\bar{w}_2 > \bar{w}_1 = w$. Suppose, instead, that a strike is not observed in the contract. If $s = 0$, then according to (4.2) either the union's offers constituted a separating equilibrium and the firm is type $\theta_2$, or the union made a single pooling offer to the firm. Therefore, if $s = 0$ it may be concluded that either: (a) $\bar{w}_2 = w$ and, from the monotonicity property, $\bar{w}_1 < \bar{w}_2 = w$ and $\bar{s}_1 > 0$, or; (b) $\bar{w}_p = w$.

In this manner, the occurrence or absence of a strike in each contract may be used to reveal information about the underlying structure of wage and strike offers. As this structure is important in the discussion below, it is convenient to summarize it in equation form, thus

For $s > 0$: $\bar{w}_1 = w < \bar{w}_2$ and $\bar{s}_1 = s$.  

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For $s = 0$: $\bar{w}_1 < \bar{w}_2 = w$ and $\bar{s}_1 > 0$, or;
\[
\bar{w}_P = w \quad \text{if union makes a pooling offer.}
\] (4.3)

In the adverse-selection model, whether the union’s offers constitute a separating or a pooling equilibrium is determined endogenously, given values for the exogenous variables. Thus it would be desirable if the econometric model allowed separating and pooling equilibria to be endogenously determined. In fact, the econometric model proposed here has precisely this feature. This can be shown as follows. The technique is to re-write the union’s constrained maximization problem as two separate maximization problems, the first being an unconstrained version of (4.1), the second being valid only in certain circumstances. Re-writing the union’s maximization problem in (4.1) ignoring the constraint $w_2 \geq w_1$ yields
\[
\max_{w_1, w_2} (1 - q)u(w_1, \theta_1) \frac{\pi(w_2, \theta_2)}{\pi(w_1, \theta_2)} + q u(w_2, \theta_2). \tag{4.4}
\]

Let the solution to this problem be $w_1^*$ and $w_2^*$. Provided $w_2^* > w_1^*$, there is no loss of generality in writing the union’s problem as (4.4) since the constraint in (4.1) would not be binding. Let $s_1^*$ be the strike offer associated with $w_1^*$ in the case where $w_2^* > w_1^*$. If $w_2^* \leq w_1^*$, the constraint in (4.1) would be binding and the union’s maximization must be re-written to take the constraint into account. In other words, if $w_2^* \leq w_1^*$, then (4.4) is no longer the relevant maximization problem; the relevant problem incorporates the binding constraint. Thus let $w_P = w_1 = w_2$ in (4.1), and notice that the ratio of profit functions becomes 1. In this case, the union’s maximization problem is simply
\[
\max_{w_P} (1 - q)u(w_P, \theta_1) + q u(w_P, \theta_2). \tag{4.5}
\]

Let the solution to this problem be $w_P^*$. The union is thought of as solving an unconstrained problem (4.4). Provided $w_2^* > w_1^*$, the solution to this problem is a separating equilibrium. However, if $w_2^* \leq w_1^*$, the union solves another maximization problem, (4.5), with the solution being a pooling equilibrium.

Clearly, the solution to (4.4) when $w_2^* > w_1^*$ corresponds to a separating solution to the original problem (4.1). Also, the solution to (4.5) when $w_2^* \leq w_1^*$ corresponds to a pooling solution to (4.1). In this manner, the solution to the unconstrained
maximization problem (4.4) is used to indicate the occurrence of separating and pooling equilibria. Equation (4.3) may now be written

For $s > 0$: $w_1^* = w < w_2^*$ and $s_1^* = s$.

For $s = 0$: $w_1^* < w_2^* = w$ and $s_1^* > 0$ if $w_2^* > w_1^*$, or;

$$w_P^* = w \text{ if } w_2^* \leq w_1^*.$$ 

where $s_1^*$ is the strike offer consistent with the solution $w_1^*$ and $w_2^*$ to (4.4).

To show that pooling and separating solutions are endogenously determined in the econometric model, the estimating equations must be specified. The values for $w_1^*, w_2^*, w_P^*$ and $s_1^*$ are derived from solving the first-order conditions from the maximization problem (4.4) and (4.5) given values for the exogenous variables $\theta_1$, $\theta_2$, and $q$. Owing to the complexity of the objective functions in (4.4) and (4.5), the corresponding first-order conditions are complicated non-linear equations, even if simple functional forms are chosen for the utility and profit functions (see the example in section 2.2.1). To reduce the complexity of what already is a complex model, the optimal values for $w_1^*, w_2^*, w_P^*$ and $s_1^*$ are approximated by a linear function of the exogenous variables in the neighbourhood of the solutions. Besides being straightforward, the linearity assumption is perfectly adequate for testing the comparative static predictions of the theory. Thus, let

$$w_{1t}^* = x_{1t}\beta_1 + u_{1t},$$

$$w_{2t}^* = x_{2t}\beta_2 + u_{2t},$$

$$w_{Pt}^* = x_{Pt}\beta_P + u_{Pt},$$

and

$$s_{1t}^* = x_{St}\beta_S + u_{St}.$$ 

where the subscript $t$ has been added to denote the $t$th contract observation. Here, $x_{1t}$, $x_{2t}$, $x_{Pt}$ and $x_{St}$ are $K_1$, $K_2$, $K_P$ and $K_S$ vectors of exogenous variables, respectively; $\beta_1$, $\beta_2$, $\beta_P$ and $\beta_S$ are similarly dimensioned vectors of coefficients to be estimated, and; $u_{1t}$, $u_{2t}$, $u_{Pt}$, and $u_{St}$ are disturbance terms. The disturbance terms could be thought of as resulting from errors in optimizing behaviour given that the independent variables are derived from a theoretical model. Equations (4.7) and (4.6) together define the econometric model of strike and wage offers. Now it is clear that the
occurrence of pooling and separating equilibria is endogenously determined in the econometric model, for the solutions to the union's unconstrained maximization problem (4.4), $w_2^*$ and $w_1^*$, are the endogenous variables. In terms of the econometric model, pooling equilibria are those instances where the low wage offer is estimated to be larger than the high wage offer.

At this point it is worthwhile summarizing the above discussion. Under the adverse-selection model of collective bargaining, the union is perceived as solving a maximization problem given by (4.1). The solution to this problem determines the union's wage and strike offers. The firm chooses the offer it prefers, thereby determining the wage and strike outcome observed in the data, as portrayed in (4.2). A key aspect of the adverse-selection model is that the union's wage and strike offers satisfy the monotonicity property. The property is used together with information on the incidence of a strike in each contract to identify the underlying structure of wage and strike offers, as shown in (4.3). Finally, the occurrence of pooling and separating equilibria is shown to be endogenous to the econometric model by redefining the union's constrained maximization problem (4.1) as an unconstrained problem (4.4), which depends on the wage offers $w_1^*$ and $w_2^*$, and then specifying linear estimating equations with additive disturbance terms for $w_1^*$ and $w_2^*$.

Before proceeding to the derivation of the likelihood functions used to obtain the coefficient estimates, three points are in order. First, the econometric model represented by (4.6) and (4.7) is applicable not just to the adverse-selection model of collective bargaining but to any empirical implementation of a model based on the theory of adverse-selection. This claim is based on the fact that all models of adverse-selection have two features in common: (i) a monotonicity property, and (ii) the notion of pooling and separating equilibria. Thus the underlying structure of the endogenous variables in any adverse-selection model can be represented by a model similar to (4.6).

Second, the structure of the model in (4.6) is similar to, though more complicated than, the structure of econometric models of markets in disequilibrium. Disequilibrium models are concerned with estimating demand and supply schedules when only the quantity transacted in the market is observed and where the quantity transacted is on the demand or the supply schedule, but generally not both. The estimation problem posed by the disequilibrium models is very similar to that in the present

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2 See Maddala (1986) for an overview of disequilibrium models.
model of estimating the $w^*_2$ and $w^*_p$ equations for the non-strike observations given only the wage outcome is observed. Furthermore, Maddala and Nelson (1974) point out that in disequilibrium models the occurrence of an observation on the demand or the supply curve is determined endogenously in the model. The point is also valid in the present model where the occurrence of a $w^*_2$ wage offer, i.e., the occurrence of a separating equilibrium for a non-strike observation, is endogenous. However, the econometric model of wage and strike offers is more complex than models of markets in disequilibrium because there are three regimes in the former—the strike observations, the non-strike separating solution observations, and the non-strike pooling solution observations—compared to two regimes—demand and supply—in the latter.

Lastly, suppose it is presumed that pooling equilibria are relatively rare events in collective bargaining. In other words, most unions are believed to make two wage-strike offers to the firm rather than a single pooling wage offer. Under this assumption it may be prudent to ignore the possibility of pooling equilibria altogether, thereby greatly simplifying the model to be estimated. Consider (4.3) where there are no pooling equilibria in the sample. As before, the occurrence of a strike signals a low wage outcome. But now, the absence of a strike merely indicates a high wage outcome, the additional complication introduced by pooling equilibria having been ruled out. Thus (4.6) may be rewritten

\[
\begin{align*}
\text{For } s > 0: & \quad w^*_1 = w < w^*_2 \quad \text{and } s^*_1 = s. \\
\text{For } s = 0: & \quad w^*_1 < w^*_2 = w \quad \text{and } s^*_1 > 0.
\end{align*}
\]

It is shown below in Section 4.3.1 that a model given by (4.8) and (4.7a), (4.7b), and (4.7d) can be estimated using the Tobit procedure. Clearly, any adverse-selection model where pooling equilibria are thought to be unlikely could be set up in the same way as (4.8), with the additional advantages that the techniques to estimate such a model are familiar and the properties of the estimators well known.

### 4.3 Likelihood Functions

Essentially, there are two econometric models under consideration. If pooling equilibria are assumed not to obtain, the econometric model is given by equations (4.7a), (4.7b), (4.7d), and (4.8). This model is called the S Model. In the more general SP Model, pooling equilibria are permitted. In this case, the econometric model is
given by equations (4.6) and (4.7). The present section derives the likelihood functions for each of the two models, and variants thereof, beginning with the likelihood function for the S Model in section 4.3.1. The likelihood function for the SP Model is derived in section 4.3.2.

4.3.1 Likelihood Functions for the S Model

The endogenous variables in the S Model are the low wage offer, \( w_1^* \), the high wage offer, \( w_2^* \), and the strike offer \( s^*_t \). Of course, the nature of the likelihood function depends on the hypothesized relationship between the disturbance terms \( u_1, u_2 \) and \( u_S \). To begin with, assume that the disturbance terms are jointly normally distributed with density function \( f(u_1, u_2, u_S) \), zero mean with covariance matrix \( \Sigma \), and serially independent. Thus the S Model of wage and strike offers is given by

\[
\begin{align*}
    w_1^* &= x_{1t} \beta_1 + u_{1t}, \\
    w_2^* &= x_{2t} \beta_2 + u_{2t}, \\
    s^*_t &= x_{St} \beta_S + u_{St},
\end{align*}
\]  

(4.9)

with \( (u_{1t}, u_{2t}, u_{St}) \) distributed trivariate normal with zero mean and covariance matrix \( \Sigma \) and where

\[
\Sigma = \begin{bmatrix}
\sigma_1^2 & \sigma_{12} & \sigma_{1S} \\
\sigma_{12} & \sigma_2^2 & \sigma_{2S} \\
\sigma_{1S} & \sigma_{2S} & \sigma_S^2
\end{bmatrix}
\]  

(4.10)

There are \( K_1 + K_2 + K_P + 6 \) parameters to be estimated in the model, corresponding to the coefficients in (4.9) and the elements of the covariance matrix \( \Sigma \).

According to (4.8), if \( s_t > 0 \), then \( w_1^* = w_t, w_2^* > w_t \) and \( s^*_t = s_t \). Substituting for the observed wage and strike outcome \( w_t \) and \( s_t \), equations (4.9) become

\[
\begin{align*}
    u_{1t} &= w_t - x_{1t} \beta_1, \\
    u_{2t} &= w_t - x_{2t} \beta_2, \\
    u_{St} &= s_t - x_{St} \beta_S.
\end{align*}
\]

Therefore, the density of an observation with \( s_t > 0 \) is given

\[
G_t = \int_{w_t - x_{2t} \beta_2}^{+\infty} f(w_t - x_{1t} \beta_1, u_{2t}, s_t - x_{St} \beta_S) du_{2t}.
\]  

(4.11)
Again, according to (4.8), if \( s_t = 0 \), then \( w_{1t}^* < w_t, w_{2t}^* = w_t \) and \( s_{1t}^* > 0 \). Substituting for the observed wage and (zero) strike outcome, equations (4.9) become

\[
\begin{align*}
    u_{1t} &< w_t - x_{1t}\beta_1, \\
    u_{2t} &= w_t - x_{2t}\beta_2, \\
    u_{5t} &> -x_{5t}\beta_S.
\end{align*}
\]

Thus, the density of an observation with \( s_t = 0 \) is

\[
H_t = \int_{-\infty}^{w_t - x_{1t}\beta_1} \int_{-\infty}^{u_{1t}} f(u_{1t}, w_t - x_{2t}\beta_2, u_{S_t}) du_{1t} du_{S_t}. \tag{4.12}
\]

Without loss of generality the observations are ordered so the \( T_S \) observations with a strike precede the \( T - T_S \) observations without a strike. Hence, the natural logarithm of the likelihood function for the S Model is simply

\[
\ln L = \sum_{t=1}^{T_S} \ln G_t + \sum_{t=T_S+1}^{T} \ln H_t. \tag{4.13}
\]

Owing to the double integral in \( H_t \) the first and second derivatives of the likelihood function in (4.13) are very complicated. This does not preclude estimation of the model, for methods which use numeric approximations to the derivatives in order to maximize the likelihood function may be used. However, the numeric approximation methods available are not sufficiently accurate (see section 4.4) and consequently there is a strong case for using analytic first and second derivatives. In view of the complexity of these derivatives, the only remaining option is to compromise the generality of the model.

**Independent Disturbance Terms**

Assume that \( u_1, u_2 \) and \( u_S \) are independently normally distributed. Equivalently, \( \Sigma = \text{diag}(\sigma_1^2, \sigma_2^2, \sigma_S^2) \). The number of parameters to be estimated is now \( K_1 + K_2 + K_P + 3 \), i.e., three less than for the full S Model described in (4.13). Equation (4.11) is now

\[
G_t = f_1(w_t - x_{1t}\beta_1) \int_{w_t - x_{2t}\beta_2}^{+\infty} f_2(u_{2t}) du_{2t} f_S(s_t - x_{5t}\beta_S) \tag{4.14}
\]
where

\[ f_1(w_t - x_{1t}\beta_1) = \frac{1}{\sigma_1 \sqrt{2\pi}} \exp \left[ \frac{-(w_t - x_{1t}\beta_1)^2}{2\sigma_1^2} \right], \quad (4.15a) \]

\[ f_2(u_{2t}) = \frac{1}{\sigma_2 \sqrt{2\pi}} \exp \left[ \frac{-u_{2t}^2}{2\sigma_2^2} \right], \quad (4.15b) \]

and

\[ f_S(s_t - x_{St}\beta_S) = \frac{1}{\sigma_S \sqrt{2\pi}} \exp \left[ \frac{-(s_t - x_{St}\beta_S)^2}{2\sigma_S^2} \right], \quad (4.15c) \]

Similarly, (4.12) is now

\[ H_t = \int_{-\infty}^{w_t - x_{1t}\beta_1} f_1(u_{1t}) \, du_{1t} \int_{-\infty}^{+\infty} f_2(w_t - x_{2t}\beta_2) \int_{-x_{St}\beta_S}^{+\infty} f_S(u_{St}) \, du_{St}, \quad (4.16) \]

where \( f_1(u_{1t}) \) and \( f_S(u_{St}) \) are defined analogously to (4.15b), and \( f_2(w_t - x_{2t}\beta_2) \) is defined analogously to (4.15a). Using the symmetry property of the normal density function, (4.14) becomes

\[ G_t = f_{1t} \, [1 - F_{2t}] \, f_{St} \]

where \( f_{1t} = f_1(w_t - x_{1t}\beta_1) \), \( f_{St} = f_S(s_t - x_{St}\beta_S) \) and

\[ F_{2t} = \int_{-\infty}^{(w_t - x_{2t}\beta_2)/\sigma_2} \phi(u) \, du, \]

where \( \phi \) is the standard normal density function. Similarly, (4.16) becomes

\[ H_t = F_{1t} \, f_{2t} \, [1 - F_{St}] \]

where \( f_{2t} = f_2(w_t - x_{2t}\beta_2) \),

\[ F_{1t} = \int_{-\infty}^{(w_t - x_{1t}\beta_1)/\sigma_1} \phi(u) \, du, \]

and

\[ F_{St} = \int_{-\infty}^{(-x_{St}\beta_S)/\sigma_S} \phi(u) \, du. \]

The log likelihood in (4.13) is now

\[
\ln L = \sum_{t=1}^{T_S} \ln f_{1t} + \ln[1 - F_{2t}] + \ln f_{St} \\
+ \sum_{t=T_S+1}^{T} \ln F_{1t} + \ln f_{2t} + \ln[1 - F_{St}].
\]

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This can also be written as

\[
\ln L = \sum_{t=1}^{T_S} \ln f_{1t} + \sum_{t=T_S+1}^{T} \ln F_{1t} + \sum_{t=1}^{T_S} \ln [1 - F_{2t}] + \sum_{t=T_S+1}^{T} \ln f_{2t}
\]

\[
+ \sum_{t=1}^{T_S} \ln f_{St} + \sum_{t=T_S+1}^{T} \ln [1 - F_{St}].
\]

Equivalently, \( \ln L \) may be written

\[
\ln L = \ln L_1 + \ln L_2 + \ln L_S. \tag{4.17}
\]

Since none of \( L_1, L_2 \) or \( L_S \) have any parameters in common, maximizing \( \ln L \) is equivalent to maximizing separately each of the log likelihoods on the right-hand side of (4.17). It is also apparent that each of \( \ln L_1, \ln L_2 \) and \( \ln L_S \) are similar to expressions encountered in models with limited dependent variables, otherwise known as Tobit models. Hence the algorithms used for Tobit models may be used to estimate the S Model with independent error terms.

**Partially Dependent Disturbance Terms**

Although the independence assumption is desirable from a pragmatic point of view, it would be more satisfying if it could be relaxed. One possible generalization is to allow the disturbance terms in the wage offer equations, \( u_1 \) and \( u_2 \), to be contemporaneously correlated maintaining the independence of \( u_1 \) and \( u_2 \) from \( u_S \). The likelihood function under partially dependent disturbance terms will separate into two components—one for the wage offers, and one for the strike offer—in much the same way as in (4.17).

The covariance matrix is now

\[
\Sigma = \begin{bmatrix}
\sigma_1^2 & \sigma_{12} & 0 \\
\sigma_{12} & \sigma_2^2 & 0 \\
0 & 0 & \sigma_S^2
\end{bmatrix}
\]

If \( g \) is the joint density for \( u_1 \) and \( u_2 \), (4.11) becomes

\[
G_t = \int_{w_t-x_{2t}\beta_2}^{+\infty} g(w_t - x_{1t}\beta_1, u_{2t}) \, du_{2t} \, f_S(s_t - x_{St}\beta_S)
\]
while (4.12) is now given
\[
H_t = \int_{-\infty}^{w_t-x_1t} g(u_{1t}, w_t - x_2t) du_{1t} \int_{-\infty}^{+\infty} f_S(u_{St}) du_{St}.
\]
The log likelihood is thus
\[
\ln L = \sum_{t=1}^{T_S} \ln \hat{g}_1(w_t - x_{1t} \beta_1) + \ln \int_{w_t-x_2t}^{+\infty} \hat{g}_2(u_{2t}|w_t - x_{1t} \beta_1) du_{2t} + \ln L_S.
\]
where \( g \) has been factored into marginal densities, denoted by bars, and conditional densities, denoted by hats. Notice the appearance of the \( \ln L_S \) term, defined in (4.17). Thus \( \ln L_S \) can be maximized separately as before. The likelihood function in (4.18), ignoring the \( \ln L_S \) term, proved difficult to estimate for reasons explained below (see section 5.2.2). In view of the difficulties encountered, alternative generalizations of the independence assumption are not considered.

4.3.2 An Alternative Model of Strike Duration

Given the assumption of independent disturbance terms, the strike offer model can be estimated separately from the wage offer equations simply by maximizing \( \ln L_S \), as discussed below (4.17). It is worthwhile pondering exactly what behaviour this model is trying to capture. Remember that, for the time being, pooling equilibria have been ruled out. In this case, according to the adverse-selection model of collective bargaining, a positive strike offer is made for every contract in the data. Of course, a positive strike outcome is observed for only some contracts because strikes occur only in type \( \theta_1 \) firms.

Contrast this view of strikes with an alternative Tobit model of strikes. Suppose there exists a variable \( \tilde{s}_t \) which measures "desired" or "potential" strike length. Let \( \tilde{s}_t = \tilde{x}_{St} \beta_S + \tilde{u}_{St} \), where \( \tilde{u}_{St} \) has a zero mean and constant variance. Only observations for which desired strike length is positive are observed as a strike in the data. For those values of \( \tilde{s}_t < 0 \), we observe only the absence of a strike. This gives rise to the Tobit model
\[
s_t = \begin{cases} 
\tilde{s}_t, & \text{if } \tilde{s}_t > 0; \\
0, & \text{otherwise.}
\end{cases}
\]

Using the same notation as above, the likelihood function for this model [Maddala (1983), p.152] is

\[
\ln \tilde{L} = \sum_{t=1}^{T_S} \ln f_{St} + \sum_{t=T_S+1}^{T} \ln F_{St}.
\] (4.20)

The likelihood function for the S Model of strike offers is

\[
\ln L = \sum_{t=1}^{T_S} \ln f_{St} + \sum_{t=T_S+1}^{T} \ln[1 - F_{St}].
\]

which differs from \( \ln \tilde{L} \) over the non-strike observations. The difference is not surprising. The alternative Tobit model regards non-strike observations as cases where the union was unwilling to strike. The S model of strike offers regards non-strike observations as cases where the union was willing to strike. In the S Model, the absence of a strike indicates the presence of a firm type which was incompatible with a positive strike outcome.

In fact, the role played by firm type in the occurrence of a strike has a very interesting implication for the S Model of strike offers. Consider again the conventional Tobit model of strikes. It is well known that the OLS estimator of \( \tilde{\beta}_S \) from a regression using only positive observations for \( s_t \) is biased and inconsistent.\(^3\) These properties arise essentially because a positive desired strike length causes a positive strike outcome. In the S Model, on the other hand, the strike offer is always positive—a positive strike outcome occurs only when the firm is a particular type and the firm's type is determined outside the model. Write the S Model in an analogous fashion to the conventional Tobit model of strikes shown in (4.19), thus

\[
s_t = s^{*}_{1t}, \quad \text{if firm is type } \theta_1; \\
0 < s^{*}_{it}, \quad \text{otherwise.}
\] (4.21)

In (4.19), whether a strike is observed depends on \( \tilde{s}_t \), the endogenous variable. In (4.21), whether a strike is observed depends on firm type, which is exogenous. The major consequence of this fact is that the OLS estimator of \( \tilde{\beta}_S \) in the strike offer equation using only positive observations of \( s_t \) has all the usual desirable properties of a least squares estimator in a regression with well behaved disturbances. Of course, this OLS estimator is not fully efficient because it ignores the information available from the non-strike subsample of the data.

\(^3\) See, for example, Amemiya (1985), p.367.
4.3.3 Likelihood Functions for the SP Model

Intuitively it seems more plausible, at least in a model of collective bargaining, that pooling equilibria will obtain in some of the observed data. The intuition is based on the casual observation that there are cases where it is simply not in the union's interest to accept the risk of a strike. In the adverse-selection model of firm and union bargaining, such cases occur when the union makes a single wage offer with no associated strike, i.e., a pooling equilibrium obtains as the solution to the union's maximization problem. Widespread evidence of pooling equilibria would explain, at least in part, the relative infrequency of strikes. It is also possible that the occurrence of pooling equilibria may enable us to explain other stylized facts about the behaviour of strikes, such as procyclical strike frequency, for example.

An econometric model of wage and strike offers which allows for the occurrence of pooling equilibria is described by equations (4.6) and (4.7). Owing to the fact that pooling is endogenous to the model, the likelihood function for the SP Model is considerably more complicated than the likelihood function for the most general S model given by (4.13). The increased complexity of the SP Model, and the lack of success with dependent disturbance terms for the more simple S Model, means that any generalization beyond independence of all the error terms in the SP Model is probably out of reach. Therefore, the SP Model is described as follows

\[
\begin{align*}
\hat{w}_{1t}^* &= x_{1t}\beta_1 + u_{1t}, \\
\hat{w}_{2t}^* &= x_{2t}\beta_2 + u_{2t}, \\
\hat{w}_{Pt}^* &= x_{Pt}\beta_P + u_{Pt}, \\
& \text{and} \\
\hat{s}_{1t}^* &= x_{St}\beta_S + u_{St},
\end{align*}
\]

with \((u_1, u_2, u_P, u_S)\) independently normally distributed with zero mean and covariance matrix \(\Sigma = \text{diag}(\sigma_1^2, \sigma_2^2, \sigma_P^2, \sigma_S^2)\). The model has \(K_1 + K_2 + K_P + K_S + 4\) parameters to be estimated, corresponding to the coefficients in each of the wage and strike offer equations and the four variances.

Using the same procedure as for the S Model, the likelihood function for the SP Model is assembled from two components—one for the observations with a strike, the other for observations without a strike. Thus, if \(s_t > 0\), it follows from (4.6)
that \( w_t = w_{1t}^*, \ w_{2t}^* > w_t \) and \( s_{1t}^* = s_t \). These conditions are exactly the same as for the strike observations in the S Model (see the discussion following (4.10)). Thus the portion of the likelihood function for the strike observations in the SP Model is identical to the corresponding portion of the likelihood for the S Model. As strikes are the result of separating equilibria, allowing pooling equilibria in the SP Model will not affect the strike observations, so it is not surprising that the two likelihood functions over the strike observations are the same. The density for a representative observation with a strike is given from (4.14) as

\[
G_t = f_{1t}[1 - F_{2t}]f_{St}.
\]

(4.22)

It is the likelihood for the non-strike observations which is affected by the assumption of pooling equilibria. The adverse-selection model predicts that an outcome of no strike is due either to a separating offer and the occurrence of a type \( \theta_2 \) firm, or to a pooling offer. From (4.6) the probability of a separating offer is simply given by \( \Pr(w_{1t}^* > w_{2t}^*) \). As stated in section 4.2 this is endogenously determined in the econometric model, and it can now be seen that the endogeneity is due to the stochastic specification in the \( w_2^* \) and \( w_1^* \) wage offer equations. Define \( \lambda_t \) as

\[
\lambda_t = \Pr\left( w_{1t}^* > w_{2t}^* \right) \\
= \Pr\left( x_{2t}\beta_2 + u_{2t} > x_{1t}\beta_1 + u_{1t} \right) \\
= \Pr\left( u_{2t} - u_{1t} > x_{1t}\beta_1 - x_{2t}\beta_2 \right).
\]

Thus \( \lambda_t \) is simply the probability of a separating offer in contract \( t \). If \( s_t = 0 \) and if the observation corresponds to a separating offer, it is known that \( w_{1t}^* = w_t, \ w_{1t}^* < w_t, \) and \( s_{1t}^* > 0 \). Hence

\[
h(w_t, s_t | \text{separating offer}) = \frac{F_{1t}f_{2t}[1 - F_{St}]}{\lambda_t}.
\]

(4.23)

If \( s_t = 0 \) and if the observation corresponds to a pooling offer, it is known that \( w_{Pt}^* = w_t \) and \( w_{1t}^* \geq w_{2t}^* \). Hence

\[
h(w_t, s_t | \text{pooling offer}) = \frac{f_{Pt}F_{vt}}{1 - \lambda_t},
\]

(4.24)

where

\[
F_{vt} = \int_{-\infty}^{v_t} \phi(u) \, du
\]
and

\[ f_{P_t} = \frac{1}{\sigma_P \sqrt{2\pi}} \exp \left[ -\frac{(w_t - x_{P_t} \beta_P)^2}{2\sigma_P^2} \right], \]

for \( v_t = (x_{1t} \beta_1 - x_{2t} \beta_2)/\sigma \) and \( \sigma^2 = \sigma_1^2 + \sigma_2^2 \) by virtue of the assumption that \( u_1 \) and \( u_2 \) are independent. The unconditional density of \( w_t \) and \( s_t \) is therefore

\[ H_t = \lambda_t h(w_t, s_t | \text{separating offer}) + (1 - \lambda_t) h(w_t, s_t | \text{pooling offer}) \]

By substitution from (4.23) and (4.24) \( H_t \) becomes

\[ H_t = F_{1t} f_{2t} [1 - F_{St}] + f_{P_t} F_{vt}. \] (4.25)

The log likelihood is given by substituting \( G_t \) and \( H_t \) from (4.22) and (4.25), respectively, into (4.13). Owing to the large number of parameters to be estimated in the SP Model, a version which examines only the wage offers is considered. Ignoring the strike offer equation, (4.22) becomes

\[ G_t = f_{1t} [1 - F_{2t}], \] (4.26)

and (4.25) becomes

\[ H_t = F_{1t} f_{2t} + f_{P_t} F_{vt}. \] (4.27)

The log likelihood for this model is found by substituting (4.26) and (4.27) into (4.13).

4.3.4 Unboundedness of the Likelihood Function for the SP Model

As stated in Section 3.2, the likelihood function of the SP Model is similar to likelihood functions encountered in models of markets in disequilibrium. Unfortunately, this similarity extends to an unboundedness property commonly found in likelihood functions for disequilibrium models where the sample separation is unknown. To illustrate that the likelihood function for the SP Model is unbounded for certain parameter values, consider the likelihood function for the model of wages described in (4.26) and (4.27). The argument may be extended easily to the more general model including strike offers. Substituting (4.26) and (4.27) into (4.13) we have

\[ \ln L = \sum_{t=1}^{T_s} \ln f_{1t} [1 - F_{2t}] + \sum_{t=T_s+1}^{T} \ln [F_{1t} f_{2t} + f_{P_t} F_{vt}] \] (4.28)

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Suppose \( w_t = x_{Pt}\beta_p \) for some \( t \) where \( t \geq T_s + 1 \). Hence \( f_{Pt} = 1/(\sigma P \sqrt{2\pi}) \) and \( f_{Pt} \to \infty \) as \( \sigma_p \to 0 \). Furthermore \( f_{Pt'} \to 0 \) as \( \sigma_p \to 0 \) for all observations \( t' \geq T_s + 1 \) where \( t' \neq t \). Suppose \( \sigma_1^2 \neq 0 \) and \( \sigma_2^2 \neq 0 \), and that both are finite. Hence \( \sigma^2 = \sigma_1^2 + \sigma_2^2 \) is finite as well. Consequently, \( f_{1t}, f_{2t}, F_{1t}, F_{2t} \) and \( F_{vt} \) are all finite. Therefore the likelihood is unbounded as \( \sigma_p \to 0 \) if \( w_t = x_{Pt}\beta_p \) for any non-strike observation.

The unboundedness property of the likelihood function in the SP Model is manifested in the portion of the function defined over the non-strike observations, i.e., for \( t = T_s + 1, \ldots, T \) in (4.28). For these observations, it is not known whether the observed wage is the result of a high wage offer or a pooling wage offer. In other words, the sample separation of the wage offers is unknown for \( t = T_s + 1, \ldots, T \). Unknown sample separation is also encountered in econometric models of markets in disequilibrium, and disequilibrium models also suffer from the unboundedness problem.

The estimation technique undertaken in the literature concerned with disequilibrium models is to find a local maximum of the likelihood function. The procedure is justified, essentially, by arguing that the global maximum is not relevant for determining the true parameter values. For example, Kiefer (1978) shows that a root of the likelihood function corresponding to a local maximum is consistent, asymptotically normal and efficient for a likelihood function in a disequilibrium model. Amemiya and Sen (1977) show, for the case of an unbounded likelihood function, that a consistent estimator of the true parameter value corresponds to a local rather than global maximum of the likelihood function. Thus the estimation procedure for the SP Model is to find a local maximum of the likelihood function.

4.4 Maximizing Algorithms, Testing Procedures and the Data

For the S Model with completely independent error terms, the likelihood functions are similar to those found in limited dependent variable Tobit models. Therefore, estimates of the wage and strike offer equation coefficients and the variances of the respective disturbance terms have identical properties to Tobit estimates. Thus the parameter estimates of the S Model are consistent and asymptotically normal. As all the parameter estimates are maximum likelihood, hypotheses may be tested using a likelihood ratio test. If \( \ln L^R \) is the restricted and \( \ln L^U \) the unrestricted value of the log likelihood function respectively, then the likelihood ratio test statistic,
\[-2[\ln L^R - \ln L^U],\] is distributed asymptotically Chi-square with degrees of freedom equal to the number of independent restrictions imposed in the restricted model. The asymptotic covariance matrix for the parameter estimates is derived in the usual way as the negative of the inverse of the Hessian of the relevant log likelihood function.

Recent evidence furnished by Joe Ritter at the University of Michigan Computing Center is critical of numeric approximations to the derivatives of non-linear functions. For example, he reports that the method of centered differences can result in approximations to the first derivatives which have the wrong sign, are several orders of magnitude different, or both, from the correct values. In addition, his experiments have revealed that the approximated Hessian is not symmetric. In view of these difficulties, and especially since the Hessian of the likelihood function is used to calculate the asymptotic standard errors of the parameter estimates, it is decided to use analytic first and second derivatives in the algorithms used to maximize the likelihood functions of the S and SP Models. The analytic derivatives for the SP model are provided in the Appendix.

Two algorithms are used to maximize the various likelihood functions: a quasi-Newton method and the quadratic hill-climbing algorithm of Goldfeld, Quandt, and Trotter (1966). Since the log likelihood function for the Tobit model is globally concave, if an algorithm converges in the S Model, it converges to a global maximum. As a check of the computer programs used here some of the Tobit estimation is carried out using the econometric program SHAZAM. In the SP Model, convergence to a local maximum is deemed to have been obtained if all the following criteria are met: (i) the largest absolute value of the gradient vector of the log likelihood function with respect to the parameters is \(1.0 \times 10^{-5}\); (ii) the Hessian is negative definite, and (iii)

\[
\begin{bmatrix}
\frac{\partial \ln L}{\partial \gamma}_{\gamma^*} \\
\frac{\partial^2 \ln L}{\partial \gamma \partial \gamma^T}_{\gamma^*}
\end{bmatrix}^{-1} \begin{bmatrix}
\frac{\partial \ln L}{\partial \gamma}_{\gamma^*}
\end{bmatrix} \leq 1.0 \times 10^{-5},
\]

where \(\gamma\) is the vector of parameters in the likelihood function \(\ln L\) and \(\gamma^*\) is the vector of parameter estimates.

There are two data sets used in the estimation of the S Model and the SP Model. When estimation involves only the wage as the dependent variable, the Full Sample

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4 I am grateful to Nadine Hoffman of the University of British Columbia Computing Centre for informing me of Joe Ritter's work. Mr. Ritter's comments are available from the author.

with 2,459 contracts is used. When the strike length is included as a dependent variable, estimation is confined to the Manufacturing Subsample with 1,516 contracts.

Before proceeding to the estimation results, the choice of independent variables in the wage and strike offer equations must be discussed. These independent variables potentially serve two purposes: (i) to test the comparative static predictions of the adverse-selection model; (ii) to control for unobserved heterogeneity which might otherwise confound the parameter estimates.

The exogenous variables in the adverse-selection model of collective bargaining are $\theta_1$ and $\theta_2$, which reflect the level of demand at the firm, and $q$, the union’s subjective probability the firm is type $\theta_2$. As discussed in Chapter 3, the empirical proxy chosen for the firm’s level of demand is the annual revenue of the firm for the fiscal year starting at least 6 months after the settlement date of the contract. Firm revenue is used as an independent variable in the wage and strike offer equations. The coefficient of variation of firm revenue measured over the period the firm is present in the data is also used as an independent variable. The variable is intended to test for the impact predicted by the theory that variability in firm type has on the dependent variables.

Referring to Table 2.1 in chapter 2, the comparative static predictions of the adverse-selection model state that low (high) wage offers are positively (negatively) related to the level of demand at the low demand firm and negatively (positively) related to the level of demand at the high demand firm type. Thus, wage offers are predicted to be positively related to own-firm type and negatively related to cross-firm type. Consider the predictions from the theory of the effects of the revenue and coefficient of variation of revenue variables on the wage and strike offers in the S Model. As pooling equilibria are ruled out, the occurrence of a strike indicates the presence of a low demand firm type and the absence of a strike indicates the presence of a high demand firm type. Therefore, over the observations with a strike, the low wage offer is predicted to be positively related to firm revenue due to the own-type effect. Over the observations without a strike, the high wage offer is predicted to be positively related to firm revenue due to the own-type effect. However, because of the cross-type effect, the sign of the effect of firm revenue on both the low and high wage offers cannot be predicted when estimation is over the full sample, though the effect of revenue should be reduced. The same argument applies to predictions for the strike offer equation; over the strike observations, strike offers and firm revenue

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are predicted to be negatively related, and over the full sample the magnitude of this relationship will be reduced and may become positive.

With respect to the variable which measures variability in firm type, the coefficient of variation of firm revenue, the predictions of the theory are much less complicated. Firms with greater variability in revenue are predicted to have lower low wage offers, higher high wage offers, and longer strikes. Another comparative static property of the theory is that the high wage offer depends positively on the wage available to union members if they quit the firm. The alternative wage for each contract is proxied by the average real hourly wage in manufacturing in the month the contract is settled. These data are assembled as follows. First, the average hourly earnings of hourly rated wage-earners in manufacturing in nominal dollars are collected for each month from 1964–86. For the period January, 1964 to February, 1982 the source of the wage data is the CANSIM Main Base where the series number is D1518. For the period March, 1982 to February, 1983 the data are from Statistics Canada, Catalogue 72–002, (March 1983, Table 18, p.108). For the period from March, 1983 onwards, the data are from the CANSIM University Base (1st Quarter, 1988) where the series number is L5607. The last group of wage rates is multiplied by 1.04035, to reflect revisions in the establishment survey, as reported in Card (1987). The manufacturing wage data are then normalized by the Consumer Price Index (1971=1.0) which is discussed in chapter 3.

For the SP Model, the possibility that there are pooling equilibria in the data requires an extra set of predictions for the pooling wage equation. The theory predicts that the pooling wage is positively related to an increase in the level of demand for both firm types. Therefore, the pooling wage is expected to be positively related to firm revenue. Recall from the predictions that the variability in firm revenue variable has no relevance in the context of the pooling wage. Thus the coefficient of variation of firm revenue is not included as an independent variable in the pooling wage equation.

As noted in Chapter 3, the incidence of strikes tends to be: (i), procyclical; (ii), positively related to the size of the bargaining unit; (iii), positively related to the length of contract negotiations; (iv), negatively related to the imposition of wage and price guidelines in 1975–78, and; (v), increasing over the sample period. To control for these factors, the following independent variables are included in the estimation of the wage and strike offer equations; the unemployment rate of males aged 25 years and older (i.e., the variable URM discussed in chapter 3), bargaining unit size (NEMP),
duration of contract negotiations (LNEG), a dummy variable which takes the value of unity during the period the Anti-Inflation Board was in existence (DAIB), and a trend term (YEAR).

The third set of independent variables is intended to control for heterogeneity across unions, regions, and industries. As in Chapter 3 heterogeneity is captured by a simple fixed effect, i.e. including a dummy variable for the union, region or industry which is expected to behave differently from the norm. All the above variables are discussed in detail in section 3.3 in chapter 3. Descriptive statistics for the data are presented in Table 3.1 of chapter 3.
CHAPTER V

Empirical Results from the Econometric Model of Wage and Strike Offers

5.1 Introduction

Broadly speaking, there are two econometric models under consideration. The S Model presumes that the observed data are generated by wage and strike offers which, at every observation, constitute a separating offer by the union to the firm. Equivalently, pooling offers are ruled out for every observed data point in the S Model. While this may or may not be a plausible hypothesis in the context of collective bargaining, it nevertheless serves to illustrate an estimation procedure for adverse-selection models where the assumption is believed, for theoretical reasons or otherwise, to be plausible. In any case, the S Model is a natural point of departure and provides the base case for the more complex SP model.

The SP Model allows for the possibility that some of the non-strike observations may have been the outcome of pooling offers. The model is designed so that the data estimate which observations correspond to separating offers, and which to pooling offers. Intuitively, it seems reasonable to expect that a significant proportion of the non-strike observations are the outcome of pooling offers. Casual empiricism suggests that there are instances in which, at least in terms of economics, it is not worth the union’s while to accept the possibility of a strike.

For the S Model there are two cases to consider. First, under the assumption of independent disturbance terms, the log likelihood function dissolves into three component parts—one for $w_2^*$, the high wage offer, one for $w_1^*$, the low wage offer, and one for $s_1^*$, the strike offer—each of which is maximized separately using a procedure similar to Tobit estimation. These results are discussed in 5.2.1 below. Next, the disturbance terms in the $w_1^*$ and $w_2^*$ equations are allowed to be contemporaneously correlated and independent from the disturbance term in the $s_1^*$ equation. Results from the estimation of this model are presented in section 5.2.2 below.

For the SP Model the results are presented in section 5.3. Section 5.3.1 contains the results from the estimation of the wage offer equations only, while section 5.3.2 discusses the results from joint estimation of the wage and strike offer equations. Section 5.4 considers the results from all of the estimated models and reconciles them with previous empirical work on the topic of wages and strikes.

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5.2 Results from the S Model

Before proceeding to the results from estimation of the wage and strike offer equations, consider a least squares regression of the natural logarithm of the (real) wage observed in each contract on the complete set of independent variables considered in the study. The regression serves as a summary description of the wage outcome data and also highlights some of the characteristics of the data. The results are shown in Table 5.1 for the Full Sample and the Manufacturing Subsample with and without industry effects. As the dependent variable is in logarithmic form, estimated coefficients are interpreted as the proportionate change in the observed real wage given a unit change in the independent variable.

The results for estimation over the Full Sample with industry dummy variables are discussed first. As expected, actual firm type is positively related to the wage outcome, as indicated by the coefficient on the revenue variable, and the effect is statistically significant at the .05 level. The predicted coefficient sign for the coefficient of variation of firm revenue is ambiguous, though the measured effect is negative and significant at the .05 level. With respect to other contract specific variables, wages are approximately 0.55% higher per additional thousand workers in the bargaining unit and 0.9% higher for each additional month of contract negotiations, and both effects are statistically significant at the .05 level. Each one (1971) dollar rise in the average manufacturing wage increases observed wages by 35%. Given that the mean contract wage is $3.58, the coefficient on MANW suggests an almost unit elastic relationship between the sample of wages in the data and the average manufacturing wage. The Anti-Inflation Board is estimated to have reduced contract wages by almost 11%, and the effect is significant. Wage settlements vary countercyclically; with each 1% rise in the unemployment rate of prime-age males, wages fall by 2.1%. Finally, over the sample period, wages are not estimated to have increased at a significant rate.1

Relative to the omitted region—multi-province agreements and contracts in the Yukon and Northwest Territories—wages are seen to be 17% lower in the Maritimes, 5.2% lower in Quebec, 3.4% lower in Ontario, 1.6% lower in the Prairies, and 12% higher in British Columbia, with the effects measured to be significant in the Maritimes, Quebec, Ontario and B.C. Curiously, contracts settled in the winter and spring are found to yield wages which are significantly lower, by 3.6% and 2.5% respectively,

1 In an unreported regression without the MANW variable, the trend term is statistically significant and its coefficient indicates that wages have increased at the rate of 3.6% per year.
Table 5.1: Coefficient Estimates of the Wage Outcome(1)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Full Sample</th>
<th>Manufacturing Subsample</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Industry Effects Included</td>
<td>Industry Effects Excluded</td>
</tr>
<tr>
<td></td>
<td>Excluded</td>
<td>Included</td>
</tr>
<tr>
<td></td>
<td>Excluded</td>
<td>Included</td>
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<tr>
<td>Constant</td>
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<td>-0.402 *</td>
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<tr>
<td></td>
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<td>(0.15)</td>
</tr>
<tr>
<td>NEMP</td>
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<td>0.00757 *</td>
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<tr>
<td></td>
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<td>(0.0019)</td>
</tr>
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<td>0.00715 *</td>
</tr>
<tr>
<td></td>
<td>(0.0014)</td>
<td>(0.0015)</td>
</tr>
<tr>
<td>MANW</td>
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<td></td>
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<td>(0.040)</td>
</tr>
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</tr>
<tr>
<td></td>
<td>(0.0036)</td>
<td>(0.0041)</td>
</tr>
<tr>
<td>YEAR</td>
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</tr>
<tr>
<td></td>
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<td>(0.0037)</td>
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<tr>
<td>DAIB</td>
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<tr>
<td></td>
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<td>(0.0016)</td>
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<tr>
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<td>(0.010)</td>
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</tr>
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<tr>
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<td>(0.023)</td>
<td>(0.033)</td>
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<tr>
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<td>(0.018)</td>
</tr>
<tr>
<td>Ontario</td>
<td>-0.0340 *</td>
<td>-0.0899 *</td>
</tr>
<tr>
<td></td>
<td>(0.016)</td>
<td>(0.017)</td>
</tr>
<tr>
<td>Prairies</td>
<td>-0.0163</td>
<td>-0.104 *</td>
</tr>
<tr>
<td></td>
<td>(0.019)</td>
<td>(0.021)</td>
</tr>
<tr>
<td>B.C.</td>
<td>0.15 *</td>
<td>0.0551 *</td>
</tr>
<tr>
<td></td>
<td>(0.015)</td>
<td>(0.024)</td>
</tr>
<tr>
<td>Autoworkers</td>
<td>0.125 *</td>
<td>0.126 *</td>
</tr>
<tr>
<td></td>
<td>(0.017)</td>
<td>(0.018)</td>
</tr>
<tr>
<td>Carpenters and</td>
<td>0.152 *</td>
<td>--</td>
</tr>
<tr>
<td>Joiners</td>
<td>(0.027)</td>
<td>--</td>
</tr>
<tr>
<td>Food and Commercial</td>
<td>-0.0144</td>
<td>-0.107 *</td>
</tr>
<tr>
<td></td>
<td>(0.024)</td>
<td>(0.038)</td>
</tr>
<tr>
<td>Workers</td>
<td>0.00668</td>
<td>0.00886 *</td>
</tr>
<tr>
<td></td>
<td>(0.010)</td>
<td>(0.011)</td>
</tr>
<tr>
<td>Steelworkers</td>
<td>0.133 *</td>
<td>0.133 *</td>
</tr>
<tr>
<td></td>
<td>(0.026)</td>
<td>(0.030)</td>
</tr>
<tr>
<td>Teamsters</td>
<td>0.0492</td>
<td>0.0393 *</td>
</tr>
<tr>
<td></td>
<td>(0.0060)</td>
<td>(0.0056)</td>
</tr>
<tr>
<td>REVENUE</td>
<td>0.0436</td>
<td>0.0875 *</td>
</tr>
<tr>
<td></td>
<td>(0.018)</td>
<td>(0.024)</td>
</tr>
</tbody>
</table>

lnL     885.613  224.409  780.557  368.342
R-squared .6719   .4383  .7194  .5166
S.E.E.    .1706   .2219  .1465  .1911

Notes: (1) Standard errors are given in parentheses. * signifies statistical significance at the .05 level.
(2) There are 2,459 observations in the Full Sample and 1,516 observations in the Manufacturing Subsample.
than wages for contracts settled during the summer months. There seems to be no rational explanation for this result. It is a particularly odd result given that there are separate controls for regions, unions and industries in the estimated equation. Relative to all other unions, wages are estimated to be 13% larger in the U.A.W., 15% larger for members of the United Brotherhood of Carpenters and Joiners, and 13% higher for Teamsters members, while the effects measured for the Food and Commercial Workers and the U.S.W. are statistically insignificant. Lastly, the industry fixed effects, which are not reported in Table 5.1, show almost all industries to have significantly higher wages than the omitted group (miscellaneous manufacturing, clothing and knitting industries). This is probably due to a combination of genuine differences in wages across industries and to differences in the base rate occupation between industries. Which factor is more important cannot be determined without information on the base rate occupation in each contract, and this information is not available on the MCA tape.

The logarithm of the observed wage equation for the Full Sample is also estimated without industry fixed effects, as reported in the second column of Table 5.1. The main consequence of ignoring industry effects is an increase in the impact of the region and union dummy variables. This is not at all surprising given that certain industries are located in specific regions of the country and bargain with particular unions. For example, the automobile industry, which traditionally has high wages, is located mostly in Ontario and many of its workers are members of the U.A.W. Notice that the effect of bargaining unit size on wages switches from being positive and significant for the equation with industry effects to being negative and significant in the equation without industry effects. Finally, the coefficient of variation of firm revenue is estimated to be significantly negatively related to the wage in the equation without industry effects.

The estimates for the Manufacturing Subsample with and without industry fixed effects are very similar to the corresponding estimates for the Full Sample. The effect of revenue on the real wage is estimated to be positive and significant, as predicted by the theory. The effect of variability in firm demand is negative and significant at the .05 level, and the coefficient is almost twice as large for the estimates from the Full

---

2 Wage levels in collective agreements are typically defined relative to the lowest paid or base rate occupation, e.g., janitors, in the bargaining unit.

3 For some reason, earlier studies using the MCA tape, e.g. Swidinsky and Vanderkamp (1982), do have information on the base rate occupation.
Sample. One important difference between the full sample and the subsample is that the seasonal dummy variables do not retain their significance in the Manufacturing Subsample estimates. Thus there seems to be some interaction between the time of year and contracts settled in the non-manufacturing sector. Summarizing the results from the Manufacturing Subsample: the real contract wage is procyclical, it depends positively on bargaining unit size and the length of negotiations. Again, the coefficient on MANW suggests a unit-elastic relationship between the real manufacturing wage and the level of observed real wages.

5.2.1 S Model Estimates—Independent Disturbance Terms

As pointed out in Section 4.3.1, maximizing the likelihood function of the S Model when the disturbance terms in the offer equations are assumed to be mutually independent is equivalent to maximizing a Tobit-type likelihood function separately for each offer equation. In addition, OLS estimation of the offer equations over the relevant subsample of data yields coefficient estimates which have the standard properties but are not fully efficient. Table 5.2 reports OLS and Tobit estimates from the low and high wage offer equations.

Consider OLS estimation of the low wage offer equation over the 501 observations which have a strike. As expected, the coefficient on the revenue variable is positive and significant at the .05 level, reflecting that the low wage offer is greater for firms with a higher level of demand. Contrary to expectation, greater variability in firm revenue is estimated to increase the low wage offer. However, the effect is not statistically significant. Low wage offers are independent of bargaining unit size, and the length of negotiations. The low wage offer exhibits a significant procyclical trend by declining 2% for every 1% increase in the unemployment rate, while the Anti-Inflation Board appears to have effected a significant 16% reduction in low wage offers. Relative to contracts settled in the summer months, low wage offers for contracts settled in the three other seasons are lower, and only the effect for the autumn is not statistically significant at .05 level. Finally, B.C. is the only region to exhibit significantly different low wage offers relative to the control group; low wage offers in B.C. appear to be 14% higher than low wage offers for multi-province, Yukon and Northwest Territories contracts.

Results from Tobit estimation of the low wage offer equation over the full sample of 2,459 observations appear in the second column of Table 5.2. The Tobit coefficients
Table 5.2: S Model - Wage Offer Equation Coefficient Estimates(1)

<table>
<thead>
<tr>
<th>Procedure</th>
<th>Low wage offer equation</th>
<th>High wage offer equation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Variable</td>
<td>OLS</td>
<td>TOBIT</td>
</tr>
<tr>
<td>Constant</td>
<td>0.231</td>
<td>-1.76 *</td>
</tr>
<tr>
<td></td>
<td>(0.27)</td>
<td>(0.36)</td>
</tr>
<tr>
<td>NEMP</td>
<td>-0.0069</td>
<td>0.0324</td>
</tr>
<tr>
<td></td>
<td>(0.0033)</td>
<td>(0.0044)</td>
</tr>
<tr>
<td>LNEG</td>
<td>0.00446</td>
<td>0.0428 *</td>
</tr>
<tr>
<td></td>
<td>(0.0028)</td>
<td>(0.0034)</td>
</tr>
<tr>
<td>MANW</td>
<td>0.548 *</td>
<td>0.503 *</td>
</tr>
<tr>
<td></td>
<td>(0.079)</td>
<td>(0.10)</td>
</tr>
<tr>
<td>URM</td>
<td>-0.0203 *</td>
<td>-0.0725 *</td>
</tr>
<tr>
<td></td>
<td>(0.0081)</td>
<td>(0.010)</td>
</tr>
<tr>
<td>YEAR</td>
<td>-0.0103</td>
<td>0.0108</td>
</tr>
<tr>
<td></td>
<td>(0.0071)</td>
<td>(0.0092)</td>
</tr>
<tr>
<td>DAIB</td>
<td>-0.162 *</td>
<td>-0.295 *</td>
</tr>
<tr>
<td></td>
<td>(0.034)</td>
<td>(0.042)</td>
</tr>
<tr>
<td>Winter</td>
<td>-0.0487 *</td>
<td>-0.0958 *</td>
</tr>
<tr>
<td></td>
<td>(0.024)</td>
<td>(0.030)</td>
</tr>
<tr>
<td>Spring</td>
<td>-0.0759 *</td>
<td>-0.112 *</td>
</tr>
<tr>
<td></td>
<td>(0.022)</td>
<td>(0.027)</td>
</tr>
<tr>
<td>Autumn</td>
<td>-0.0247</td>
<td>-0.0826 *</td>
</tr>
<tr>
<td></td>
<td>(0.022)</td>
<td>(0.028)</td>
</tr>
<tr>
<td>Maritimes</td>
<td>-0.0241</td>
<td>0.0482</td>
</tr>
<tr>
<td></td>
<td>(0.059)</td>
<td>(0.067)</td>
</tr>
<tr>
<td>Quebec</td>
<td>-0.00567</td>
<td>0.161 *</td>
</tr>
<tr>
<td></td>
<td>(0.040)</td>
<td>(0.045)</td>
</tr>
<tr>
<td>Ontario</td>
<td>0.0353</td>
<td>0.179 *</td>
</tr>
<tr>
<td></td>
<td>(0.040)</td>
<td>(0.046)</td>
</tr>
<tr>
<td>Prairies</td>
<td>0.0954</td>
<td>0.205 *</td>
</tr>
<tr>
<td></td>
<td>(0.057)</td>
<td>(0.063)</td>
</tr>
<tr>
<td>B.C.</td>
<td>0.142 *</td>
<td>0.315 *</td>
</tr>
<tr>
<td></td>
<td>(0.046)</td>
<td>(0.053)</td>
</tr>
<tr>
<td>Autoworkers</td>
<td>0.115 *</td>
<td>0.477 *</td>
</tr>
<tr>
<td></td>
<td>(0.026)</td>
<td>(0.038)</td>
</tr>
<tr>
<td>Steelworkers</td>
<td>0.0725 *</td>
<td>0.208 *</td>
</tr>
<tr>
<td></td>
<td>(0.022)</td>
<td>(0.028)</td>
</tr>
<tr>
<td>REVENUE</td>
<td>0.0375 *</td>
<td>0.0374 *</td>
</tr>
<tr>
<td></td>
<td>(0.036)</td>
<td>(0.045)</td>
</tr>
<tr>
<td>Variance(2)</td>
<td>0.0302</td>
<td>0.106 *</td>
</tr>
<tr>
<td></td>
<td>(0.069)</td>
<td>(0.069)</td>
</tr>
</tbody>
</table>

lnL (s>0) 175.768 -285.5856 -336.6460
lnL (s=0) 175.8768 -371.0394 79.5547 15.0957
Total -656.6250 -321.5503

Chi-squared statistic(3) 711.9810 1297.6182
Observations 501 2459 1958 2459

Notes: (1) See Note (1) for Table 5.1. For the Tobit estimates, the asymptotic standard errors are given in parentheses.
(2) For the OLS results this is the variance of the estimate.
(3) Under the null hypothesis that all the coefficients except the constant term are zero.
are in most cases larger in absolute value than the OLS estimates. Note that the coefficient for the variable measuring variability in firm revenue is negatively related to the low wage offer, as predicted by the theory, though the effect is not statistically significant. Under Tobit estimation, a 1% increase in the unemployment rate reduces the low wage offer by about 7%, and during the period the Anti-Inflation Board was in effect low wage offers are estimated to have been reduced by 30%. All but one of the regional dummy variables is now significant, and all of the seasonal dummy variables are also significant.

Some of the Tobit coefficient estimates for the low wage offer equation do not appear to be intuitively plausible. For example, the U.A.W. is estimated to have low wage offers which are 48% greater than those of other unions. And low wage offers in B.C. are estimated to be 32% greater than the low wage offers for the excluded region. However, these estimates do not contradict the evidence concerning the behaviour of the observed wage reported in Table 5.1. The higher observed wages in B.C. and for the U.A.W. could be due to higher low wage offers. In this context, the likelihood function value for the Tobit estimates may be split into two values, one for the strike observations and one for the non-strike observations. The Tobit likelihood value for the strike observations is then directly comparable to the likelihood value from the OLS estimation. As reported in Table 5.2, the log likelihood value drops from 175.8 to -285.6 in moving from the OLS to the Tobit estimates. The Tobit coefficient estimates are also used to calculate an $R^2$ value for the 501 strike observations. The $R^2$ from the OLS regression is .5116 compared to .0396 from the Tobit estimates. Thus, when the non-strike observations are included in the estimation of the low wage offer equation, the resulting coefficient estimates explain almost none of the variation in the observed wages for the strike observations.

To investigate why the Tobit estimates fit the strike subsample so poorly, the following exercise is performed. The estimated Tobit coefficients are used to construct a fitted low wage offer for every observation in the sample, $\hat{w}_{it}^*$. According to the theory, the low wage offer should be less than the observed wage for all contracts without a strike. Thus the fitted low wage offer for each contract, $\hat{w}_{it}^*$, is compared to the observed wage for the same contract, $w_t$, for all contracts without a strike. The result is that $\hat{w}_{it}^* < w_t$ for all observations $t$ with no strike. That is, the Tobit estimates successfully impose the theoretical restriction on the data in every case for which the restriction applies. Next, the OLS coefficients in Table 5.2 are used to
construct a low wage offer for each non-strike observation. In this case, \( \tilde{\omega}_{1t} > \omega_t \) in just over 80% of the contracts without a strike. That is, the theoretical restriction that the low wage offer should be less than the observed non-strike wage is violated in the majority of contracts. Therefore, it appears that the reason the Tobit estimates fit the strike subsample so poorly is that the restriction the estimates impose is not in accordance with the data.

Of course, there is another restriction which has been imposed in the estimation of the S Model besides the monotonicity property—namely the assumption of no pooling equilibria. It could simply be that the latter restriction is the underlying cause of the problem. It will be interesting to see how the SP Model estimates fit the low wage equation for the non-strike observations, as the model allows for the occurrence of pooling equilibria.

Returning to Table 5.2, the two right-hand columns present the OLS estimates of the high wage offer equation over the 1,958 non-strike observations and the Tobit estimates over the full sample. Notice that the revenue variable is positive and significant for the OLS estimates—as predicted by the theory. On the other hand, variability in firm revenue is estimated to reduce significantly the high wage offer, which is counter to the theoretical prediction. However, the effect is reduced in magnitude moving from the OLS to the Tobit estimates. Generally speaking, the observation that the Tobit estimates are larger (in absolute value) than the corresponding OLS estimates, which was noted for the low wage offer equation estimates, applies to the high wage offer equation estimates as well. However, in contrast to the low wage offer equation, the Tobit estimates fit the subsample of data used for the OLS estimation reasonably well. The \( R^2 \) from the OLS regression over the non-strike observations is .4108 and the \( R^2 \) calculated using the Tobit estimates is .3930. And although the log likelihood value drops, from 79.6 to 15.1, moving from the OLS to the Tobit estimates, the change is not nearly as dramatic as for the low wage equation. The results in Table 5.2 indicate that high wage offers are procyclical and that the offers were reduced by 17% during the period the Anti-Inflation Board was in effect. In addition, high wage offers are significantly lower in the winter, spring and autumn relative to the summer months. Finally, high wage offers are larger in B.C. and for the United Autoworkers and the United Steelworkers.

In summary, Tobit estimation of the low wage offer equation is possibly hampered by the assumption that rules out pooling equilibria. Tobit estimation of the high

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wage offer equation, on the other hand, seems to be satisfactory. In addition, two things have been accomplished. First, the estimated models are an example of the application of an estimation strategy which applies to models of adverse-selection where pooling equilibria are believed to be unlikely events. Second, if there are pooling equilibria present in the data, it is not surprising that the S Model estimates for the low wage equation are implausible. Based on the evidence so far it is reasonable to expect the SP Model to perform much better.

Results from estimation of the strike offer equation and the so-called alternative Tobit model of strikes are presented in Table 5.3. The first column of the table lists the coefficient estimates from maximum likelihood estimation of the alternative Tobit strike model in (4.20). The two right-hand columns list the estimates of the strike offer equation with the OLS estimates for the positive strike duration observations on the left, and the maximum likelihood estimates over the entire sample on the right.

Consider the alternative Tobit strike model results. The coefficients are interpreted as the change in desired or potential strike length given a unit change in the independent variable. For example, each extra month of negotiations is estimated to increase desired strike length by just under 17 days. A 1% increase in the unemployment rate reduces desired strike length by roughly 9 days. The estimated impact of the Anti-Inflation Board is to reduce desired strike length by over a month. The unconditional mean strike duration is 14.4 days, while, conditional on the occurrence of a strike, mean duration is 55.5 days. According to the alternative Tobit model there are no significant seasonal or regional differences in desired strike length. The U.A.W. and U.S.W. are estimated to have substantially longer potential strikes. For example, the U.A.W. are estimated to have a desired strike duration which is two months longer relative to all other unions.

This last result indicates a weakness in the conventional Tobit model; if a variable increases the likelihood of a positive desired strike length, it also increases the conditional duration. Since the U.A.W. are known to have a significantly higher incidence of strikes relative to other unions, it is probably the case that the Tobit model estimates this as a large effect on desired strike length. Finally, both the revenue and

4. There appears to be no empirical study of strike duration using this Tobit model in the literature.
5. The coefficients may also be used to calculate changes in the probability of a positive desired strike length.
### Table 5.3: S Model - Strike Offer Equation and Alternative Tobit Strike Model Coefficient Estimates\(^{(1)}\)

<table>
<thead>
<tr>
<th>Variable</th>
<th>TOBIT Coefficient</th>
<th>OLS Coefficient</th>
<th>Tobit Coefficient</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-139.34</td>
<td>-25.317</td>
<td>69.378</td>
</tr>
<tr>
<td></td>
<td>(96.939)</td>
<td>(80.408)</td>
<td>(56.837)</td>
</tr>
<tr>
<td>NEMP</td>
<td>2.7108</td>
<td>-1.0860</td>
<td>-1.8160 *</td>
</tr>
<tr>
<td></td>
<td>(1.4411)</td>
<td>(.97062)</td>
<td>(.76481)</td>
</tr>
<tr>
<td>LNEG</td>
<td>16.793 *</td>
<td>10.957 *</td>
<td>5.1264 *</td>
</tr>
<tr>
<td></td>
<td>(1.0223)</td>
<td>(1.79715)</td>
<td>(.59518)</td>
</tr>
<tr>
<td>MANW</td>
<td>62.316 *</td>
<td>41.917 *</td>
<td>35.376 *</td>
</tr>
<tr>
<td></td>
<td>(27.207)</td>
<td>(22.345)</td>
<td>(15.950)</td>
</tr>
<tr>
<td>URM</td>
<td>-9.2817 *</td>
<td>-3.5578</td>
<td>.17162</td>
</tr>
<tr>
<td></td>
<td>(2.7843)</td>
<td>(2.2294)</td>
<td>(1.5881)</td>
</tr>
<tr>
<td>YEAR</td>
<td>-2.4176</td>
<td>-1.7260</td>
<td>-2.0417</td>
</tr>
<tr>
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<td>(2.5066)</td>
<td>(2.0597)</td>
<td>(1.4683)</td>
</tr>
<tr>
<td>DAIB</td>
<td>-48.332 *</td>
<td>-22.924 *</td>
<td>-11.127</td>
</tr>
<tr>
<td></td>
<td>(11.402)</td>
<td>(9.4339)</td>
<td>(6.6801)</td>
</tr>
<tr>
<td>Winter</td>
<td>-9.6281</td>
<td>3.0618</td>
<td>8.7389</td>
</tr>
<tr>
<td></td>
<td>(8.1045)</td>
<td>(6.7918)</td>
<td>(4.8787)</td>
</tr>
<tr>
<td>Spring</td>
<td>-13.671</td>
<td>-22.196 *</td>
<td>-9.9965 *</td>
</tr>
<tr>
<td></td>
<td>(7.2433)</td>
<td>(5.8984)</td>
<td>(4.0712)</td>
</tr>
<tr>
<td>Autumn</td>
<td>-1.0311</td>
<td>14.162 *</td>
<td>13.079 *</td>
</tr>
<tr>
<td></td>
<td>(7.4989)</td>
<td>(6.0563)</td>
<td>(4.5335)</td>
</tr>
<tr>
<td>Maritimes</td>
<td>-26.669</td>
<td>-6.3414</td>
<td>-4.1116</td>
</tr>
<tr>
<td></td>
<td>(21.127)</td>
<td>(17.927)</td>
<td>(11.739)</td>
</tr>
<tr>
<td>Quebec</td>
<td>-6.5120</td>
<td>13.966</td>
<td>4.2408</td>
</tr>
<tr>
<td></td>
<td>(15.279)</td>
<td>(12.887)</td>
<td>(8.4633)</td>
</tr>
<tr>
<td>Ontario</td>
<td>-1.0745</td>
<td>13.606</td>
<td>2.0215</td>
</tr>
<tr>
<td></td>
<td>(15.283)</td>
<td>(12.962)</td>
<td>(8.4222)</td>
</tr>
<tr>
<td>Prairies</td>
<td>-28.230</td>
<td>.39764</td>
<td>10.716</td>
</tr>
<tr>
<td></td>
<td>(22.304)</td>
<td>(20.580)</td>
<td>(12.295)</td>
</tr>
<tr>
<td>B.C.</td>
<td>-7.7594</td>
<td>-15.342 *</td>
<td>-20.976 *</td>
</tr>
<tr>
<td></td>
<td>(17.782)</td>
<td>(14.374)</td>
<td>(9.6782)</td>
</tr>
<tr>
<td>Autoworkers</td>
<td>59.993 *</td>
<td>-7.5658</td>
<td>-20.394 *</td>
</tr>
<tr>
<td></td>
<td>(9.3293)</td>
<td>(6.5581)</td>
<td>(5.0973)</td>
</tr>
<tr>
<td>Steelworkers</td>
<td>33.618 *</td>
<td>5.3522</td>
<td>-1.6689</td>
</tr>
<tr>
<td></td>
<td>(7.5566)</td>
<td>(6.0525)</td>
<td>(4.3737)</td>
</tr>
<tr>
<td>REVENUE</td>
<td>-1.7860</td>
<td>-4.8884 *</td>
<td>-4.1515 *</td>
</tr>
<tr>
<td></td>
<td>(3.8033)</td>
<td>(2.5893)</td>
<td>(2.0465)</td>
</tr>
<tr>
<td>C.V. REVENUE</td>
<td>-18.584</td>
<td>8.3212</td>
<td>6.6169</td>
</tr>
<tr>
<td></td>
<td>(15.813)</td>
<td>(14.175)</td>
<td>(9.6162)</td>
</tr>
<tr>
<td>Variance</td>
<td>5828.8644</td>
<td>1709.9052</td>
<td>1448.8792 *</td>
</tr>
</tbody>
</table>

InL: \(-2650.9514\)

\(\text{R-squared} : 0.4732\)

\(\text{Chi-squared statistic} : 401.6744\)

Observations: 1516

Notes: (1) See notes for Table 5.2.
variation in revenue variables are statistically insignificant in the alternative Tobit model of strikes.

Now, examine the Tobit estimates of the strike offer equation. In terms of the log likelihood, the strike offer model with a value of $-2158.2$ is preferred to the alternative Tobit model with a value of $-2651.0$. Contrary to the alternative Tobit model estimates, there is evidence of seasonal variation in strike offers. Relative to the summer months, strike offers are estimated to be longer in the autumn and shorter in the spring. For each additional thousand members in the bargaining unit, strike offers are estimated to decrease by just under two days. Every extra month of contract negotiations increases the strike offer by just over 5 days. In contrast to the alternative Tobit model, there is no statistically significant evidence of cyclical variation in strike offers and neither is there any evidence of an effect on strike offers by the Anti-Inflation Board. It is also interesting to note that the U.A.W. is estimated to have significantly shorter strike offers, by almost 20 days, than other unions. Finally, the effect of the variability in firm revenue is estimated to be positive, as predicted by the theory, but not statistically significant. Furthermore, firm revenue is found to have a significant negative influence on strike offers for the observations with a strike, i.e., the OLS estimates, as predicted by the theory. Over the entire sample, the effect of firm revenue on the strike offer is reduced, also as expected.

5.2.2 S Model Estimates—Dependent Disturbance Terms

Allowing for contemporaneous correlation between the error terms in the low and high wage offer equations yields the likelihood function in (4.18). An attempt was made to estimate this model using the independent variables listed in Table 5.2, but the attempt was unsuccessful. Regardless of the chosen coefficient starting values, the optimizing algorithms would venture into regions of the parameter space where the correlation coefficient for the disturbance terms is +1 and where $\beta_1$ approached $\beta_2$. To see why this might happen, consider the wage offer equations. If $s_t > 0$, it is known that $w^*_1 = w_t$ and $w^*_2 > w_t$, which by substitution implies

$$w_t = x_t \beta_1 + u_{1t},$$

$$w_t < x_t \beta_2 + u_{2t},$$

(5.1)

where $x_t$ has been substituted for $x_{1t}$ and $x_{2t}$ as the same set of independent variables is used in the low and high wage offer equation. If $s_t = 0$, it is known that $w^*_1 < w_t$
and $w^*_2 = w_t$, thus

$$w_t > x_t \beta_1 + u_{1t},$$

$$w_t = x_t \beta_2 + u_{2t}. \quad (5.2)$$

It is clear from (5.1) and (5.2) that as $\beta_1 \to \beta_2$, the correlation between $u_{1t}$ and $u_{2t}$ will become almost perfect.

The obvious way around this problem is to choose an explanatory variable in $x_{1t}$ that is not in $x_{2t}$, or vice versa. According to the theory there is a natural candidate for this additional independent variable. The comparative statics predictions of the adverse-selection model state that the low wage offer is independent of $q$, the union’s subjective probability the firm is the high demand type. Thus, according to the theory, a proxy for $q$ may be included in the $x_{2t}$ vector of explanatory variables and left out of the $x_{1t}$ vector.

An attempt was made to follow this procedure. The proxy chosen for $q$ is the Index of Industrial Production (IIP) for the same year as the effective date of the contract. The annual index is constructed as follows. First, the seasonally adjusted quarterly Gross Domestic Product at factor cost in 1981 prices is summed across quarters for each year 1964–86 to arrive at an annual figure. Second, the annual series for GDP at factor cost is normalized with 1981 as the base year resulting in the IIP. The advantage with using the IIP as a proxy for $q$ is that it is readily available. However, it is a long way from an ideal proxy for the union’s probability the firm is the high demand type. For example, it would be more in the spirit of the model to have a variable at the level of individual firms rather than at the economy-wide level. In addition, the aggregate Index is certain to be highly collinear with the unemployment rate of males age 25 and over, which is also an element of $x_{2t}$. One possibility is to construct a different IIP for each industry in the data.

The model is re-estimated with the addition of the proxy for $q$ as an explanatory variable in the high wage equation. Again, the attempt was unsuccessful. This time, the algorithms succeeded in reducing the coefficient on the IIP to zero and then proceeded to the same point in parameter space as before. However, it should be noted that OLS estimation of the high wage offer equation over the non-strike observations produces a coefficient value on the proxy for $q$ which is not significantly different from zero. Perhaps if a better proxy is found for $q$, the S Model for wage...
offers with dependent error terms will be estimable.

5.3 Results from the SP Model

Estimation of the SP Model is undertaken in two stages. First, the low, high and pooling wage offer equations are estimated over the larger contract data set with 2,459 observations. Next, the entire model, i.e., the three wage offer equations and the strike offer equation, is estimated over the data set with 1,516 observations for which strike duration information are available.

5.3.1 SP Model Results—Wage Offer Equations

The likelihood function is not globally concave for the SP Model, unlike the likelihood for the S Model, and in fact it is unbounded for certain parameter values (see Section 4.3.2). The estimation procedure is to find the highest local maximum. Two sets of starting values are used for the parameters $\beta_1, \beta_2, \beta_P, \sigma_1^2, \sigma_2^2, \sigma_P^2$. First, the OLS estimates from Table 5.2 are used as starting values. The low wage offer equation coefficients and variance of the estimate are used as starting values for $\beta_1$ and $\sigma_1^2$ respectively. The high wage offer equation coefficients and variance of the estimate are used as starting values for $\beta_2$ and $\beta_P$, and $\sigma_2^2$ and $\sigma_P^2$. The rationale for using the same starting values for the high and pooling wage offer equation parameters is that, since both offers occur in the absence of a strike, the OLS estimates over the non-strike observations, in some sense, are robust indicators of the underlying parameters. The Tobit estimates from Table 5.2 are also used as starting values for the SP Model of wage offers, and again identical starting values are used for the parameters in the high and pooling wage offer equations. Remarkably, both sets of starting values converged to the same point which is reported in Table 5.4.

The value of the log likelihood at the local maximum is $-218.1$ which compares very favourably to the log likelihood for the S Model, $-978.1$.\(^8\) The Chi-squared test statistic for the null hypothesis that all the coefficients in each wage offer equation except the constant term are zero has a value of 1319.0 with 23 degrees of freedom. The most interesting aspects of the estimates in Table 5.4 are as follows. The number of employees does not significantly affect the low or high wage offer equations, but larger bargaining units are estimated to have significantly lower pooling wage offers.

\(^8\) The value of the log likelihood for the S Model is found by simply adding the values for the low and high wage offer equation estimates, as shown in (4.17).
### Table 5.4: SP Model - Wage Offer Equation Coefficient Estimates

<table>
<thead>
<tr>
<th>Variable</th>
<th>Low wage offer</th>
<th>High wage offer</th>
<th>Pooling wage offer</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-0.322 (0.28)</td>
<td>-0.958 * (0.15)</td>
<td>0.957 * (0.42)</td>
</tr>
<tr>
<td>NEMP</td>
<td>0.00407 (0.0032)</td>
<td>0.00279 (0.0019)</td>
<td>~0.00811 * (0.0042)</td>
</tr>
<tr>
<td>LNEG</td>
<td>0.00877 * (0.0027)</td>
<td>0.0168 * (0.0017)</td>
<td>0.0162 * (0.0035)</td>
</tr>
<tr>
<td>MANW</td>
<td>0.390 * (0.085)</td>
<td>0.302 * (0.043)</td>
<td>0.621 * (0.11)</td>
</tr>
<tr>
<td>URM</td>
<td>-0.0302 * (0.0078)</td>
<td>-0.0461 * (0.0043)</td>
<td>-0.0292 * (0.0013)</td>
</tr>
<tr>
<td>YEAR</td>
<td>0.00478 (0.0075)</td>
<td>0.0183 * (0.0040)</td>
<td>0.0205 * (0.0011)</td>
</tr>
<tr>
<td>DAIB</td>
<td>-0.101 * (0.034)</td>
<td>-0.127 * (0.017)</td>
<td>-0.0177 * (0.041)</td>
</tr>
<tr>
<td>REVENUE</td>
<td>0.0371 * (0.0097)</td>
<td>0.0478 * (0.0057)</td>
<td>0.0676 * (0.016)</td>
</tr>
<tr>
<td>C.V. REVENUE</td>
<td>0.0935 * (0.036)</td>
<td>0.0550 * (0.022)</td>
<td></td>
</tr>
<tr>
<td>Variance</td>
<td>0.0443 * (0.0032)</td>
<td>0.0170 * (0.0014)</td>
<td>0.0677 * (0.0037)</td>
</tr>
</tbody>
</table>

**Conditional descriptive statistics:**

<table>
<thead>
<tr>
<th>Number of observations</th>
<th>481</th>
<th>1,665</th>
<th>313</th>
</tr>
</thead>
<tbody>
<tr>
<td>R-squared</td>
<td>0.3996</td>
<td>0.2957</td>
<td>0.2923</td>
</tr>
</tbody>
</table>

**Notes:**

(1) See the notes for Table 5.2. The value of the log likelihood from the full model with 2,459 observations is -218.1044. The log likelihood value for the model restricted to a constant term in each equation is -877.6025, yielding a Chi-squared test statistic of 1318.9962 with 23 degrees of freedom.

(2) The descriptive statistics are conditional on the sample separation from the predicted values of the maximum likelihood coefficient estimates. The R-squared is calculated as \([1 - \text{Var}(u)/\text{Var}(w)]\) for each equation.
Pooling wage offers are much less sensitive to fluctuations in the unemployment rate and the trend term than either the low or high wage offers. The Anti-Inflation Board is estimated to have significantly reduced the high and the low wage offer, while the pooling wage offer was largely unaffected. As predicted by the theory, pooling wage offers depend positively on the revenue at the firm, and high wage offers depend positively on the variability of firm revenue. Also as predicted by the theory, the high wage and the pooling wage offer depend positively on the alternative wage available to union members, as indicated by the coefficient on MANW. Contrary to the theory, low wage offers are estimated to be positively related to the variability in firm revenue.

Recall from the discussion of the empirical implementation in Section 4.2 that the wage offers $w_1^*$ and $w_2^*$ are from an unconstrained version of the union’s maximization problem. If $w_2^* \leq w_1^*$ then a pooling solution obtains. Using the parameter estimates in Table 5.4 an estimated value for the low wage offer, $\hat{w}_1^t$, and for the high wage offer, $\hat{w}_2^t$, is calculated for every observation $t$. Contracts for which $\hat{w}_2^t \leq \hat{w}_1^t$ are classified as pooling offers. According to the theory, all strike observations are the result of separating offers, so if the estimated model were to predict perfectly then $\hat{w}_1^t < \hat{w}_2^t$ for all 501 strike observations. It turns out that there are 20 strike observations with $\hat{w}_2^t < \hat{w}_1^t$, representing a “success” rate of 96%. Estimated low and high wage offers are also calculated for the 1958 non-strike observations, yielding 1,665 separating offers and 293 pooling offers. Thus, over the full sample, there are 313 predicted pooling offers representing 12.7% of all contracts.

To get an idea of how well the predicted wage offers correspond to the observed wage outcomes, an $R^2$ value is calculated for each wage offer equation as follows. All contracts which have a strike and for which $\hat{w}_1^t < \hat{w}_2^t$ are classified as wage outcomes from low wage offers.\(^9\) All contracts without strikes and for which $\hat{w}_1^t < \hat{w}_2^t$ are classified as wage outcomes from high wage offers, while those for which $\hat{w}_2^t \leq \hat{w}_1^t$ are classified as pooling wage outcome. Based on this sample separation, an $R^2$ value is calculated for each class of observations as $1 - \text{Var}(\hat{u}_it)/\text{Var}(w_t)$, where $\hat{u}_it = w_t - \hat{w}_it$ is the estimated residual in the $i$th offer equation for $i = 1, 2, P$. As reported in Table 5.4, the $R^2$ for the low wage equation is .3996, for the high wage equation the $R^2$ value is .2957 and for the pooling equation the $R^2$ is .2923. The SP Model provides a reasonable fit to the observed wage data, especially considering the cross-sectional nature of the data. It was hypothesized in the discussion of results for the S Model

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\(^9\) See equation (4.3).
that the SP Model would provide a better fit of the wage data. The evidence presented in Table 5.4 supports this hypothesis.

An interesting feature of the SP Model is that, for every observation, the probability of a separating equilibrium may be calculated as follows

\[ \Pr(\text{separating offer in contract } t) = \Pr(w_{2t}^* > w_{1t}^*) \]
\[ = \Pr(u_{2t} - u_{1t} > x_t \beta_1 - x_t \beta_2) \]
\[ = 1 - \Phi(u_t), \]

where \( x_t = x_{1t} = x_{2t}, \Phi = (x_t \beta_1 - x_t \beta_2)/\sigma, \sigma^2 = \sigma_1^2 + \sigma_2^2, \) and \( \Phi \) is the standard normal distribution function. The impact of the \( j \)th independent variable on the probability of a separating equilibrium in the \( t \)th contract is thus

\[ \frac{\partial \Pr(\text{separating offer in contract } t)}{\partial x_{jt}} = -\frac{\partial \Phi(u_t)}{\partial x_{jt}} \]
\[ = -\frac{\phi(u_t)}{\sigma} (\beta_{1j} - \beta_{2j}), \]

where \( \phi \) is the standard normal density function. Since \( \phi(u_t)/\sigma > 0 \) for all \( t \), \( \partial \Pr(\text{separating offer in contract } t)/\partial x_{jt} > 0 \) if and only if \( (\beta_{2j} - \beta_{1j}) > 0 \). Therefore, simply by taking the difference between the coefficients from the high and low wage offer equations we may determine the effect of an independent variable on the probability of a separating equilibrium. Furthermore, as the coefficient estimates are asymptotically normal, the standard error of the difference \( (\beta_{2j} - \beta_{1j}) \) may be calculated from the asymptotic covariance matrix of the \( \beta \)'s.

Table 5.5 reports the coefficient differences and the corresponding standard errors for the estimates in Table 5.4. Thus, separating equilibria are more likely for longer negotiations, less likely the higher the unemployment rate, more likely in recent years and less likely during the Anti-Inflation Board period. According to the theory, separating equilibria are more likely the greater the degree of variability in firm revenue. This prediction is not confirmed for the results presented in Table 5.5 where the effect of the coefficient of variation of revenue on the probability of a separating equilibrium is negative and not significantly different from zero.

The adverse-selection model of strikes predicts that strikes occur only when the union’s offers constitute a separating equilibrium. In addition, the probability of a strike in contract \( t \) is \( (1 - q_t) \Pr(\text{separating offer in contract } t) \). Since \( 0 < q_t < 1 \) for
Table 5.5: SP Model - Difference Between the High and Low Wage Offer Equation Coefficient Estimates

<table>
<thead>
<tr>
<th>Variable</th>
<th>Difference between High and Low wage offer coefficient estimate (1)</th>
<th>Asymptotic t-statistic (2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-.636 (.29)</td>
<td>- 2.23</td>
</tr>
<tr>
<td>NEMP</td>
<td>-.00128 (.0031)</td>
<td>- .408</td>
</tr>
<tr>
<td>LNEG</td>
<td>.00807 (.0027)</td>
<td>3.03</td>
</tr>
<tr>
<td>MANW</td>
<td>-.0875 (.082)</td>
<td>- 1.07</td>
</tr>
<tr>
<td>URM</td>
<td>-.0159 (.0086)</td>
<td>- 1.86</td>
</tr>
<tr>
<td>YEAR</td>
<td>.0135 (.0075)</td>
<td>1.81</td>
</tr>
<tr>
<td>DAIB</td>
<td>-.0258 (.033)</td>
<td>- .780</td>
</tr>
<tr>
<td>REVENUE</td>
<td>.0107 (.010)</td>
<td>1.06</td>
</tr>
<tr>
<td>C.V. REVENUE</td>
<td>-.0386 (.034)</td>
<td>- 1.15</td>
</tr>
</tbody>
</table>

Notes: (1) The coefficient estimates are shown in Table 5.4. Any discrepancies between the differences given above and differences calculated directly from Table 5.4 are due to rounding error.

(2) Asymptotic standard errors are shown in parentheses. The critical value of the t-statistic at the .10 level, based on 2430 degrees of freedom, is 1.64.
all \( t \), the probability of a strike and the probability of a separating equilibrium are positively related. Therefore, the results in Table 5.5 may be related to the probability of a strike. Provided \( q_t \) is independent of \( x_t \) there will be a monotonic relationship between the probability of a strike and the differences reported in Table 5.5. In this context, it is interesting to note that the results suggest that the probability of a strike increases with the length of contract negotiations, decreases with the unemployment rate, increases over the period 1964–86 and decreases during the Anti-Inflation Board period—exactly the same results reported in the logit model of strike incidence in chapter 3. There is some evidence, then, that the stylized facts about strike incidence can be explained, at least in part, by the SP Model of wage offers.

5.3.2 SP Model Results—Wage and Strike Offer Equations

The most ambitious econometric model proposed here is joint estimation of the wage and strike offer equations in the model which allows for the occurrence of pooling equilibria. The results from estimation of the model are presented in Table 5.6. Note that estimation is performed over the data set with 1,516 observations from the manufacturing sector, as these are the only data with strike duration information. Also, seasonal effects are included as independent variables for the strike offer equation and not in the wage offer equations.

As predicted by the adverse-selection model, firm revenue is significantly positively related to the pooling wage offer, variability in firm revenue is significantly positively related to the high wage offer (at the .10 level) and positively related to the strike offer. Counter to the theoretical prediction, variability in firm revenue is significantly positively related to the low wage offer. There is evidence that firm revenue is negatively related to strike offers and positively related to both high and low wage offers, though the predictions of the theory are ambiguous in all these cases. Relative to the summer months, strike offers are significantly shorter in the spring and significantly longer in the autumn by roughly 12 and 14 days respectively. There is no statistically significant evidence of any cyclical trend in strike offers, nor of any effect of the Anti-Inflation Board on strike offers.

Generally speaking, the wage offer equation estimates in Table 5.6 are similar to the estimates presented in Table 5.4 for the full sample. This indicates the robustness of the econometric model to data samples with slightly different characteristics. The high wage offer is more sensitive than the low wage offer to the length of negotiations,
Table 5.6: SP Model - Wage and Strike Offer Equation Coefficient Estimates(1)

<table>
<thead>
<tr>
<th>Equation:</th>
<th>Low wage offer</th>
<th>High wage offer</th>
<th>Pooling wage offer</th>
<th>Strike offer</th>
</tr>
</thead>
<tbody>
<tr>
<td>Variable:</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>-.551</td>
<td>-.829 *</td>
<td>.757 *</td>
<td>49.7</td>
</tr>
<tr>
<td></td>
<td>(.29)</td>
<td>(.17)</td>
<td>(.042)</td>
<td>(61.)</td>
</tr>
<tr>
<td>NEMP</td>
<td>.00610</td>
<td>.0107 *</td>
<td>.0234 *</td>
<td>-2.09 *</td>
</tr>
<tr>
<td></td>
<td>(.0040)</td>
<td>(.0027)</td>
<td>(.010)</td>
<td>(.81)</td>
</tr>
<tr>
<td>LNEG</td>
<td>.0110 *</td>
<td>.0230 *</td>
<td>.00531 *</td>
<td>6.18 *</td>
</tr>
<tr>
<td></td>
<td>(.0031)</td>
<td>(.0019)</td>
<td>(.0052)</td>
<td>(.65)</td>
</tr>
<tr>
<td>MANW</td>
<td>.036 *</td>
<td>.295 *</td>
<td>.583 *</td>
<td>32.9</td>
</tr>
<tr>
<td></td>
<td>(.087)</td>
<td>(.047)</td>
<td>(.12)</td>
<td>(18.)</td>
</tr>
<tr>
<td>URM</td>
<td>-.0353 *</td>
<td>-.0423 *</td>
<td>.0180</td>
<td>-.239</td>
</tr>
<tr>
<td></td>
<td>(.0084)</td>
<td>(.0053)</td>
<td>(.015)</td>
<td>(1.8)</td>
</tr>
<tr>
<td>YEAR</td>
<td>.0104</td>
<td>.0164 *</td>
<td>-.0245 *</td>
<td>-1.80</td>
</tr>
<tr>
<td></td>
<td>(.0078)</td>
<td>(.0044)</td>
<td>(.0011)</td>
<td>(1.6)</td>
</tr>
<tr>
<td>DAIB</td>
<td>-.115 *</td>
<td>-.135 *</td>
<td>-.177 *</td>
<td>-11.0</td>
</tr>
<tr>
<td></td>
<td>(.035)</td>
<td>(.019)</td>
<td>(.048)</td>
<td>(7.4)</td>
</tr>
<tr>
<td>Winter</td>
<td>--</td>
<td>--</td>
<td>--</td>
<td>5.73</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(5.4)</td>
</tr>
<tr>
<td>Spring</td>
<td>--</td>
<td>--</td>
<td>--</td>
<td>-12.3 *</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(4.5)</td>
</tr>
<tr>
<td>Autumn</td>
<td>--</td>
<td>--</td>
<td>--</td>
<td>14.0</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td>(4.9)</td>
</tr>
<tr>
<td>REVENUE</td>
<td>.0397 *</td>
<td>.0604 *</td>
<td>.120 *</td>
<td>-6.83 *</td>
</tr>
<tr>
<td></td>
<td>(.0098)</td>
<td>(.0064)</td>
<td>(.022)</td>
<td>(2.1)</td>
</tr>
<tr>
<td>C.V. REVENUE</td>
<td>.0984 *</td>
<td>.0383</td>
<td>-</td>
<td>9.00</td>
</tr>
<tr>
<td></td>
<td>(.046)</td>
<td>(.0028)</td>
<td></td>
<td>(11.)</td>
</tr>
<tr>
<td>Variance</td>
<td>.0309 *</td>
<td>.0144 *</td>
<td>.0418 *</td>
<td>1552.9 *</td>
</tr>
<tr>
<td></td>
<td>(.0025)</td>
<td>(.0012)</td>
<td>(.0036)</td>
<td>(114.4)</td>
</tr>
</tbody>
</table>

Conditional(2) descriptive statistics:

<table>
<thead>
<tr>
<th>Number of observations</th>
<th>387</th>
<th>1043</th>
<th>86</th>
<th>387</th>
</tr>
</thead>
<tbody>
<tr>
<td>R-squared</td>
<td>.5450</td>
<td>.3953</td>
<td>.2412</td>
<td>.3782</td>
</tr>
</tbody>
</table>

Notes: (1) See the notes for Table 5.2. The value of the log likelihood for the full model with 1516 observations is -2009.398. The log likelihood value for the model restricted to a constant term in each equation is -2670.5752, yielding a Chi-squared test statistic of 1322.3542 with 34 degrees of freedom.

(2) See note (2) for Table 5.4.
the unemployment rate, the trend term and the Anti-Inflation Board in both sets of results. The results in Tables 5.4 and 5.6 indicate that the pooling wage is not very sensitive to the aggregate business cycle. Furthermore, the result that the pooling wage has decreased over the sample period, reported in Table 5.4, is confirmed for the manufacturing subsample.

As before, the coefficient estimates from the \( w_1^* \) and \( w_2^* \) offer equations may be used to predict which observations correspond to pooling equilibria and which to separating equilibria. For the estimates in Table 5.6, 387 of the 393 strike observations are predicted to be the outcome of separating equilibria, representing a "success" rate of 99%. For the total 1,516 contracts only 86 or just over 5% are predicted to be the outcome of pooling offers. Thus, the manufacturing subsample has proportionately half as many pooling equilibria as the full sample of 2,459 contracts. Given that there are proportionately more strikes in the manufacturing subsample, the apparent birth of pooling equilibria there is not in contradiction with the data.

5.4 Comparison of the Results from the Econometric Model of Wage and Strike Offers to the Empirical Literature.

As stated periodically throughout the present study, there are several aspects concerning the behaviour of strikes which appear in one empirical study after another. Perhaps the most well known result is that strike incidence is procyclical. According to the adverse-selection model of collective bargaining, there is a systematic relationship between the union's wage and strike offers. One of the most striking aspects of the results from the estimation of the SP Model is that the estimated wage offers are entirely consistent with the stylized fact that strikes are procyclical. Furthermore, the result is seen to be a consequence of the fact that the high wage offer is more sensitive than the low wage offer to fluctuations in the business cycle. Another stylized fact concerning strike behaviour is that strike activity was significantly reduced during the period the Anti-Inflation Board was administering wage and price guidelines. Results from the SP Model suggest that strike activity was reduced during this period because the high wage offers of unions were reduced more than the low wage offers. Recent empirical evidence suggests that while strike incidence is procyclical, strike duration is countercyclical. Estimation results from the SP Model indicate that strike offers are

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10 The strike incidence in the manufacturing sector is 25% compared to 20% in the overall sample.
not sensitive to the business cycle. Thus, according to the results here, the observed countercyclical behaviour of strikes is due to factors other than unions' strike offers.

The essential differences between the present econometric model and previous empirical studies are twofold. First, is the notion that wage and strike outcomes observed in collective bargaining contract data are the result of offers made by the union to the firm in order to resolve a problem of asymmetric information. Second, is the prediction of the theoretical model that these wage and strike offers are systematically related and the recognition that the relationships may be used to derive an estimable model. The econometric model derived here, it is believed, is applicable in other instances where asymmetric information plays an important role generating economic data. This claim is based on the fact that the key condition used to derive the econometric model of offers is a monotonicity property which is fundamental to all models of adverse-selection.
CHAPTER VI

Summary, Conclusions and Topics for Further Research

6.1 Summary

The aim of the present dissertation is to examine, theoretically and empirically, a model of collective bargaining between unions and firms based on the theory of asymmetric information. Specifically, the firm is assumed to be one of two "types," where each type is assumed to have a different level of output demand. The union knows that the firm is one of the two types, but it does not know which type the firm is. The model is characterized as a problem of adverse-selection—the union designs wage and strike offers to maximize expected utility over the firm types subject to self-selection constraints. The optimal wage and strike offers are shown to possess a monotonicity property under an assumption which guarantees that the higher demand firm type is willing to pay larger wages for a given reduction in strike length. The solution to the union's problem takes one of two forms: in the case of a separating equilibrium, the union makes a high wage–no strike offer and a low wage–strike offer; in the case of a pooling equilibrium, the union makes a single wage–no strike offer. Comparative static results are derived for all the exogenous variables in the model. For a separating equilibrium solution, it is demonstrated that wage offers depend positively on own-firm-type and negatively on cross-firm-type. The strike offer depends positively on the level of demand at the high demand firm, and negatively on the level of demand at the low demand firm type. For a pooling equilibrium solution, it is shown that the wage depends positively on the level of demand at both the high and the low demand firm type. In addition, it is shown that the union's strike threshold probability depends positively on the difference between the levels of demand at the two firm types.

The set-up of the theoretical model follows Hayes (1984). However, in the present study, more attention is given to the properties of the firm profit function and the union utility function than in Hayes (1984). This allows new comparative static properties to be derived. In particular, the comparative statics with respect to firm type have not appeared in the asymmetric information literature on strikes. As well, the notion of the strike threshold probability and its comparative static properties is also new to the literature. Furthermore, it turns out that the comparative static properties of the model generalize to the case where the number of firm types is
greater than two without any additional assumptions than were required to derive the results in the two-type model.

For empirical implementation of the model it is necessary to have wage and strike information from a series of collective agreements and data on the level of demand at the firm. Thus data on annual firm revenue is added to 2,459 contracts from the Major Collective Agreements computer tape for Canada covering the period 1964-86. The level of demand at the firm is proxied by the firm revenue for the fiscal year which ends at least six months after the effective date of the contract. The difference between the levels of demand at the two firm types is proxied by the coefficient of variation of firm revenue over the entire period the firm is present in the contract data. An econometric model of strike incidence is specified, following the conventional approach in the literature represented by Gunderson, Kervin and Reid (1986), with explanatory variables suggested by the adverse-selection model and additional variables designed to control for cyclical and trend effects and various forms of unobserved heterogeneity. The results indicate that, as predicted by the theory, the incidence of strikes is negatively related to the level of demand at the firm; and positively related to the difference between the levels of demand at the two firm types. The estimation results are robust to changes in the definition of firm revenue, the estimation technique and the choice of explanatory variables.

The data assembled here represent the first time firm-specific data have been used to test the adverse-selection model of strikes. Previous studies, e.g., McConnell (1987a, 1987b), proxy the level of demand at the firm using industry data. The pooled regressions of firm revenue also represent the first attempt to examine the causality between strikes and revenue.

In the adverse-selection model of collective bargaining, the wage level is an endogenous variable as well as the occurrence and the length of the strike. The adverse-selection model also predicts that the wage and strike offers are systematically related according to the monotonicity property. The analysis of chapter 4 recognizes both these facts in building an econometric model of wage and strike offers. The econometric problem posed is to model wage and strike offers when only the wage and strike outcome is observed in the data. The likelihood function used to derive the coefficient estimates of the wage and strike offer equations is derived using the monotonicity property implied by the theoretical model. An attractive feature of the SP Model is that the occurrence of separating and pooling equilibria is endogenously
determined by the data. If it is assumed that pooling equilibria are not observed in the data, the likelihood function simplifies to the Tobit function resulting in the S Model.

Results from the S Model indicate that the restriction which rules out pooling equilibria is probably not valid. Nevertheless, estimation of the S Model is an illustration of an estimation procedure which would apply in cases where there are a priori reasons to expect that pooling equilibria do not occur. Results from the SP Model indicate, for the sample of 2,459 contracts, pooling equilibria occur in approximately 12% of contract negotiations. All the unambiguous predictions of the theoretical model are confirmed, with one exception. Thus wage offers are estimated to depend positively on own-firm-type and strike offers are found to depend negatively on the level of demand at the firm. The variability of demand at the firm is estimated to increase the high wage offer and increase the strike offer, following the theoretical predictions, but to increase the low wage offer, counter to a prediction of the theory.

A particularly interesting result from the SP Model is that the high wage offer is estimated to be more sensitive than the low wage offer to various aspects of the collective bargaining environment. For example, the high wage offer is found to respond to the cyclical conditions, the length of negotiations and the trend term more than the low wage offer. Since the probability of a separating equilibrium in each contract is determined by the difference between the coefficient estimates from the high and low wage offer equations, the result that the high wage offer is more sensitive to the bargaining environment than the low wage offer indicates that separating equilibria are more likely during upswings in the business cycle, after longer negotiations and over time. Since the probability of a separating equilibrium is monotonically related to the probability of a strike, these results are consistent with the regularities found in the estimation of strike incidence equations that strikes are procyclical, more likely after long negotiations and over time.

6.2 Further Research

Several avenues of future research are suggested by the estimation results and the data set assembled in the present thesis. One line of investigation exploits the fact that the firm revenue data assembled here could be used to test more thoroughly the causation between strikes and the level of demand. According to the theory, strikes are the result of the presence of a low demand firm type. That is, low demand at the
firm causes strikes. Thus, using firm revenue data and information on the occurrence of a strike, it should be possible to examine the causality between revenue and strikes.

Clearly, the result that separating equilibria follow certain systematic trends with respect to the bargaining environment requires further investigation. One of the most robust results in the empirical literature on strikes is that strike activity follows a procyclical trend. The finding that wage offers also follow a procyclical trend, therefore, and that this behaviour is entirely consistent with the facts is potentially an important result in the context of the empirical testing of not only adverse-selection models of strikes but also of adverse-selection models in general.

Finally, the econometric model of wage and strike offers is based upon the monotonicity property of the offers derived from the adverse-selection model. In addition, the maximum likelihood procedure allows the data to estimate which observations correspond to pooling equilibria and which to separating equilibria. Since all adverse-selection models have a monotonicity property, the model proposed here is, in principle, applicable to the estimation of alternative empirical models based on adverse-selection theory.
BIBLIOGRAPHY


APPENDIX A:

ESTIMATES OF THE STRIKE INCIDENCE EQUATION USING DIFFERENT STRIKE INFORMATION
Table A.1: OLS Estimates of Strike Incidence for the Manufacturing Subsample Using the MCA Tape Strike Information(1)

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Notes: (1) There are 1,516 contracts. Proportion of contracts signed after a strike is 372/1516 = .245.

(2) Significance is denoted by ** at the .05 level and * at the .10 level, where the critical values, respectively, are 1.65 and 1.28 for the one-tailed test (when the expected sign is unambiguous) and 1.96 and 1.65 for the two-tailed test (when the expected sign is ambiguous).
Table A.2: OLS Estimates of Strike Incidence for the Manufacturing Subsample Using the Card (1987) Strike Information (1)

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Notes: (1) There are 1,516 contracts. The proportion of contracts signed after a strike is 409/1516 = .270.

(2) See notes for Table A.1.
APPENDIX B:

THE DERIVATIVES OF THE LIKELIHOOD FUNCTION FOR THE SP MODEL

This Appendix lists the first and second derivatives of the likelihood function for the SP Model of wage and strike offers discussed in chapter 4. The first and second derivatives for the likelihood function are used by the maximizing algorithms to compute iterative estimates, to examine convergence criteria, and to calculate the standard errors of the coefficients, as described in chapter 4. Throughout the Appendix it is assumed that the data are ordered so that the first $T_S$ observations correspond to contracts with a strike while the remaining $T - T_S$ observations correspond to contracts without a strike.

The likelihood function for the SP Model is given by:

$$
\ln L = \sum_{t=1}^{T_S} \ln G_t + \sum_{t=T_S+1}^{T} \ln H_t,
$$

for

$$
G_t = f_{1t}[1 - F_{2t}]f_{St},
$$
$$
H_t = F_{1t}f_{2t}[1 - F_{St}] + f_{Pt}F_{vt},
$$

and where

$$
f_{it} = \frac{1}{\sigma_i\sqrt{2\pi}} \exp \left[-\frac{(w_t - x_{it}\beta_i)^2}{2\sigma_i^2}\right], \quad \text{for } i = 1, 2, P,
$$
$$
f_{St} = \frac{1}{\sigma_S\sqrt{2\pi}} \exp \left[-\frac{(s_t - x_{St}\beta_S)^2}{2\sigma_S^2}\right],
$$
$$
F_{it} = \int_{-\infty}^{w_t-x_{it}\beta_i/\sigma_i} \phi(u) \, du, \quad \text{for } i = 1, 2,
$$
$$
F_{St} = \int_{-\infty}^{s_t-x_{St}\beta_S/\sigma_i} \phi(u) \, du,
$$
$$
F_{vt} = \int_{-\infty}^{(x_{1t}\beta_1-x_{2t}\beta_2)/\sigma} \phi(u) \, du,
$$

and where $\phi$ is the standard normal density function and $\sigma^2 = \sigma_1^2 + \sigma_2^2$. 

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Henceforth, to economize on notation, the subscript $t$ is dropped. Let $z_i = (w - x_i \beta_i)/\sigma_i$ for $i = 1, 2, P \text{ and } z_S = (s - x_S \beta_S)/\sigma_S$, where $x_i$ and $\beta_i$ are $K_i$ vectors for $i = 1, 2, P, S$. It follows that

$$f_i = \frac{1}{\sigma_i \sqrt{2\pi}} e^{-z_i^2/2}, \quad \text{for } i = 1, 2, P, S,$$

$$F_i = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{z_i} e^{-u^2/2} du, \quad \text{for } i = 1, 2, S,$$

and

$$F_v = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{v} e^{-u^2/2} du.$$

It is straightforward to show that

$$\frac{\partial f_i}{\partial \beta_i} = \frac{f_i z_i}{\sigma_i} x_i,$$

$$\frac{\partial f_i}{\partial \sigma_i^2} = \frac{f_i}{2\sigma_i^2} (z_i^2 - 1),$$

$$\frac{\partial F_i}{\partial \beta_i} = -f_i x_i,$$

and

$$\frac{\partial F_i}{\partial \sigma_i^2} = -\frac{f_i z_i}{2\sigma_i^2}, \quad \text{for } i = 1, 2, P, S.$$

Similarly,

$$\frac{\partial f_v}{\partial \beta_1} = -\frac{f_v v}{\sigma} x_1,$$

$$\frac{\partial f_v}{\partial \sigma_1^2} = \frac{f_v}{2\sigma^2} (v^2 - 1),$$

$$\frac{\partial F_v}{\partial \beta_1} = f_v x_1,$$

$$\frac{\partial F_v}{\partial \sigma_1^2} = -\frac{f_v v}{2\sigma^2},$$

$$\frac{\partial f_v}{\partial \beta_2} = \frac{f_v v}{\sigma} x_2,$$

$$\frac{\partial f_v}{\partial \sigma_2^2} = \frac{\partial f_v}{\partial \sigma_1^2}.$$
\[ \frac{\partial F_v}{\partial \beta_2} = -f_v x_2, \]

and

\[ \frac{\partial F_v}{\partial \sigma_v^2} = \frac{\partial F_v}{\partial \sigma_v^2}. \]

There are \( K = K_1 + K_2 + K_P + K_S + 4 \) parameters in the likelihood function, corresponding to the coefficients in each of the offer equations and the four variances, of which there is one for each equation. It follows that the first derivatives of the log likelihood are given by

\[
\frac{\partial \ln L}{\partial \omega_i} = \sum_{t=1}^{T_S} \frac{\partial \ln G_t}{\partial \omega_i} + \sum_{t=T_S+1}^{T} \frac{1}{H} \frac{\partial H}{\partial \omega_i}, \quad \text{for } i = 1, 2, \ldots, K, \tag{B.1}
\]

for \( \omega \in \Omega \), the set of parameters of the likelihood function \( L \).

Ignoring the \( t \) subscript, the first derivatives of \( \ln G \) are given by

\[
\frac{\partial \ln G}{\partial \beta_1} = \frac{z_1}{\sigma_1} x_1, \quad \frac{\partial \ln G}{\partial \sigma_1^2} = \frac{(z_1^2 - 1)}{2\sigma_1^2}, \\
\frac{\partial \ln G}{\partial \beta_2} = \frac{f_2}{(1 - F_2)} x_2, \quad \frac{\partial \ln G}{\partial \sigma_2^2} = \frac{f_2 z_2}{2\sigma_2(1 - F_2)}, \\
\frac{\partial \ln G}{\partial \beta_S} = \frac{z_S}{\sigma_S} x_S, \quad \frac{\partial \ln G}{\partial \sigma_S^2} = \frac{(z_S^2 - 1)}{2\sigma_S^2}.
\]

Clearly,

\[ \frac{\partial \ln G}{\partial \beta_P} = 0 \quad \text{and} \quad \frac{\partial \ln G}{\partial \sigma_P^2} = 0. \]

Again, dropping the \( t \) subscript, the first derivatives of \( H \) are given by

\[
\frac{\partial H}{\partial \beta_1} = [-f_1 f_2(1 - F_S)] + f_P f_v x_1, \\
\frac{\partial H}{\partial \beta_2} = \left[ \frac{F_1 f_2(1 - F_S) z_2}{\sigma_2} - f_P f_v \right] x_2,
\]

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\[
\begin{align*}
\frac{\partial H}{\partial \beta_P} &= \frac{f_P F_v z_P}{\sigma_P} x_P, \\
\frac{\partial H}{\partial \beta_S} &= F_1 f_2 f_s s, \\
\frac{\partial H}{\partial \sigma_1^2} &= \frac{f_1 f_2 (1 - F_S) z_1}{2\sigma_1} - \frac{f_P f_v v}{2\sigma}, \\
\frac{\partial H}{\partial \sigma_2^2} &= \frac{F_1 f_2 (1 - F_S) (z_2^2 - 1)}{2\sigma_2} - \frac{f_P f_v v}{2\sigma}, \\
\frac{\partial H}{\partial \sigma_P^2} &= \frac{f_P F_v (z_P^2 - 1)}{2\sigma_P^2}, \\
\frac{\partial H}{\partial \sigma_S^2} &= \frac{F_1 f_2 f_S z_S}{2\sigma_S}.
\end{align*}
\]

The first derivatives of the log likelihood are derived by making the appropriate substitutions from the first derivatives of \(\ln G\) and \(H\) into (B.1).

The \(K \times K\) matrix of second derivatives of the log likelihood, the Hessian, has as a typical element

\[
\frac{\partial^2 \ln L}{\partial \omega_i \partial \omega_j} = \sum_{t=1}^{T_S} \frac{\partial^2 \ln G_t}{\partial \omega_i \partial \omega_j} + \sum_{t=T_S+1}^{T} \frac{H_t (\partial^2 H_t/\partial \omega_i \partial \omega_j) - (\partial H_t/\partial \omega_i) (\partial H_t/\partial \omega_j)}{H_t^2}, \quad \text{for } i, j = 1, 2, \ldots, K.
\]

As \(\ln G_t\) and \(H_t\) are continuous functions, the Hessian is a symmetric matrix.
The second derivatives of $G$ are as follows

\[
\frac{\partial^2 \ln G}{\partial \beta_1 \partial \beta'_1} = -\frac{1}{\sigma_1^2} x_1 x'_1,
\]
\[
\frac{\partial^2 \ln G}{\partial \beta_1 \partial \sigma_1^2} = -\frac{z_1}{\sigma_1^3} x_1,
\]
\[
\frac{\partial^2 \ln G}{\partial \beta_2 \partial \beta'_2} = \left[ \frac{f_2 z_2}{\sigma_2 (1 - F_2)} - \frac{f_2^2}{(1 - F_2)^2} \right] x_2 x'_2,
\]
\[
\frac{\partial^2 \ln G}{\partial \beta_2 \partial \sigma_2^2} = \left[ \frac{f_2 (z_2^2 - 1)}{2\sigma_2^2 (1 - F_2)} - \frac{f_2^2 z_2}{2\sigma_2 (1 - F_2)^2} \right] x_2,
\]
\[
\frac{\partial^2 \ln G}{\partial \beta_S \partial \beta'_S} = -\frac{1}{\sigma_S^2} x_S x'_S,
\]
\[
\frac{\partial^2 \ln G}{\partial \beta_S \partial \sigma_S^2} = -\frac{z_S}{\sigma_S^3} x_S,
\]
\[
\frac{\partial^2 \ln G}{(\partial \sigma_1^2)^2} = \frac{1}{2\sigma_1^4},
\]
\[
\frac{\partial^2 \ln G}{(\partial \sigma_2^2)^2} = \frac{f_2 z_2 (z_2^3 - 3) - f_2^2 z_2^2}{2\sigma_2^2 (1 - F_2) - 4\sigma_2^2 (1 - F_2)^2},
\]
\[
\frac{\partial^2 \ln G}{(\partial \sigma_S^2)^2} = \frac{1}{2\sigma_S^4},
\]

and

\[
\frac{\partial^2 \ln G}{\partial \beta_i \partial \beta'_j} = 0 \quad i \neq j; \quad i, j = 1, 2, S,
\]
\[
\frac{\partial^2 \ln G}{\partial \beta_i \partial \sigma_j} = 0 \quad i \neq j; \quad i, j = 1, 2, S,
\]

while

\[
\frac{\partial^2 \ln G}{\partial \beta_i \partial \beta'_P} = 0, \quad \frac{\partial^2 \ln G}{\partial \beta_i \partial \sigma_P^2} = 0,
\]
\[
\frac{\partial \ln G}{\partial \sigma_i^2 \partial \beta'_P} = 0, \quad \frac{\partial \ln G}{\partial \sigma_i^2 \partial \sigma_P^2} = 0, \quad \text{for} \ i = 1, 2, P.
\]

The second derivatives of $H$ are given by

\[
\frac{\partial^2 H}{\partial \beta_1 \partial \beta'_1} = \left[ -\frac{f_1 f_2 (1 - F_{S}) z_1}{\sigma_1} - \frac{v f P f_v}{\sigma} \right] x'_1 x_1,
\]
\[
\frac{\partial^2 H}{\partial \beta_1 \partial \beta'_2} = \left[ -\frac{f_1 f_2 (1 - F_{S}) z_2}{\sigma_2} + \frac{v f P f_v}{\sigma} \right] x'_1 x_2,
\]
\[
\frac{\partial^2 H}{\partial \beta_1 \partial \beta'_P} = \frac{f_P f_v z_P}{\sigma_P} x'_1 x_P,
\]
\[
\frac{\partial^2 H}{\partial \beta_1 \partial \beta'_S} = -f_1 f_2 f_S x'_1 x_S,
\]
\[
\frac{\partial^2 H}{\partial \beta_1 \partial \sigma^2_1} = \left[ -\frac{f_1 f_2 (1 - F_S)(z_1^2 - 1)}{2\sigma_1^2} + \frac{f_P f_v (v^2)}{2\sigma^2} \right] x'_1,
\]
\[
\frac{\partial^2 H}{\partial \beta_2 \partial \beta'_1} = \left[ -\frac{f_1 f_2 (1 - F_S)(z_2^2 - 1)}{2\sigma_2^2} + \frac{f_P f_v (v^2)}{2\sigma^2} \right] x'_1,
\]
\[
\frac{\partial^2 H}{\partial \beta_1 \partial \sigma^2_2} = \frac{f_P f_v (z_P^2 - 1)}{2\sigma_P^2} x'_1,
\]
\[
\frac{\partial^2 H}{\partial \beta_2 \partial \sigma^2_2} = -\frac{f_1 f_2 f_S z_S}{2\sigma_S} x'_1,
\]
\[
\frac{\partial^2 H}{\partial \beta_2 \partial \beta'_2} = \left[ -\frac{f_1 f_2 (1 - F_S)(z_2^2 - 1)}{\sigma_2^2} - \frac{v f_P f_v}{\sigma} \right] x'_2 x_2,
\]
\[
\frac{\partial^2 H}{\partial \beta_2 \partial \beta'_P} = -\frac{f_P f_v z_P}{\sigma_P} x'_2 x_P,
\]
\[
\frac{\partial^2 H}{\partial \beta_2 \partial \beta'_S} = \frac{f_1 f_2 f_S z_S}{\sigma_2} x'_2 x_S,
\]
\[
\frac{\partial^2 H}{\partial \beta_2 \partial \sigma^2_1} = \left[ -\frac{f_1 f_2 (1 - F_S)(z_1 z_2)}{2\sigma_1 \sigma_2} - \frac{f_P f_v (v^2 - 1)}{2\sigma^2} \right] x'_2,
\]
\[
\frac{\partial^2 H}{\partial \beta_2 \partial \sigma^2_2} = \left[ -\frac{f_1 f_2 (1 - F_S)(z_2^2 - 3)}{2\sigma_2^3} - \frac{f_P f_v (v^2)}{2\sigma^2} \right] x'_2,
\]
\[
\frac{\partial^2 H}{\partial \beta_2 \partial \beta'_P} = -\frac{f_P f_v (z_P^2 - 1)}{2\sigma_P^2} x'_2,
\]
\[
\frac{\partial^2 H}{\partial \beta_2 \partial \beta'_S} = -\frac{f_1 f_2 f_S z_S}{2\sigma_2 \sigma_S} x'_2,
\]
\[
\frac{\partial^2 H}{\partial \beta_P \partial \beta'_P} = \frac{F_v f_P (z_P^2 - 1)}{\sigma_P^2} x'_p x_P,
\]
\[
\frac{\partial^2 H}{\partial \beta_P \partial \beta'_S} = 0,
\]
\[
\frac{\partial^2 H}{\partial \beta_P \partial \sigma_P^2} = -\frac{v f_P f_v z_P}{2 \sigma_P} x_P',
\]
\[
\frac{\partial^2 H}{\partial \beta_P \partial \sigma_P^2} = -\frac{v f_P f_v z_P}{2 \sigma_P} x_P',
\]
\[
\frac{\partial^2 H}{\partial \beta_P \partial \sigma_P^2} = \frac{F_v f_P z_P (z_P^2 - 3)}{2 \sigma_P^3} x_P',
\]
\[
\frac{\partial^2 H}{\partial \beta_P \partial \sigma_P^2} = 0,
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \beta_S'} = \frac{F_1 f_2 f_3 z_S}{\sigma_S} x_S' x_S,
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = \frac{f_1 f_2 f_3 z_1}{2 \sigma_1} x_S',
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = \frac{F_1 f_2 f_3 (z_2^2 - 1)}{2 \sigma_2^2} x_S',
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = 0,
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = \frac{F_1 f_2 f_3 (z_2^2 - 1)}{2 \sigma_2^2} x_S',
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = -\frac{v f_P f_v (v^2 - 3)}{4 \sigma_3^3},
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = -\frac{v f_P f_v (v^2 - 3)}{4 \sigma_3^3},
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = -\frac{F_1 f_2 f_3 z_S}{4 \sigma_1 \sigma_2},
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = -\frac{f_1 f_2 f_3 z_1 z_S}{4 \sigma_1 \sigma_2},
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = \frac{F_1 f_2 (z_2^4 - 6 z_2^2 + 3)}{4 \sigma_2^4} - \frac{v f_P f_v (v^2 - 3)}{4 \sigma_3^3},
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = \frac{v f_P f_v (v^2 - 3)}{4 \sigma_3^3},
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = \frac{F_1 f_2 f_3 (z_2^2 - 1) z_S}{4 \sigma_2^2 \sigma_S},
\]
\[
\frac{\partial^2 H}{\partial \beta_S \partial \sigma_S^2} = \frac{F_1 f_2 f_3 (z_2^2 - 1) z_S}{4 \sigma_2^2 \sigma_S},
\]

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\[
\frac{\partial^2 H}{(\partial \sigma_P^2)^2} = \frac{F_v f_P (z_P^2 - 6z_P^2 + 3)}{4\sigma_P^4},
\]
\[
\frac{\partial^2 H}{\partial \sigma_P^2 \partial \sigma_S^2} = 0,
\]
and
\[
\frac{\partial^2 H}{(\partial \sigma_S^2)^2} = \frac{F_1 f_2 f_S z_S (z_S^2 - 3)}{4\sigma_S^4},
\]

In the above expressions \(z_1, z_2, z_P, z_S, v, f_1, f_2, f_P, f_S, f_v, F_1, F_2, F_P, F_S\) and \(F_v\) are scalars; \(x_1\) is a \(K_1\) vector, \(x_2\) is a \(K_2\) vector, \(x_P\) is a \(K_P\) vector, and \(x_S\) is a \(K_S\) vector; \(x_1^T x_1\) is a \(K_1 \times K_1\) matrix, \(x_1^T x_2\) is a \(K_1 \times K_2\) matrix, \(x_1^T x_P\) is a \(K_1 \times K_P\) matrix, \(x_1^T x_S\) is a \(K_1 \times K_S\) matrix, \(x_2^T x_2\) is a \(K_2 \times K_2\) matrix, \(x_2^T x_P\) is a \(K_2 \times K_P\) matrix, \(x_2^T x_S\) is a \(K_2 \times K_S\) matrix, \(x_P^T x_P\) is a \(K_P \times K_P\) matrix, \(x_S^T x_S\) is a \(K_S \times K_S\) matrix, and \(x_P^T x_S\) is a \(K_P \times K_S\) matrix. The second derivatives of the log likelihood are derived by making the appropriate substitutions from the second derivatives of \(G\) and \(H\) into (B.2).