An Econometric Model of Trade Union Membership Growth in Canada, 1925-1966

by

Michael G. Abbott

Queen's University at Kingston and Princeton University

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1. INTRODUCTION

The sporadic and irregular character of the growth of trade union membership has long stimulated considerable debate, primarily among labor economists and economic historians, over the identity of the major causal factors underlying observed fluctuations in aggregate union membership growth and over the prospects for future expansion of the trade union movement. This debate has produced an extensive historical and institutional literature in which can be found numerous theories linking variations in the growth of union membership with various causal factors of an economic, political, social and legal-institutional nature. More recently, investigators have adopted econometric methods -- principally those of the linear regression model -- in attempts to test systematically some of the hypotheses advanced in the institutional literature concerning the determinants of union membership growth. These econometric investigations have differed
considerably with respect to the types of data they employ and the particular measures of unionization or union growth they attempt to explain. The present study belongs in the category of econometric studies that are based on aggregate time series data and that endeavor to model intertemporal variations in the rate of change of aggregate trade union membership.\textsuperscript{2} Several such studies for various countries have appeared over the past decade or so; most notable among these are the seminal analysis by Ashenfelter and Pencavel (1969) of American trade union growth over the period 1904-1960, and the monograph by Bain and Elsheikh (1976) in which a model of union growth is proposed and estimated on British data for the period 1892-1970.\textsuperscript{3}

Like its predecessors, the present study proposes a single equation linear regression model designed to represent movements in the annual rate of change of aggregate trade union membership in Canada. Several variants of the proposed model are estimated and tested on annual Canadian time series data for the period 1925-1966; in addition, the \textit{ex post} forecasting performance of the estimated models is investigated using observations for the years 1967-1972. The proposed model of Canadian union growth is something of a composite or hybrid model in the sense that its specification draws extensively on previous models advanced in earlier econometric studies of aggregate union growth. In particular, it incorporates many of the hypotheses respecting observable determinants of union growth that are embodied in the Ashenfelter-Pencavel (1969) model, the Bain-Elsheikh (1976) model, and the model proposed by Swidinsky (1974) to represent Canadian union membership growth over the period 1911-1970. Where the present model differs from these previous ones is in the empirical measurement and speci-
fication of many of the causal factors that are suggested by these hypothe-
ses. Furthermore, in contrast to earlier investigations, the current study
devotes considerable attention to assessing the empirical adequacy of the
proposed econometric model of Canadian union growth. Recent advances in
testing for different types of model misspecification and in testing the
specification of non-nested regression models are employed for this pur-
pose, in an effort to address systematically some of the empirical issues
raised in comments on previous models.4

In general terms, the empirical results obtained appear to confirm
the appropriateness of the postulated model in representing the sample data
for the estimation period 1925-1966, and to support most of the hypotheses
incorporated in the model respecting the determinants of intertemporal
variations in Canadian union membership growth. However, the analysis of
the model's ex post forecasting performance for the years 1967-1972 yields
somewhat mixed results, and suggests that the model may not provide a
fully satisfactory representation of Canadian union growth in the post-1966
period.

The remainder of the paper is organized as follows. Section 2
briefly describes the data set used in the study and the manner in which it
is employed in the empirical analysis. Section 3 outlines the specification
of the proposed model. Section 4 presents and interprets the estimates of
several variants of the model for the estimation period 1925-1966 and, for
comparative purposes, provides estimates of a modified version of the
Ashenfelter-Pencavel (1969) model. Section 5 reports the results of
several tests for misspecification, and Section 6 presents some tests of
model specification based on both nested and non-nested hypothesis testing.
procedures. Section 7 examines the forecasting accuracy of the models for the post-sample forecast period 1967-1972. Finally, Section 8 contains a summary of major empirical results and some suggestions for further research.

2. THE DATA

This section briefly outlines the data set on which the present investigation of Canadian union membership growth is based, and describes the manner in which these data are employed in the empirical analysis. Detailed descriptions of the definitions of all variables used in the analysis and of the sources from which they were compiled are given in the Data Appendix.

A continuous series of annual figures on aggregate trade union membership in Canada is available from the year 1911. Unfortunately, however, tolerably consistent annual time series for many of the explanatory variables included in the proposed model are available only as far back as 1921. Consequently, the full data set employed in the present study begins in 1921 and comprises annual time series observations for the period 1921-1972. Furthermore, because the dependent and several explanatory variables are expressed in annual percentage change form, and because too the proposed model specifies fixed or distributed lags on some of the regressors, complete joint observations are available only for the years 1925-1972, which therefore constitutes the complete sample period on which the empirical analysis is performed.
The present study's approach to the sample data reflects in part considerations respecting public policy towards labor relations and collective bargaining in Canada during the sample period 1925-1972. Several major legislative instruments of Canadian labor policy were in effect for portions of the complete sample period. The Industrial Disputes Investigation (IDI) Act of 1907 was the major piece of federal labor legislation in force during the early years of the sample period. It introduced the principles of work stoppage suspension and compulsory conciliation in dispute resolution and, in effect, "was designed to accomplish a forced recognition of a bargaining relationship, thereby avoiding the necessity of using the strike or lockout in recognition issues." In 1944, under its wartime emergency powers, the federal Cabinet proclaimed Order-in-Council P.C. 1003, also known as the Wartime Labour Relations Regulations. P.C. 1003 suspended the IDI Act and, in its place, established a national labor relations policy which combined the IDI Act principles of work stoppage suspension and compulsory conciliation in dispute resolution with the U.S. Wagner Act (1935) principles and administrative techniques of compulsory recognition and compulsory bargaining. The policy implemented by P.C. 1003 clearly enunciated for the first time the right of unions to exist and the right of employees to union representation, and put in place administrative agencies to enforce these rights. The Industrial Relations and Disputes Investigation (IRDI) Act of 1948 repealed the suspended IDI Act of 1907 and replaced the temporary wartime policy established by P.C. 1003 in 1944. Although it contained several minor innovations, the IRDI Act essentially retained all major features of the labor relations policy established by P.C. 1003, including most importantly the dual principles of compulsory
conciliation and work stoppage suspension on the one hand, and of compulsory recognition and bargaining on the other. Thus, the IRDI Act of 1948 represented no significant shift in policy but rather a postwar continuation of the labor relations framework implemented by P.C.1003 under the wartime emergency powers of the federal government. The last major development in Canadian labor legislation during the sample period occurred in March, 1967 with the enactment of the Public Service Staff Relations Act (PSSRA). This Act granted for the first time full collective bargaining rights to federal government employees, and effectively extended to federal civil servants most of the rights and procedures respecting representation that non-governmental employees had secured pursuant to P.C. 1003.

Although some federal employees were organized in one way or another prior to 1967, it was the stated view of the Public Service Alliance of Canada in April, 1971 that "the extent of their present organization and their acceptance by the employer are due mainly to the PSSRA."

In view of these major legislative landmarks in the development of labor relations policy in Canada, the complete sample period 1925-1972 is partitioned into three subperiods, each of which corresponds roughly to a major phase in the evolution of industrial relations policy in Canada. The first subperiod comprises the years 1925 to 1943 inclusive, when the Industrial Disputes Investigation Act (1907) was the dominant policy instrument in force. The second subperiod comprises the years 1944 to 1966 inclusive, the period during which the dual policy contained in P.C.1003 and its postwar successor, the Industrial Relations and Disputes Investigation Act, provided the basic legal and institutional framework for private sector unionism in Canada. The third subperiod includes the years 1967 to
1972 inclusive that immediately followed the introduction of collective bargaining for federal government employees under the provisions of the Public Service Staff Relations Act (1967).

Because the legal and administrative framework governing union organization and collective bargaining was substantially different in these subperiods, it is quite possible that any model of union membership growth may exhibit significant parameter differences across the three subperiods. In recognition of this possibility, the present investigation focuses mainly on the first two subperiods for purposes of formulating, estimating and testing several variants of a model of Canadian union membership growth; accordingly, the estimation period sample consists of the 42 annual observations for the years 1925-1966. The six observations for the third subperiod, 1967-1972, are reserved primarily to evaluate the ex post forecasting accuracy of the equations estimated for the 1925-1966 period. Thus, the scope of the present investigation is limited to modelling the growth of Canadian union membership in the era preceding widespread public sector unionism, when private sector industrial unionism emerged and matured to become an important and permanent institutional feature of the Canadian economy.
3. SPECIFICATION OF THE MODEL

Like previous econometric models of union growth, the present model employs the linear regression framework in an attempt to provide an empirical representation of intertemporal movements in aggregate union membership growth. The dependent variable is $\Delta T_t = ((T_t - T_{t-1})/T_t) \cdot 100$, where $\Delta T_t$ denotes the annual percentage rate of change in $T_t$, total Canadian union membership in year $t$. The regressor set comprises variables that are intended to represent the various determinants of $\Delta T_t$. Of the numerous factors that have been proposed as determinants of union growth, those identified with the business cycle theories propounded by John R. Commons (1918), Selig Perlman (1923) and Lloyd Ulman (1955a, 1955b), among others, have played a central role in both the institutional and econometric literature.\textsuperscript{12} Other factors that also have been incorporated in various ways in both institutional and econometric explanations of union growth include: the prevailing organizational and institutional environment; the attitudes and behavior of workers, unions and employers; and the sentiments and policy actions of government. Most of these factors cannot be measured directly, and therefore must be represented indirectly by observable proxy variables.\textsuperscript{13} This section identifies the various factors incorporated in the present model of Canadian union membership growth by specifying \textit{seriatim} both the measurable explanatory variables by which these factors are empirically represented and the hypotheses respecting the \textit{ceteris paribus} relationship of each explanatory variable to $\Delta T_t$. 
3.1 The Rate of Change of Unionizable Employment

An essential element of business cycle theories of union growth is the proposition that the expected returns to and costs of union membership, as perceived by workers, employers and unions, depends systematically on cyclical variations in aggregate labor market conditions and in real economic activity. But while this proposition has been recognized and incorporated in virtually all previous econometric models of union growth, there has emerged no consensus on the most appropriate measure for representing empirically intertemporal variations in aggregate labor market conditions. The two most commonly used measures of real labor market conditions in models of aggregate union growth are: (1) the rate of change of total employment in potentially unionizable sectors of the economy, originally proposed by Ashenfelter and Peacock (1969); and (2) the level of the aggregate unemployment rate (or changes in the aggregate unemployment rate), advocated by Bain and Elsheikh (1976).

Good a priori arguments can and have been advanced to justify the use of each of these empirical measures of aggregate labor market conditions. Basically, these arguments contend that several underlying determinants of union growth vary directly with the rate of change of unionizable employment and inversely with the aggregate unemployment rate. Included among these determinants are: the ability of unions to secure negotiated wage and non-wage improvements in the terms of employment, which constitutes one of the perceived benefits to workers of union membership; the opportunity costs to employers of actively resisting union organizing efforts, as measured by reduced or lost output and sales stemming from work
disruptions or stoppages; the ability of employers to pass on to customers the increased costs arising from unionization; and the net returns to, and hence the intensity of, unions' organizing and recruiting efforts. Although it is certainly possible to give counterexamples, these arguments are usually taken to imply that the rate of change of unionizable employment should have a positive effect, and the aggregate unemployment rate a negative effect, on the rate of change of union membership. Furthermore, because union formation and expansion is an inherently lengthy and costly process, and because unionized workers are generally reluctant to relinquish their membership when economic conditions worsen, a short finite distributed lag relationship is typically specified between union membership growth and either the rate of change of employment or the aggregate unemployment rate.

There are really two distinct (though related) issues involved in choosing between these two indicators of aggregate labor market conditions. The first is: which of the two measures is conceptually less ambiguous and empirically more reliable as an indicator of the cyclical performance of the aggregate labor market? The second is: which of the two measures is more closely related empirically to the rate of change of aggregate union membership?

With respect to the first issue, there appears to be no compelling reason for preferring on purely conceptual grounds one or the other of the two measures as "the" appropriate indicator of cyclical variations in labor market conditions. Neither is there a clearcut choice between the two on the basis of their empirical reliability and consistency as indicators of cyclical labor market conditions, although a superficially plausible case
can be made for perhaps preferring the percentage change in unionizable 
employment over the aggregate unemployment rate. With respect to the 
second issue — i.e., which of the two measures is the better predictor of 
union growth — the present investigation takes the position that the 
choice between the rate of change of unionizable employment and the aggre-
gate unemployment rate is essentially an empirical question that can only 
be decided on empirical grounds. Accordingly, in Section 6, both nested 
and non-nested hypothesis testing procedures are employed to test the spe-
cification of the proposed model of Canadian union growth when one or the 
other of the two measures is used to represent cyclical variations in real 
labor market conditions. In addition to yielding evidence on the question 
of which, if either, of the corresponding alternative models provides a 
valid representation of the Canadian sample data, these tests may also help 
to select empirically the variable that better explains Canadian union mem-
bership growth over the period 1925-1966. Partly in anticipation of these 
test results, the present model incorporates the effect on Canadian union 
membership growth of cyclical variations in real labor market conditions by 
specifying a four-year unrestricted distributed lag on

\[ \Delta E_{t-1} = \left( \frac{(E_t - E_{t-1})}{E_{t-1}} \right) \cdot 100, \]

the annual percentage change in non-agricultural paid employment, where \( E_t \) 
denotes total non-agricultural paid employment in year \( t \). More specifi-
cally, the regressor set is specified to include the values of \( \Delta E_{t-1} \) for 
\( i = 0, 1, 2, 3 \), on the expectation that both the individual lag coef-
ficients and the sum of the lag coefficients will be positively signed.
3.2 The Rate of Change of Money Prices

Like the rate of change of unionizable employment or the unemployment rate, the rate of change of money prices has been included in virtually all previous econometric models of union growth to serve as a separate indicator of business cycle influences on union membership growth. There are several reasons for expecting the rate of price inflation to exert a positive ceteris paribus effect on the rate of change of aggregate union membership. Many of these are encompassed by the threat effect, according to which workers are induced to become or to remain union members in order to defend their standard of living and to promote their real wage gain expectations in the face of rising consumer prices.\(^{16}\) In addition to the threat effect, higher rates of price inflation may also increase workers' uncertainty about future rates of change in prices and real wages, raise information costs about individual wage increases, and alter customary relative wage relationships — all of which may increase the attractiveness of union membership as a vehicle for expressing and protecting the economic interests of workers. Finally, insofar as rapidly rising prices have historically been associated with increased business prosperity, increases in the rate of price inflation may stimulate union growth by reducing employers' resistance to workers' demands for union recognition and collective bargaining.

The effect on Canadian union membership growth of the rate of change of money prices is represented in the present model by the variable

\[ \Delta P_t = \frac{(P_t - P_{t-1})}{P_t} \cdot 100, \]

the annual percentage change in the Consumer Price Index, where \( P_t \) denotes the all-items Consumer Price Index in year \( t \).
Furthermore, both the role of unions as defensive organizations and the observed relationship between the rate of price inflation and the average relative wage advantage of unionized workers suggest that the ceteris paribus effect of \( \Delta P_t \) on \( \Delta T_t \) may itself depend on the value of \( \Delta P_t \). Accordingly, to allow for possible nonlinearities in the effect of \( \Delta P_t \) on \( \Delta T_t \), both the first and second powers of \( \Delta P_t \) are included as regressors in the estimating equations. It is not obvious a priori what sign the coefficient of \( \Delta P_t^2 \) should be expected to take. All that can be predicted with reasonable certainty is that the coefficient of \( \Delta P_t \) should be positive.

3.3 The Density of Union Membership

Union density represents the extent to which the population of potential union members is already unionized; it is usually measured by the ratio of total union membership to total unionizable employment. Although existing union density is frequently viewed as a constraint on current increases in union membership, there are two alternative hypotheses concerning the direction of the effect of union density on union growth.

According to the so-called saturation hypothesis, the difficulty of further extending union membership increases as the proportion of potentially organizeable employment that is already unionized rises, partly because there are fewer unorganized workers left to recruit, partly because the remaining unorganized workers are less willing to join unions, and partly because the costs of recruiting those unorganized workers who are left tend to rise. In short, for a variety of reasons, there are thought to exist diminishing
returns to unions' organizing efforts.\textsuperscript{18} The saturation hypothesis thus implies that union density will have a negative \textit{ceteris paribus} effect on the rate of growth of union membership. However, there are also reasons for expecting union density to have a positive effect on union membership growth. Most of these relate to the growing acceptance of unionization by both employers and employees and the increasing social/political pressures to accept unionization that accompany the spread and maturation of unionism and collective bargaining. Contrary to the saturation hypothesis, these acceptance and demonstration effects imply a positive effect of union density on membership growth.

As in previous models, the effect of existing union density on Canadian union membership growth is incorporated in the present model by including in the regressor set the variable \( D_{t-1} = (T_{t-1}/E_{t-1}) \cdot 100 \), the one-period lagged value of total union membership \((T_{t-1})\) as a percentage of the one-period lagged value of total unionizable employment \((E_{t-1})\). Further, following the original example of Ashenfelter and Pencavel (1969), two variants of the relationship between \(\Delta T_t \) and \( D_{t-1} \) are specified: first, a linear variant, which specifies the density regressor as simply \( D_{t-1} \); and second, a nonlinear variant, which specifies the density regressor as \( D_{t-1}^{-1} \), the reciprocal of the percentage of potentially unionizable employment that is already organized. The saturation hypothesis implies a negative sign for the coefficient of \( D_{t-1}^{-1} \) and a positive sign for the coefficient of \( D_{t-1} \); conversely, the acceptance/demonstration hypothesis implies a positive sign for the coefficient of \( D_{t-1} \) and a negative sign for the coefficient of \( D_{t-1}^{-1} \).
3.4 Worker Discontent

An often-cited reason for individuals' deciding to join or form a union is to give voice to and to seek redress of workers' grievances with existing terms or conditions of employment. Thus, one of the recognized functions of unions is to serve as "agencies of protest and dissatisfaction"—i.e., as vehicles for the expression and redress of individual and collective worker grievances. Obviously, the existence and extent of worker discontent is not directly measurable, a fact which probably explains why few previous studies have attempted explicitly to incorporate the role of worker discontent in econometric models of union growth: in order even to account crudely for worker discontent in such models, it is necessary to adopt strong (some would say heroic) assumptions about the observable correlates of the existence and intensity of worker dissatisfaction.

One attempt to incorporate the effect of worker discontent on union growth is the Ashenfelter-Pencavel (1969) model of American union membership growth. Ashenfelter and Pencavel represent the current stock of worker grievances by the step function $\lambda^{(t-\delta)}u^P_t$, where $u^P_t$ is the "peak" value of the unemployment rate in the year of the most recent business cycle trough, $t$ denotes the current year, $\delta$ denotes the year of the most recent cyclical trough, and $\lambda$ is a decay parameter that reflects the rate of decline of workers' grievances and which therefore is constrained to take values in the closed unit interval $[0, 1]$. Implicit in this specification is the assumption that the stock of worker grievances depends directly on the unemployment rate in the most recent cyclical trough, not
on experience during the current cycle.

An alternative hypothesis explored here is that worker discontent varies directly with the severity of business cycle contractions and, further, that the response of union membership growth to workers' current grievances occurs during, rather than after, the contractionary period in which they arise. One proxy measure of worker discontent suggested by this hypothesis is the variable \( DD_t(U_t-U_t^*) \), where \( DD_t \) is a binary dummy variable that takes the value 1 in years of cyclical contraction and the value 0 in years of cyclical expansion, \( U_t \) is the aggregate civilian unemployment rate in year \( t \), and \( U_t^* \) is the unemployment rate that prevailed in the year of the preceding business cycle peak. In addition to the assumption that the effects of worker discontent on union growth occur without any appreciable lag, the proxy variable \( DD_t(U_t-U_t^*) \) also assumes (i) that worker grievances are essentially a phenomenon of cyclical recessions, and (ii) that the number and intensity of these grievances vary directly with the severity of the current recession, as measured by the magnitude of the difference between the current unemployment rate \( (U_t) \) and the unemployment rate at the end of the previous cyclical expansion \( (U_t^*) \). Thus, the variable \( DD_t(U_t-U_t^*) \) is included in the regressor set of the present model as a proxy measure of worker discontent, although it can more literally be interpreted as simply representing the effect on union growth of the occurrence and severity of business cycle contractions. In accordance with either interpretation, the coefficient of \( DD_t(U_t-U_t^*) \) is expected to be positive.
3.5 The Rate of Change of American Union Membership

Even a casual inspection of the historical data on aggregate trade union membership in Canada and the United States reveals that the patterns of aggregate union growth in the two countries have been broadly similar throughout much of the 20th century, although, as Swidinsky (1974, p. 437) observes, "there are significant differences in timing and in magnitudes of expansion and contraction." In view of the relative sizes of the two economies, one immediate possibility suggested by these similarities is that the growth of American union membership may have systematically influenced the growth of union membership in Canada. The sources of such an influence are to be found not only in the important role of the international — i.e., American-based — union in Canadian union development, but also in the prevalence of American ownership and/or control of Canadian enterprises, the high degree of economic interdependence between the two countries, and the dominant influence of the American media in shaping Canadian attitudes and expectations.21 The effect of American union growth on Canadian union growth is incorporated in the present model by including as a regressor the variable $\Delta T_{US,t} = \{(T_{US,t} - T_{US,t-1})/T_{US,t}\} \cdot 100$, the percentage change in American union membership, where $T_{US,t}$ is total membership in American unions exclusive of Canadian membership in U.S.-based international unions. The coefficient associated with $\Delta T_{US,t}$ is expected to have a positive sign, in accordance with the hypothesis that increases (decreases) in the growth rate of American union membership will generate, via the manifold economic and institutional linkages between the two countries, increases (decreases) in the growth rate of union membership in Canada.22
3.6. Union Militancy

The notion of union militancy or aggressiveness to date has received little attention in econometric models of union growth. One reason for this inattention is doubtless the fact that there exists no satisfactory proxy measure of union militancy -- in particular, a measure that is distinct from various numerical indicators of the potential strength of unions, such as union density or the rate of change of union membership. If, however, union militancy is interpreted as the extent to which unions actually exercise their bargaining power in pursuit of their members' interests, a case can be made for using same measure of changes in strike activity as an index of variations in union militancy. Just such a case is made by Swidinsky (1974), who argues that one observable consequence of increased union militancy is increased strike activity arising from more aggressive recruiting and bargaining on the part of unions. Without claiming to have identified "the" correct proxy measure of changes in union militancy, the present study follows Swidinsky's (1974) lead and includes as a regressor the variable $\Delta S_t = [(S_t - S_{t-1})/S_t] \times 100$, the percentage change in $S_t$, the total number of strikes beginning in year $t$.

One potential econometric problem with using $\Delta S_t$ as a predictor of $\Delta T_t$ is that, since increased union membership growth and increased strike frequency may reasonably be viewed as concomitant manifestations of more aggressive union behavior in organizing and/or bargaining, $\Delta T_t$ and $\Delta S_t$ might well be jointly determined. Consequently, to avoid potential simultaneity problems, the present model was initially specified to include only recent lagged values of $\Delta S$ -- in particular, the values of $\Delta S_{t-1}$ for
i = 1, 2, 3. But, as reported in Section 4, preliminary estimates of the model consistently failed to produce coefficient estimates for either \( \Delta S_{t-1} \) or \( \Delta S_{t-3} \) that even remotely approached statistical significance. Thus, in the interests of parsimony, the model is specified to include only \( \Delta S_{t-2} \), the two-year lagged value of the percentage change in aggregate strike starts, as a means of indirectly testing the union militancy hypothesis.

A second problem is essentially one of interpretation, and involves predicting a priori the sign of the coefficient of \( \Delta S_{t-2} \). Insofar as changes in aggregate strike starts reflect changes in union militancy, \( \Delta S_{t-2} \) may be expected to have a positive effect on \( \Delta T_t \). Alternatively, to the extent that increases (decreases) in strike starts are seen by workers and employers as indicating increases (decreases) in a component of the costs of unionization, \( \Delta S_{t-2} \) would be expected to have a negative effect on \( \Delta T_t \). These two alternative interpretations of the lagged change in strike starts imply that, strictly speaking, the sign of the coefficient of \( \Delta S_{t-2} \) is ambiguous a priori.

3.7 Major Legislative Enactments

Much of the institutional literature on the development of trade unionism emphasizes the role of legislation and prevailing public attitudes respecting union recognition and collective bargaining in stimulating or retarding union membership growth. Consequently, some previous econometric investigations of union growth, such as Ashenfelter and Pencavel's (1969) for the U.S., have attempted to account for the impact on union growth of governmental attitudes and actions towards union organization, although
others, such as Bain and Elsheikh's (1976) for the U.K., have found it unnecessary to do so. On balance, however, there appears to exist ample evidence to support the view that government action "to promote union recognition has had a profound effect on the growth of trade unions by tending to neutralize or at least contain employer opposition to them" (Bain (1981, p. 21)).

But while there may be good reasons for postulating effects of government policies on union growth, there is no consensus on how precisely to allow for these effects in econometric models of union membership growth. The present investigation adopts the expedient, if not altogether satisfactory, method of using a dummy variable to allow for the effect on Canadian union membership growth of the Federal Cabinet's proclamation in 1944 of Order-in-Council P.C. 1003. This dummy variable is denoted as $D_{44t}$, and is defined to equal one for all years from 1944 and zero for all years before 1944. Since P.C. 1003 in effect reduced the expected costs and risks to workers of forming or joining a union, it is expected to have increased the growth rate of Canadian union membership in the period following its enactment; accordingly, the coefficient of $D_{44t}$ is predicted to be positive.

The dummy variable $D_{44t}$ also enters the estimating equations interactively with the density variable $D_{t-1}$, or its reciprocal, $D_{t-1}^{-1}$. The resulting interaction variables $D_{44t} \cdot D_{t-1}$ and $D_{44t} \cdot D_{t-1}^{-1}$ provide a test of the hypothesis that the provisions of P.C. 1003 respecting union recognition may have altered the effect of existing union density on the current rate of union membership growth. Specifically, the hypothesis to be tested is that the value of $\Delta T_t / \Delta D_{t-1}$ prior to 1944 is less than the value of
$\Delta T_t/\Delta D_{t-1}$ from 1944 on. This hypothesis implies a positive coefficient for the regressor $D44_t \cdot D_{t-1}$, and a negative coefficient for the regressor $D44_t \cdot D_{t-1}^{-1}$.

3.8 Summary: The Estimating Equations

The arguments and hypotheses presented in this section may be summarized in terms of the following pair of estimating equations:

(1A) \[
\Delta T_t = \alpha_0 + \alpha_1 \Delta P_t + \alpha_2 \Delta P_t^2 + \sum_{i=0}^{i=3} \alpha_{3i} \Delta E_{t-i} + \alpha_4 D_{t-1} \\
+ \alpha_5 D44_t (U_t - U^*_t) + \alpha_6 \Delta T_{US,t} + \alpha_7 \Delta S_{t-2} + \alpha_8 D44_t \\
+ \alpha_9 D44_t \cdot D_{t-1} + u_{A,t} ;
\]

(1B) \[
\Delta T_t = \beta_0 + \beta_1 \Delta P_t + \beta_2 \Delta P_t^2 + \sum_{i=0}^{i=3} \beta_{3i} \Delta E_{t-i} + \beta_4 D_{t-1}^{-1} \\
+ \beta_5 D44_t (U_t - U^*_t) + \beta_6 \Delta T_{US,t} + \beta_7 \Delta S_{t-2} + \beta_8 D44_t \\
+ \beta_9 D44_t \cdot D_{t-1}^{-1} + u_{B,t} .
\]

The $\alpha$'s and $\beta$'s are (constant) regression parameters, and $u_{A,t}$ and $u_{B,t}$ are random error terms, each of which is assumed to be NID with zero mean and constant variance. Equation (1A), which incorporates the linear form of the union density effect, is referred to as the linear variant of the
model; equation (1B), which incorporates the reciprocal specification of the union density effect, is referred to as the nonlinear variant of the model. In accordance with the foregoing discussion, the following sign hypotheses are postulated for the regression coefficients of equations (1A) and (1B), respectively:

(A) $a_1 > 0, \ a_2 < 0, \ a_{3i} > 0 \ (i=0,1,2,3), \ a_4 \leq 0,$

$\ a_5 > 0, \ a_6 > 0, \ a_7 \leq 0, \ a_8 > 0, \ a_9 > 0; \$

(B) $\beta_1 > 0, \ \beta_2 < 0, \ \beta_{3i} > 0 \ (i=0,1,2,3), \ \beta_4 \leq 0,$

$\beta_5 > 0, \ \beta_6 > 0, \ \beta_7 \leq 0, \ \beta_8 > 0, \ \beta_9 < 0.$
4. ESTIMATES OF THE MODEL

This section presents and evaluates the OLS parameter estimates for the period 1925-1966 of two variants of the proposed model of Canadian union membership growth — the linear variant given by equation (1A), and the nonlinear variant given by equation (1B). In addition, estimates are reported for two restricted versions of equations (1A) and (1B): one restricted version, given by equations (1A.1) and (1B.1), sets equal to zero the coefficient on the interaction terms \(D_{44t} \cdot D_{t-1}\) and \(D_{44t} \cdot D_{t-1}^{-1}\), respectively; the second, given by equations (1A.2) and (1B.2), constrains to zero the coefficient on the intercept shift dummy variable \(D_{44t}\). These restricted versions of the two basic estimating equations correspond to specific hypotheses concerning the impact on \(\Delta T_t\) of the federal labor relations policy established by P.C. 1003 in 1944. Finally, in order to provide an empirical basis of comparison for the model herein proposed, a modified version of the Ashenfelter-Pencavel (1969) model is formulated and estimated on the same Canadian time series data for 1925-1966 to which the present model is fitted.

OLS estimates and summary statistics for the estimation period 1925-1966 are given in Table 1 for the linear variant, equation (1A), and in Table 2 for the nonlinear variant, equation (1B). In general, the estimated equations provide a good fit to the sample data for the years 1925-1966: the values of the coefficient of determination \(R^2\) are all in the neighborhood of 0.90, and the sample values of the F-statistic for testing the joint significance of the slope parameters are uniformly far in excess of their respective 1 percent critical values. Moreover, a
substantial majority of the regression parameters are estimated with a high degree of precision.28

The estimates of all equations in Tables 1 and 2 strongly support several of the hypotheses proposed in the previous section. The estimated distributed lag coefficients on $\Delta E_{t-1}$ substantially confirm the hypothesis that current and recent past rates of change in unionizable employment exert a positive effect on the rate of change of union membership. All the parameter estimates for $\Delta E_t$, $\Delta E_{t-1}$ and $\Delta E_{t-2}$ have the expected positive sign, and all are statistically significant on a two-tailed test at the 1 percent level. However, contrary to expectations, the parameter estimates for $\Delta E_{t-3}$ are all negatively signed and, though numerically small, are all significant at the 1 percent or 5 percent significance level. Nevertheless, the sums of the estimated lag coefficients on $\Delta E_{t-i}$ over $i = 0,1,2,3$ are uniformly large and positive, and are easily significant at the 1 percent test level: they imply that the long-run response (over a four year period) to a 1 percentage point increase in the annual rate of change of unionizable employment is equal to an increase of approximately 2.4 percentage points in the annual growth rate of Canadian union membership.

Finally, attempts to constrain the shape or form of the distributed lag on $\Delta E_{t-1}$ by imposing Almon polynomial restrictions consistently indicated that such restrictions were inconsistent with the sample data.29 Consequently, given the lag length of four years ($i=0,1,2,3$), only an unrestricted distributed lag relationship between $\Delta T_t$ and $\Delta E_{t-1}$ appears to be consistent with the estimation period sample data.30
TABLE 1

Ordinary Least Squares Estimates of Equation (1A):
Estimation Period 1925-1966

Dependent Variable = $\Delta T_t$
(t-ratios in parentheses)

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Equation (1A)</th>
<th>Equation (1A.1)</th>
<th>Equation (1A.2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>5.005</td>
<td>5.505</td>
<td>9.322</td>
</tr>
<tr>
<td></td>
<td>(1.456)</td>
<td>(2.548)</td>
<td>(2.995)</td>
</tr>
<tr>
<td>$\Delta P_t$</td>
<td>0.692</td>
<td>0.699</td>
<td>0.795</td>
</tr>
<tr>
<td></td>
<td>(4.536)</td>
<td>(4.811)</td>
<td>(5.079)</td>
</tr>
<tr>
<td>$\Delta P_t^2$</td>
<td>-0.0367</td>
<td>-0.0367</td>
<td>-0.0351</td>
</tr>
<tr>
<td></td>
<td>(-2.866)</td>
<td>(-2.916)</td>
<td>(-2.560)</td>
</tr>
<tr>
<td>$\Delta E_t$</td>
<td>1.441</td>
<td>1.441</td>
<td>1.393</td>
</tr>
<tr>
<td></td>
<td>(9.719)</td>
<td>(9.877)</td>
<td>(8.848)</td>
</tr>
<tr>
<td>$\Delta E_t^{-1}$</td>
<td>0.484</td>
<td>0.484</td>
<td>0.467</td>
</tr>
<tr>
<td></td>
<td>(4.395)</td>
<td>(4.465)</td>
<td>(3.960)</td>
</tr>
<tr>
<td>$\Delta E_t^{-2}$</td>
<td>0.851</td>
<td>0.853</td>
<td>0.863</td>
</tr>
<tr>
<td></td>
<td>(7.746)</td>
<td>(7.933)</td>
<td>(7.335)</td>
</tr>
<tr>
<td>$\Delta E_t^{-3}$</td>
<td>-0.352</td>
<td>-0.349</td>
<td>-0.280</td>
</tr>
<tr>
<td></td>
<td>(-3.034)</td>
<td>(-3.087)</td>
<td>(-2.332)</td>
</tr>
<tr>
<td>$D_{t-1}$</td>
<td>-0.801</td>
<td>-0.837</td>
<td>-1.091</td>
</tr>
<tr>
<td></td>
<td>(-3.334)</td>
<td>(-5.834)</td>
<td>(-4.936)</td>
</tr>
<tr>
<td>$DD_{t}(U - U^*)_{t}$</td>
<td>1.855</td>
<td>1.863</td>
<td>1.876</td>
</tr>
<tr>
<td></td>
<td>(8.167)</td>
<td>(8.479)</td>
<td>(7.713)</td>
</tr>
<tr>
<td>$\Delta T_{US, t}$</td>
<td>0.274</td>
<td>0.277</td>
<td>0.293</td>
</tr>
<tr>
<td></td>
<td>(4.782)</td>
<td>(5.199)</td>
<td>(4.823)</td>
</tr>
<tr>
<td>$\Delta S_{t-2}$</td>
<td>-0.0435</td>
<td>-0.0434</td>
<td>-0.0421</td>
</tr>
<tr>
<td></td>
<td>(-3.417)</td>
<td>(-3.468)</td>
<td>(-3.090)</td>
</tr>
<tr>
<td>$D_{44_t}$</td>
<td>15.690</td>
<td>14.525</td>
<td>--</td>
</tr>
<tr>
<td></td>
<td>(2.341)</td>
<td>(5.691)</td>
<td>--</td>
</tr>
<tr>
<td>$D_{44 \ast D}_{t-1}$</td>
<td>-0.0568</td>
<td>--</td>
<td>0.594</td>
</tr>
<tr>
<td></td>
<td>(-0.189)</td>
<td></td>
<td>(4.747)</td>
</tr>
<tr>
<td>$\sum_{i=3}^{i=a} \Delta E_{t-i}$</td>
<td>2.424</td>
<td>2.429</td>
<td>2.443</td>
</tr>
<tr>
<td></td>
<td>(9.921)</td>
<td>(10.145)</td>
<td>(9.333)</td>
</tr>
</tbody>
</table>
TABLE 1 (Continued)

Ordinary Least Squares Estimates of Equation (1A):
Estimation Period 1925-1966

Dependent Variable = $\Delta T_t$
(t-ratios in parentheses)

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Equation (1A)</th>
<th>Equation (1A.1)</th>
<th>Equation (1A.2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Summary Statistics</td>
<td>$b$/</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\ln L$</td>
<td>-86.7111</td>
<td>-86.7369</td>
<td>-90.3461</td>
</tr>
<tr>
<td>SSR</td>
<td>152.764</td>
<td>152.951</td>
<td>181.633</td>
</tr>
<tr>
<td>$s$</td>
<td>2.295</td>
<td>2.258</td>
<td>2.461</td>
</tr>
<tr>
<td>$R^2$</td>
<td>0.912</td>
<td>0.912</td>
<td>0.895</td>
</tr>
<tr>
<td>$F(k-1,T-k)$</td>
<td>25.066(12,29)</td>
<td>28.250(11,30)</td>
<td>23.359(11,30)</td>
</tr>
<tr>
<td>$d_1$</td>
<td>2.21</td>
<td>2.22</td>
<td>2.03</td>
</tr>
</tbody>
</table>

Notes: The number of observations in the estimation period sample is 42. Definitions and sources of the variables are given in the Data Appendix.

$^a/\sum_{i=3}^{3} \Delta E_{t-i}$ denotes the sum of the estimated coefficients on $\Delta E_{t-i}$ for $i = 0, \ldots, 3$; the coefficients on $\Delta E_{t-i}$ are estimated without restrictions.

$^b/$ For each equation: $\ln L$ is the maximized log-likelihood value; SSR is the sum of squared OLS residuals; $s$ is the standard error of the regression computed as $s = \left[ SSR/(T-k) \right]^{1/2}$, where $T = 42$ is the number of observations and $k$ is the number of regression parameters; $R^2$ is the coefficient of determination; $F(k-1,T-k)$ is the conventional $F$-statistic for testing the null hypothesis that all $(k-1)$ slope coefficients are jointly equal to zero; and $d_1$ is the standard Durbin-Watson statistic for testing the null hypothesis that the error terms are non-autoregressive against the alternative hypothesis that the error terms are first-order autoregressive.
TABLE 2

Ordinary Least Squares Estimates of Equation (1B):
Estimation Period 1925-1966

Dependent Variable = ΔT  
(t-ratios in parentheses)

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Equation (1B)</th>
<th>Equation (1B.1)</th>
<th>Equation (1B.2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-19.240</td>
<td>-23.726</td>
<td>-23.139</td>
</tr>
<tr>
<td></td>
<td>(-4.496)</td>
<td>(-5.013)</td>
<td>(-6.734)</td>
</tr>
<tr>
<td>ΔP_t</td>
<td>0.699</td>
<td>0.809</td>
<td>0.740</td>
</tr>
<tr>
<td></td>
<td>(4.618)</td>
<td>(4.701)</td>
<td>(4.878)</td>
</tr>
<tr>
<td>ΔP_t^2</td>
<td>-0.0353</td>
<td>-0.0332</td>
<td>-0.0355</td>
</tr>
<tr>
<td></td>
<td>(-2.774)</td>
<td>(-2.247)</td>
<td>(-2.735)</td>
</tr>
<tr>
<td>ΔE_t</td>
<td>1.443</td>
<td>1.368</td>
<td>1.434</td>
</tr>
<tr>
<td></td>
<td>(9.755)</td>
<td>(8.046)</td>
<td>(9.520)</td>
</tr>
<tr>
<td>ΔE_t-1</td>
<td>0.472</td>
<td>0.445</td>
<td>0.467</td>
</tr>
<tr>
<td></td>
<td>(4.308)</td>
<td>(3.503)</td>
<td>(4.181)</td>
</tr>
<tr>
<td>ΔE_t-2</td>
<td>0.830</td>
<td>0.834</td>
<td>0.838</td>
</tr>
<tr>
<td></td>
<td>(7.626)</td>
<td>(6.594)</td>
<td>(7.563)</td>
</tr>
<tr>
<td>ΔE_t-3</td>
<td>-0.392</td>
<td>-0.287</td>
<td>-0.370</td>
</tr>
<tr>
<td></td>
<td>(-3.344)</td>
<td>(-2.180)</td>
<td>(-3.119)</td>
</tr>
<tr>
<td>D_{t-1}^1</td>
<td>178.827</td>
<td>242.083</td>
<td>230.097</td>
</tr>
<tr>
<td></td>
<td>(3.247)</td>
<td>(4.018)</td>
<td>(5.273)</td>
</tr>
<tr>
<td>DD_{t} (U - U*)_{t}</td>
<td>1.868</td>
<td>1.894</td>
<td>1.931</td>
</tr>
<tr>
<td></td>
<td>(8.158)</td>
<td>(7.122)</td>
<td>(8.416)</td>
</tr>
<tr>
<td>ΔT_{US,t}</td>
<td>0.279</td>
<td>0.308</td>
<td>0.309</td>
</tr>
<tr>
<td></td>
<td>(4.798)</td>
<td>(4.608)</td>
<td>(5.547)</td>
</tr>
<tr>
<td>ΔS_{t-2}</td>
<td>-0.0443</td>
<td>-0.0433</td>
<td>-0.0439</td>
</tr>
<tr>
<td></td>
<td>(-3.497)</td>
<td>(-2.936)</td>
<td>(-3.400)</td>
</tr>
<tr>
<td>D44_t</td>
<td>-9.018</td>
<td>10.227</td>
<td>--</td>
</tr>
<tr>
<td></td>
<td>(-1.479)</td>
<td>(3.902)</td>
<td>--</td>
</tr>
<tr>
<td>D44_{t} * D_{t-1}</td>
<td>515.551</td>
<td>--</td>
<td>307.002</td>
</tr>
<tr>
<td></td>
<td>(3.397)</td>
<td></td>
<td>(5.365)</td>
</tr>
<tr>
<td>Σ_{i=0}^{i=3} ΔE_{t-i}</td>
<td>2.353</td>
<td>2.360</td>
<td>2.369</td>
</tr>
<tr>
<td></td>
<td>(9.731)</td>
<td>(8.400)</td>
<td>(9.625)</td>
</tr>
</tbody>
</table>
TABLE 2 (Continued)

Ordinart Least Squares Estimates of Equation (1B):

Estimation Period 1925-1966

Dependent Variable = \Delta T

(t-ratios in parentheses)

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Equation (1B)</th>
<th>Equation (1B.1)</th>
<th>Equation (1B.2)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Summary Statistics</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>%nL</td>
<td>-86.5846</td>
<td>-93.6179</td>
<td>-88.1113</td>
</tr>
<tr>
<td>SSR</td>
<td>151.846</td>
<td>212.255</td>
<td>163.297</td>
</tr>
<tr>
<td>s</td>
<td>2.288</td>
<td>2.660</td>
<td>2.333</td>
</tr>
<tr>
<td>(R^2)</td>
<td>0.913</td>
<td>0.878</td>
<td>0.906</td>
</tr>
<tr>
<td>(F(k-1,T-k))</td>
<td>25.233(12,29)</td>
<td>19.595(11,30)</td>
<td>26.288(11,30)</td>
</tr>
<tr>
<td>(d_1)</td>
<td>2.28</td>
<td>1.96</td>
<td>2.25</td>
</tr>
</tbody>
</table>

Notes: See the notes to Table 1.
The coefficient estimates for $\Delta P_t$ and $\Delta P_t^2$ in Tables 1 and 2 imply a positive and downwardly concave quadratic relationship of $\Delta T_t$ with $\Delta P_t$. The parameter estimates for $\Delta P_t$ are uniformly positive and statistically significant at the 1 percent test level on a two-tailed test, while those for $\Delta P_t^2$ are all negative and small, but statistically different from zero at either the 5 percent or 1 percent significance level. The implied values of $\partial \Delta T_t / \partial \Delta P_t$ for the entire estimation period 1925-1966, and for the sub-periods 1925-1943 and 1944-1966, evaluated at the respective sample mean values of $\Delta P_t$, are tabulated in Table 3 for each of the six estimating equations. These mean estimates of $\partial \Delta T_t / \partial \Delta P_t$ range from about 0.5 to 0.8 and are all significantly different from zero at the 1 percent test level. The relationship between $\partial \Delta T_t / \partial \Delta P_t$ is negative linear. Numerically, $\partial \Delta T_t / \partial \Delta P_t$ is positive for values of $\Delta P_t$ less than about 9.5 percent; statistically, the point estimates of $\partial \Delta T_t / \partial \Delta P_t$ are significantly different from zero at test levels of 5 percent or less for all values of $\Delta P_t$ less than about 5.0 percent. Thus, since only 4 of the 42 sample values of $\Delta P_t$ exceed 5.0 percent during the 1925-1966 estimation period, the coefficient estimates for $\Delta P_t$ and $\Delta P_t^2$ are consistent with a significant positive effect of $\Delta P_t$ on $\Delta T_t$ for all years except those characterized by historically high rates of price inflation.  

With respect to the relationship between $\Delta T_t$ and $D_{t-1}$, the parameter estimates in Table 1 and 2 provide strong confirmation of the saturation hypothesis, according to which existing union density is expected to have a negative effect on the rate of union membership growth. For both the 1925-1943 and 1944-1966 sub-periods, all coefficient estimates for $D_{t-1}$ in Table 1 are negative and significantly different from zero at the 1 percent
### TABLE 3

Estimates of the Marginal Effects on $\Delta T$ of $\Delta P$ and $\Delta D_{-1}$ in Equations (1A) and (1B)

(t-ratios in parentheses)

<table>
<thead>
<tr>
<th>Equation</th>
<th>Estimated Values of $\partial \Delta T / \partial \Delta P$ for:</th>
<th>Estimated Values of $\partial \Delta T / \partial \Delta D_{-1}$ for:</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1)</td>
<td>$\Delta P = 1.483$</td>
<td>$\Delta P = -0.091$</td>
</tr>
<tr>
<td>(1A)</td>
<td>0.583</td>
<td>0.699</td>
</tr>
<tr>
<td></td>
<td>(4.081)</td>
<td>(4.553)</td>
</tr>
<tr>
<td>(1A.1)</td>
<td>0.590</td>
<td>0.706</td>
</tr>
<tr>
<td></td>
<td>(4.352)</td>
<td>(4.827)</td>
</tr>
<tr>
<td>(1A.2)</td>
<td>0.691</td>
<td>0.802</td>
</tr>
<tr>
<td></td>
<td>(4.768)</td>
<td>(5.087)</td>
</tr>
<tr>
<td>(1B)</td>
<td>0.594</td>
<td>0.706</td>
</tr>
<tr>
<td></td>
<td>(4.204)</td>
<td>(4.633)</td>
</tr>
<tr>
<td>(1B.1)</td>
<td>0.710</td>
<td>0.815</td>
</tr>
<tr>
<td></td>
<td>(4.451)</td>
<td>(4.706)</td>
</tr>
<tr>
<td>(1B.2)</td>
<td>0.635</td>
<td>0.747</td>
</tr>
<tr>
<td></td>
<td>(4.490)</td>
<td>(4.891)</td>
</tr>
</tbody>
</table>

**Note:** All estimates of $\partial \Delta T / \partial \Delta P$ and $\partial \Delta T / \partial \Delta D_{-1}$ (and the corresponding t-ratios) are calculated from the OLS coefficient estimates in Tables 1 and 2. The values of $\Delta P$ and $D_{-1}$ are the sample mean values of these variables for each indicated period.
test level,\textsuperscript{32} while all coefficient estimates for $D_{t-1}$ in Table 2 are positive and statistically significant at the 1 percent level. Columns (4) and (5) of Table 3 contain the estimated values of $\beta\Delta T_t/\beta D_{t-1}$ for the sub-periods 1925-1943 and 1944-1966 that are implied by the parameter estimates in Tables 1 and 2: those for equation (1A) and its restricted versions are independent of $D_{t-1}$; those for equation (1B) and its restricted versions are inversely proportional to the squared value of $D_{t-1}$, and are evaluated at the respective mean values of $D_{t-1}$ for the two sub-periods.\textsuperscript{34} In accordance with the saturation hypothesis, all estimates of the mean value of $\beta\Delta T_t/\beta D_{t-1}$ given in Table 3 are negatively signed, and all are significantly different from zero at the 1 percent test level. For both the sub-periods 1925-1943 and 1944-1966, the point estimates of $\beta\Delta T_t/\beta D_{t-1}$ implied by the estimates of $\beta_4$ and $\beta_9$ in equation (1B) are negative but increasing in $D_{t-1}$: i.e., the negative effect on $\Delta T_t$ of a 1 percentage point increase in $D_{t-1}$ becomes smaller in absolute terms as the value of $D_{t-1}$ increases. Thus, the parameter estimates of equation (1B), though obviously in accordance with the saturation hypothesis, may also be construed as being consistent with the acceptance/demonstration hypothesis in the sense that higher values of $D_{t-1}$ are associated with smaller negative values of $\beta\Delta T_t/\beta D_{t-1}$.

All coefficient estimates for the regressor $DD_t(U_t-U_t^*)$ in Tables 1 and 2 are positive, numerically large, statistically significant on a two-tailed test at the 1 percent test level, and exhibit very little variation across equations. Conditional on the present interpretation of the variable $DD_t(U_t-U_t^*)$, these findings appear to substantiate the importance of worker grievances as a causal determinant of observed
variations in Canadian union membership growth. At the very least, they indicate that the occurrence and severity of cyclical contractions, for whatever reasons, exerted a strong positive effect on aggregate union membership growth in Canada during the 1925-1966 period.

The estimates in Tables 1 and 2 provide unambiguous empirical support for the proposition that the growth rate of American union membership had a positive effect on Canadian union membership growth over the 1925-1966 estimation period. The coefficient estimates for $\Delta T^{US, t}$ are all positively signed and significantly different from zero at the 1 percent test level, and range in magnitude from 0.27 to 0.31. In addition, neither equation (1A) nor (1B) yields any evidence of a finite distributed lag relationship between $\Delta T_t$ and $\Delta T^{US, t}$, a finding which suggests that the response of Canadian union growth to American union growth is adequately captured by the contemporaneous rate of change of U.S. union membership.

Contrary to the interpretation of $\Delta S_{t-2}$ as a proxy measure of union militancy, the coefficient estimates for $\Delta S_{t-2}$ in Tables 1 and 2 are uniformly negative and statistically significant on a two-tailed test at the 1 percent level. Furthermore, the finding of a negative and significant effect of $\Delta S_{t-2}$ on $\Delta T_t$ is unaltered when the fixed two-year lag is replaced by an unrestricted three-year distributed lag on $\Delta S_{t-i}$ for $i=1,2,3$. As pointed out in Section 3.6, one possible explanation for this finding is that variations in $\Delta S_{t-2}$ are proxying the effect on union growth of variations in one of the costs of unionization — namely, the costs to workers and employers of work stoppages; these costs are likely to exert a negative influence on the growth rate of union membership. Although the estimates of this negative effect are extremely small in absolute terms,
their statistical significance is at least consistent with this alternative explanation.

The parameter estimates in Tables 1 and 2 provide somewhat mixed evidence concerning the effect on $\Delta T_t$ of the 1944 enactment of Order-in-Council P.C. 1003 by the federal government. Recall that equations (1A) and (1B) incorporate two distinct hypotheses respecting the effect on Canadian union growth of the pro-union provisions of P.C. 1003. The first asserts that the adjusted mean value of $\Delta T_t$ was higher during the 1944-1966 period following enactment of P.C. 1003 than it was during the 1925-1943 period; this hypothesis implies a positive sign for the coefficient of the intercept shift dummy variable $D44_t$ in both equations. The second hypothesis contends that the effect of existing union density on union membership growth, $\partial \Delta T_t / \partial D_{t-1}$, was algebraically larger during the 1944-1966 period compared with the 1925-1943 period; this hypothesis implies a positive sign for the coefficient of the interaction term $D44_t \cdot D_{t-1}$ in equation (1A), and a negative sign for the coefficient of the interaction term $D44_t \cdot D_{t-1}^{-1}$ in equation (1B).

The estimates of both equations (1A) and (1B) confirm the joint significance of the coefficients of $D44_t$ and either $D44_t \cdot D_{t-1}$ or $D44_t \cdot D_{t-1}^{-1}$ at the 1 percent test level. However, the estimates of equations (1A) and (1B) yield quite different inferences concerning the validity of the two hypotheses when each is considered individually. In equation (1A), the coefficient estimate for $D44_t$ has the expected positive sign and is statistically significant, while that for $D44_t \cdot D_{t-1}$ is negative but statistically insignificant even at very high test levels. These results imply that, when tested against equation (1A), equation (1A.2) must be rejected,
whereas equation (1A.1) is not. In equation (1B), the coefficient estimate for \( D_{44t} \) is negative (contrary to expectations) but is statistically insignificant at the 10 percent level, while that for \( D_{44t-1}^{-1} \) is positive (again contrary to expectations) and statistically significant at the 1 percent test level. These results imply that, when tested against equation (1B), equation (1B.1) is rejected but equation (1B.2) is not rejected at conventional test levels. Although the positive coefficient estimates for \( D_{44t-1} \) in equations (1B) and (1B.2) imply that the value of \( \partial \Delta t / \partial D_{t-1} \) was significantly smaller in the 1944-1966 sub-period than in the 1925-1943 sub-period for a given value of \( D_{t-1} \), the estimated mean values of \( \partial \Delta t / \partial D_{t-1} \) for the two periods, given in Table 3, convey a somewhat different impression. In particular, when evaluated at the respective mean values of \( D_{t-1} \) in the two sub-periods, the mean estimates of \( \partial \Delta t / \partial D_{t-1} \) derived from equation (1B) are virtually identical for the two sub-periods (-0.779 for 1925-1943 and -0.770 for 1944-1966), while those derived from equation (1B.2) are -0.603 for the 1925-1943 period and -0.596 for the 1944-1966 period. Thus, the mean estimates of \( \partial \Delta t / \partial D_{t-1} \) for the two sub-periods from equation (1B.2) at least partially corroborate, while those from equation (1B) do not contradict, the hypothesis that the effect of union density on Canadian union growth was algebraically larger (but still negative) during the 1944-1966 period compared with the 1925-1943 period.\(^{39}\)

A final observation concerns the coefficient estimates for \( D_{44t} \) in equations (1A) and (1A.1). Although statistically they support the prediction of a positive coefficient for \( D_{44t} \), these coefficient estimates are so large that they cast doubt on both the appropriateness and the interpre-
tation of the dummy variable specification of the intercept shift. An alternative and perhaps better way to model the adjustment to the new labor relations framework established by P.C. 1003 might be to specify a polynomial (or some other) transition function for the intercept parameter.⁴⁰ Be that as it may, the implausibly large coefficient estimates for D₄₄ in equations (1A) and (1A.1) suggest that the implied intercept shifts may be capturing not only the stimulative effects of the pro-union provisions of P.C. 1003, but also the positive effects of other factors that distinguish the post-World War II period from the earlier period but that are not explicitly accounted for elsewhere in the present model.

This section concludes by presenting the parameter estimates obtained by fitting a somewhat modified version of the original Ashenfelter-Pencavel (1969) model of union growth to the Canadian time series data for the estimation period 1925-1966. These estimates of the modified Ashenfelter-Pencavel model are intended to provide a benchmark for assessing the empirical performance of the model of Canadian union growth proposed in this study.

Two variants of the modified Ashenfelter-Pencavel model were estimated: variant A incorporates the linear form of the union density effect; variant B incorporates the nonlinear (i.e. reciprocal) form of the union density effect. These two variants of the modified Ashenfelter-Pencavel model correspond respectively to the following two estimating equations for ΔTₖ:
(2A) \[ \Delta T_t = \gamma_0 + \gamma_1 \Delta P_t + \sum_{i=0}^{i=3} \gamma_{2i} \Delta E_{t-i} + \gamma_3 D_{t-1} \]

\[ + \gamma_4 \left[ \lambda(t-\theta) U^F_t \right] + \gamma_5 D_{44t} + v_{A,t} ; \]

(2B) \[ \Delta T_t = \delta_0 + \delta_1 \Delta P_t + \sum_{i=0}^{i=3} \delta_{2i} \Delta E_{t-i} + \delta_3 D_{t-1} \]

\[ + \delta_4 \left[ \lambda(t-\theta) U^F_t \right] + \delta_5 D_{44t} + v_{B,t} . \]

All variables were defined and measured as in equations (1A) and (1B). In keeping with the original Ashenfelter-Pencavel (1969) specification, the distributed lag on current and past employment changes, \( \Delta E_{t-i} \) (i=0,1,2,3), was approximated by a second-degree Almon polynomial lag with the far endpoint constraint imposed.\(^{41}\) The intercept shift dummy variable \( D_{44t} \) was used in place of Ashenfelter and Pencavel's (1969) original proxy measure of public and legislative sentiments towards unionism.\(^{42}\) As mentioned in Section 3.4, the function \( \lambda(t-\theta) U^F_t \) represents Ashenfelter and Pencavel's specification of the effect on union growth of worker discontent; its coefficients \( \gamma_4 \) and \( \delta_4 \) are therefore predicted to be positive.\(^{43}\) The remaining coefficients in equations (2A) and (2B) are predicted to have the same signs as the corresponding coefficients in equations (1A) and (1B).

Table 4 presents the GLS parameter estimates of equations (2A) and (2B) for the estimation period 1925-1966. In the interests of brevity, these estimates are not discussed in any detail. However, a few general features of the results warrant brief mention. First, for both equations,
the value of \( \lambda \) that minimized the sum of squared residuals over the closed unit interval was found to be \( \lambda = 0 \). Literally interpreted, this finding implies an instantaneous and complete response of \( \Delta T_t \) to \( u_t^p \), an implication that casts some doubt on the plausibility of this particular specification of the effect of worker discontent on Canadian union growth. Second, the coefficient estimate for \( \Delta P_t \) is positive in both equations, but is at best only marginally significant. Third, the estimates of the Almon-constrained lag coefficients on \( \Delta E_{t-1} \) and of their sum over \( i = 0,1,2,3 \) are positive and statistically significant. Finally, the hypothesis of a negative effect of \( D_{t-1} \) on \( \Delta T_t \) is statistically supported by the coefficient estimate for \( D_{t-1} \) in equation (2A), but not by that for \( D_{t-1}^{-1} \) in equation (2B).

On balance, the empirical performance of the modified Ashenfelter-Pencavel model on the Canadian data for the 1925-1966 estimation period is not as favorable as that of the present model given by equations (1A) and (1B). Equations (2A) and (2B) not only produce substantially worse fits to the data, but also provide little or very marginal support for several of the hypotheses incorporated in them. Nonetheless, this comparison does not necessarily imply that equations (1A) and (1B) provide valid representations of the sample data; evidence on this question is presented in Section 6. The following section prefaces this evidence by reporting the results of testing equations (1A) and (1B) for various types of specification error.
### TABLE 4

Ordinary Least Squares Estimates of Equations (2A) and (2B), A Modified Ashenfelter-Pencavel Model: Estimation Period 1925-1966

Dependent Variable = $\Delta T_t$; $\lambda = 0.0$

(t-ratios in parentheses)

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Equation (2A)</th>
<th>Equation (2B)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>7.686</td>
<td>-5.895</td>
</tr>
<tr>
<td></td>
<td>(1.995)</td>
<td>(-0.871)</td>
</tr>
<tr>
<td>$\Delta P_t$</td>
<td>0.357</td>
<td>0.444</td>
</tr>
<tr>
<td></td>
<td>(1.456)</td>
<td>(1.753)</td>
</tr>
<tr>
<td>$\Delta E_t$</td>
<td>0.909</td>
<td>0.792</td>
</tr>
<tr>
<td></td>
<td>(4.012)</td>
<td>(3.450)</td>
</tr>
<tr>
<td>$\Delta E_{t-1}$</td>
<td>0.554</td>
<td>0.511</td>
</tr>
<tr>
<td></td>
<td>(4.943)</td>
<td>(4.426)</td>
</tr>
<tr>
<td>$\Delta E_{t-2}$</td>
<td>0.284</td>
<td>0.285</td>
</tr>
<tr>
<td></td>
<td>(2.765)</td>
<td>(2.647)</td>
</tr>
<tr>
<td>$\Delta E_{t-3}$</td>
<td>0.0995</td>
<td>0.115</td>
</tr>
<tr>
<td></td>
<td>(1.194)</td>
<td>(1.321)</td>
</tr>
<tr>
<td>$D_t$</td>
<td>-0.548</td>
<td>--</td>
</tr>
<tr>
<td></td>
<td>(-2.096)</td>
<td>--</td>
</tr>
<tr>
<td>$D_{t-1}$</td>
<td>--</td>
<td>85.078</td>
</tr>
<tr>
<td></td>
<td>--</td>
<td>(0.922)</td>
</tr>
<tr>
<td>$\lambda (t-\theta) P_{Ut}$</td>
<td>0.506</td>
<td>0.427</td>
</tr>
<tr>
<td></td>
<td>(2.243)</td>
<td>(1.804)</td>
</tr>
<tr>
<td>$D44_t$</td>
<td>5.970</td>
<td>0.504</td>
</tr>
<tr>
<td></td>
<td>(1.396)</td>
<td>(0.142)</td>
</tr>
<tr>
<td>$\Sigma_{i=0}^{3} \frac{a_i}{\Delta E_{t-i}}$</td>
<td>1.847</td>
<td>1.703</td>
</tr>
<tr>
<td></td>
<td>(4.943)</td>
<td>(4.426)</td>
</tr>
</tbody>
</table>
TABLE 4 (Continued)

Ordinary Least Squares Estimates of Equations (2A) and (2B), A Modified Ashenfelter-Pencavel Model: Estimation Period 1925-1966

Dependent Variable = ΔT_t; λ = 0.0

(t-ratios in parentheses)

<table>
<thead>
<tr>
<th>Regressor</th>
<th>Equation (2A)</th>
<th>Equation (2B)</th>
</tr>
</thead>
<tbody>
<tr>
<td>lnL</td>
<td>-116.689</td>
<td>-118.669</td>
</tr>
<tr>
<td>SSR</td>
<td>636.772</td>
<td>699.739</td>
</tr>
<tr>
<td>s</td>
<td>4.265</td>
<td>4.471</td>
</tr>
<tr>
<td>R²</td>
<td>0.634</td>
<td>0.597</td>
</tr>
<tr>
<td>F(x-1,T-k)</td>
<td>10.082(6,35)</td>
<td>8.649(6,35)</td>
</tr>
<tr>
<td>d₁</td>
<td>1.64</td>
<td>1.59</td>
</tr>
</tbody>
</table>

Notes: See the notes to Table 1.

\[ \frac{1}{3} \sum_{i=0}^{3} \Delta E_{t-i} \] denotes the sum of the estimated coefficients on \( \Delta E_{t-i} \) for \( i = 0, \ldots, 3 \). The coefficients on \( \Delta E_{t-1} \) (i=0,...,3) are estimated subject to the 2 independent linear restrictions implied by a second-degree Almon polynomial lag with the far endpoint constraint imposed; the endpoint constraint restricts the lag coefficient on \( \Delta E_{t-4} \) to equal zero. The sample values of the F-statistic for the null hypothesis that the distributed lag on \( \Delta E_{t-i} \) (i=0,...,3) is consistent with these 2 Almon restrictions are \( F(2,33) = 1.142 \) for equation (2A) and \( F(2,33) = 0.929 \) for equation (2B), compared with the 5 percent critical value \( F_{0.05}(7,33) = 3.29 \).
5. TESTS FOR MODEL MISSPECIFICATION

Although casual inspection of the OLS parameter estimates and summary statistics in Tables 1 and 2 suggests that the proposed model of Canadian union membership growth provides both a satisfactory fit to the sample data and tolerably precise estimates of the individual parameters, such inspection by itself is insufficient to establish the empirical veracity of the estimating equations. The present section therefore tests equations (1A) and (1B) for several possible types of model misspecification. Specifically, it investigates the validity of three distinct assumptions on which the OLS estimates in Tables 1 and 2 are predicated: first, the assumption that the error terms in equations (1A) and (1B) are non-autoregressive or serially uncorrelated; second, the assumption that the parameters of each equation are constant over the sample space; and third, the assumption that the error terms of each equation have zero mean, an assumption that Ramsey (1969) has shown is equivalent to the absence of a variety of specification errors such as omitted variables, incorrect functional form and correlation of the error term with one or more regressors. Tests for each of these three types of model misspecification are considered seriatim.
5.1 Tests for Autoregressive Errors

Tables 1 and 2 present among the reported summary statistics the sample values of the conventional Durbin-Watson \( d_1 \) statistic, which can be used to perform an exact bounds test of the null hypothesis that the error terms are serially uncorrelated against the alternative hypothesis that the error terms are first-order autoregressive. Despite the fact that the values of \( d_1 \) for both equations (1A) and (1B) are reasonably close to 2, they all fall within the inclusive region. Consequently, the Durbin-Watson test against first-order autoregressive errors yields inconclusive results for both equations.

A determinate test of the hypothesis of non-autoregressive errors is provided by either of two asymptotic test procedures, one of which is the alternative test proposed by Durbin (1970), the other the Lagrange Multiplier test recently formulated by Godfrey (1978) and by Breusch and Pagan (1980). Although they differ in their choice of test statistic, both these procedures involve essentially the same operational framework — namely, maximum likelihood (or OLS) estimation of the postulated regression equation under the null hypothesis of non-autoregressive errors, followed by OLS estimation of a linear artificial regression equation from which the sample value of the required test statistic is readily computed.

The main computational features of the test procedures may be outlined in terms of the linear regression model

\[
Y_t = \sum_{j=1}^{k} \beta_j X_{jt} + u_t, \quad t = 1, \ldots, T,
\]

where the regressors \( X_{jt} \) may include lagged values of the dependent variable. The null hypothesis is that the error terms \( \{u_t\} \) are generated by a zero-order autoregressive, or AR(0), process such that \( u_t \) is
NID \((0, \sigma^2_u)\) for \(t = 1, \ldots, T\). The alternative hypothesis specifies that the 
\[ u_t = \rho_1 u_{t-1} + \rho_2 u_{t-2} + \cdots + \rho_n u_{t-n} + \epsilon_t \]  
where \(\epsilon_t\) is NID \((0, \sigma^2_e)\) for \(t = 1, \ldots, T\). The test of the null hypothesis of AR(0) errors against 
the alternative hypothesis of AR(n) errors thus amounts to testing 
\[ H_0: \rho_i = 0 \text{ for all } i = 1, \ldots, n \text{ against } H_1: \rho_i \neq 0 \text{ for all } i = 1, \ldots, n. \]  
Both procedures employ the same two-step computational algorithm. First, 
the proposed regression equation is estimated by OLS under the null 
hypothesis, and the resulting OLS residuals \(\hat{u}_t\) \((t=1, \ldots, T)\) are retrieved. 
Second, the required test statistic is obtained from an artificial 
regression equation of the form 
\[ \hat{u}_t = \sum_{i=1}^{n} a_i \hat{u}_{t-i} + \sum_{j=1}^{k} b_j X_{jt} + v_t, \quad t = 1, \ldots, T \]  
(5.1) 
where \(\hat{u}_{t-1}\) is the \(i\)-th lagged value of the OLS residuals (the first \(i\) 
sample values of which are set equal to zero), the \(a_i\)'s and \(b_j\)'s are 
constant coefficients, and \(v_t\) is a random error term. 

Durbin's (1970) so-called alternative test exploits the orthogonality 
of the least squares residuals \(\hat{u}_t\) and the regressors \(X_{jt}\) \((j=1, \ldots, k)\). 
This orthogonality property implies that a test of \(H_0\) against \(H_1\) is equivalent 
asymptotically to any valid large sample test of the \(n\) linear restrictions 
\(a_i = 0\) for all \(i = 1, \ldots, n\) in the artificial regression equation 
(5.1). One such test is that based on the asymptotic F-statistic.
\[ F^*(n) = \frac{(T-k-n) \left( \sum_{t=1}^{T} \hat{\epsilon}_t^2 - \sum_{t=1}^{T} \hat{\phi}_t^2 \right)}{\sum_{t=1}^{T} \hat{\phi}_t^2} \]  

(5.2)

where \( \sum_{t=1}^{T} \hat{\epsilon}_t^2 \) is the sum of squared residuals when equation (5.1) is estimated under \( H_0 \) with \( a_i = 0 \) \((i=1,\ldots,n)\), and \( \sum_{t=1}^{T} \hat{\phi}_t^2 \) is the sum of squared residuals when equation (5.1) is estimated under \( H_1 \) with \( a_i \neq 0 \) \((i=1,\ldots,n)\). On the assumption that the restrictions \( a_i = 0 \) \((i=1,\ldots,n)\) in equation (5.1) are true, \( F^*(n) \) is distributed asymptotically as a chi-square variate with \( n \) degrees of freedom; i.e., under \( H_0 \), \( F^*(n) \sim \chi^2(n) \).

Alternatively, as Godfrey (1978) and Breusch and Pagan (1980) have recently demonstrated, the Lagrange Multiplier test of \( H_0: \rho_i = 0 \) for all \( i = 1,\ldots,n \) against \( H_1: \rho_i \neq 0 \) for all \( i = 1,\ldots,n \) is given by the test statistic

\[ M(n) = T \cdot R^2 \]

\[ = \frac{T \cdot \left( \sum_{t=1}^{T} \hat{\epsilon}_t^2 - \sum_{t=1}^{T} \hat{\phi}_t^2 \right)}{\sum_{t=1}^{T} \hat{\epsilon}_t^2} \]  

(5.3)

where \( R^2 = \frac{\left( \sum_{t=1}^{T} \hat{\epsilon}_t^2 - \sum_{t=1}^{T} \hat{\phi}_t^2 \right)}{\sum_{t=1}^{T} \hat{\epsilon}_t^2} \) is simply the coefficient of determination from OLS estimation of artificial regression equation (5.1). On the assumption that \( H_0 \) is true, \( M(n) \) is distributed asymptotically as a chi-square variate with \( n \) degrees of freedom; i.e., under \( H_0 \), \( M(n) \sim \chi^2(n) \).
Although they have the same asymptotic distribution under the null hypothesis $H_0$ and therefore are asymptotically equivalent, Durbin's alternative test based on $F^*(n)$ and the Lagrange Multiplier test based on $M(n)$ may yield different inferences in finite samples, because the two test statistics have different finite sample values.\(^4\)\(^7\) For this reason, both tests were used to test for autoregressive errors in equations (1A) and (1B).\(^4\)\(^8\) For each equation, the null hypothesis of AR(0) errors was tested against two alternative hypotheses: (1) the hypothesis that the errors \(\{u_t\}\) are generated by an AR(1) process of the form \(u_t = \rho_1 u_{t-1} + \varepsilon_t\); and (2) the hypothesis that the errors \(\{u_t\}\) are generated by an AR(2) process of the form \(u_t = \rho_1 u_{t-1} + \rho_2 u_{t-2} + \varepsilon_t\).

The results of the tests for autoregressive errors in both the unrestricted and restricted versions of equations (1A) and (1B) are tabulated in Table 5. They indicate uniform non-rejection of the null hypothesis of AR(0) errors against both alternative autoregressive error specifications at nominal significance levels of 25 percent or higher. The test results therefore produce no evidence of autoregressive errors in the proposed models of Canadian union membership growth.\(^4\)\(^9\)
<table>
<thead>
<tr>
<th>Equation</th>
<th>Tests for Autoregressive Errors</th>
<th>Tests for Parameter Constancy</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$H_1: AR(1)$ Errors</td>
<td>$H_1: AR(2)$ Errors</td>
</tr>
<tr>
<td></td>
<td>$F^*(1)$ Value</td>
<td>$M(1)$ Value</td>
</tr>
<tr>
<td>(1A)</td>
<td>0.410 0.605</td>
<td>0.430 0.659</td>
</tr>
<tr>
<td>(1A.1)</td>
<td>0.438 0.626</td>
<td>0.469 0.693</td>
</tr>
<tr>
<td>(1A.2)</td>
<td>0.011 0.015</td>
<td>1.253 1.798</td>
</tr>
<tr>
<td>(1B)</td>
<td>0.678 0.991</td>
<td>0.654 0.991</td>
</tr>
<tr>
<td>(1B.1)</td>
<td>0.013 0.019</td>
<td>3.000 4.066</td>
</tr>
<tr>
<td>(1B.2)</td>
<td>0.551 0.785</td>
<td>0.857 1.247</td>
</tr>
</tbody>
</table>
TABLE 5 (Continued)

Notes: \( a/ \) Under the null hypothesis of AR(0) errors, the test statistics \( F^*(n) \) and \( M(n) \) are distributed asymptotically as \( \chi^2(n) \), where \( n = 1, 2 \). Selected values of \( \chi^2_{0.10}(1) \), the 100\( \alpha \) percent critical value of the chi-square distribution with 1 degree of freedom, are: \( \chi^2_{0.10}(1) = 2.706 \); \( \chi^2_{0.05}(1) = 3.841 \); and \( \chi^2_{0.01}(1) = 6.635 \). Selected 100\( \alpha \) percent critical values of \( \chi^2_{0.10}(2) \) are: \( \chi^2_{0.10}(2) = 4.605 \); \( \chi^2_{0.05}(2) = 5.91 \); and \( \chi^2_{0.01}(2) = 9.210 \).

\( b/ \) maxF, denotes the sample value of the sequential Chow test statistic \( F_\tau \) for which the difference between \( F_\tau \) and its corresponding 5 percent critical value is a maximum over \( \tau = k+1, k+2, \ldots, T \), where \( k \) is the number of regression parameters in the estimating equation and \( T \) is the number of sample observations. The figures in parentheses are the numerator and denominator degrees of freedom for maxF, and are equal to \( (T-\tau) \) and \( (\tau-k) \), respectively.

\( c/ \) maxQ, denotes the sample value of the cusum-of-squares test statistic \( Q_\tau \) for which the absolute deviation of \( Q_\tau \) from \( E(Q_\tau) \) is a maximum over \( \tau = k+1, k+2, \ldots, T \), where \( E(Q_\tau) = (\tau-k)/(T-k) \) is the expectation of \( Q_\tau \) under the null hypothesis of parameter constancy.

\( d/ \) The (U) indicates that the corresponding 5 percent critical value is the upper 95 percent confidence limit for maxQ, under the null hypothesis of parameter constancy.
5.2 Tests for Parameter Non-Constancy

The formulation and estimation of equations (1A) and (1B) are based on the assumption that their respective parameters are time invariant, i.e., are constant over the sample space. This assumption of parameter constancy corresponds to the null hypothesis $H_0: \beta_T = \beta$ and $\sigma^2_T = \sigma^2$ for all $t = 1, \ldots, T$, where $\beta_T$ denotes a $k$-vector of regression parameters and $\sigma^2_T$ denotes the error variance in a univariate linear regression model, and where the time subscript "t" indicates that the parameters may be time dependent. This section reports the results of testing the null hypothesis of parameter constancy against the alternative hypothesis of parameter non-constancy over time, i.e., against $H_1: \beta_T \neq \beta$ and/or $\sigma^2_T \neq \sigma^2$ for all $t = 1, \ldots, T$.

Two sequential procedures — the sequential prediction interval, or Chow (1960), test and the cusum-of-squares test developed by Brown, Durbin and Evans (1975) — are employed to investigate the intertemporal parameter stability of equations (1A) and (1B). Although computationally somewhat more demanding than the conventional split-sample procedures used to test the hypothesis of parameter constancy in previous studies of aggregate union membership growth, these sequential procedures have a distinct advantage: unlike the split-sample tests, they do not require that the investigator have prior information identifying the point(s) in time when a shift or change in the parameters of the model may have occurred. They therefore represent a more comprehensive means for detecting whether a particular regression relationship has shifted over the period spanned by the sample data.
The sequential prediction interval test is a straightforward generalization of the familiar split-sample Chow (1960) test for parameter equality between two sample sub-periods. It involves evaluating for each \( r = k+1, k+2, \ldots, T \) the Chow (1960) test statistic

\[
F_r = \frac{(S_T - S_r)/(T-r)}{S_r/(r-k)}
\]  

(5.4)

where \( S_T \) is the sum of squared OLS residuals from estimation of the model on all \( T \) sample observations, \( S_r \) is the sum of squared OLS residuals from estimation of the model over the first \( r \) sample observations, and \( k \) denotes the number of free regression parameters in the model. Under the null hypothesis of parameter constancy, \( F_r \) is distributed as \( F(T-r,r-k) \), the central \( F \)-distribution with \( (T-r) \) numerator degrees of freedom and \( (r-k) \) denominator degrees of freedom. The occurrence of sample values of \( F_r > F_{\alpha}(T-r,r-k) \), the upper \( \alpha \) significance point of the \( F(T-r,r-k) \) distribution, implies rejection of the null hypothesis of parameter constancy at the \( 100\alpha \) percent significance level.

The cusum-of-squares test proposed by Brown, Durbin and Evans (1975) is based on the behavior of a linear regression model's recursive residuals \( w_r \) for \( r = k+1, k+2, \ldots, T \), where the \( r \)-th recursive residual \( w_r \) is an appropriately normalized one-period-ahead prediction error obtained from least squares estimation of the regression equation on the first \( (r-1) \) sample observations. The cusum-of-squares test statistic \( Q_r \) is a normalized cumulative sum of the squared values of these recursive residuals, which are related to successive values of the sum of squared OLS residuals.
by the equality \( w_r^2 = S_r - S_{r-1} \) for \( r = k+1, k+2, \ldots, T \); it takes the form

\[
Q_r = \frac{\sum_{t=k+1}^{T} w_t^2}{\sum_{t=k+1}^{T} w_t^2/S_T} = \frac{S_r}{S_T}, \quad r = k+1, k+2, \ldots, T
\]  

(5.5)

Since \( Q_r = 0 \) for \( r < k \) and \( Q_r = 1 \) for \( r = T \), it follows from (5.5) that sample values of \( Q_r \) are non-decreasing in \( r \) and lie in the unit interval between zero and one. Under the null hypothesis \( H_0 \), the cuseum-of-squares test statistic has expectation \( E(Q_r) = (r-k)/(T-k) \) for \( r = k+1, k+2, \ldots, T \); accordingly, if the null hypothesis of parameter constancy is true, the plot of \( Q_r \) against \( r \) should lie along this mean value line. To test the significance of deviations of sample values of \( Q_r \) from the corresponding expected values \( E(Q_r) \) under \( H_0 \), a pair of parallel confidence lines given by

\[
Q_r = (r-k)/(T-k) \pm c_\alpha
\]  

(5.6)

is constructed such that the resulting confidence band will totally contain any series of \( Q_r \) sample values with probability \( (1-\alpha) \), provided that \( H_0 \) is true. The significance values \( c_\alpha \) for alternative significance levels \( \alpha \) and for alternative values of \( T \) and \( k \) are contained in Table 1 of Durbin (1969). Thus, a sample value of \( Q_r \) such that either \( Q_r > (r-k)/(T-k) + c_\alpha \) or \( Q_r < (r-k)/(T-k) - c_\alpha \) indicates rejection of the null hypothesis of parameter constancy at the 100\( \alpha \) percent significance level, and is interpreted as signifying the occurrence of a departure from parameter
constancy at or somewhat before the time $t = r$.\footnote{51}

The results of both tests of parameter constancy over the estimation period 1925–1966 are reported in Table 5 for the unrestricted and restricted versions of equations (1A) and (1B). The value of the cusum-of-squares test statistic $Q_r$ reported for each equation is that for which the absolute deviation of $Q_r$ from $E(Q_r)$ attains a maximum over the interval $r = k+1, k+2, \ldots T$; this value is referred to as the maximum value of $Q_r$, and is denoted by $\text{max} Q_r$. The 5 percent critical value reported with each maximum value of $Q_r$ is equal to either (i) the corresponding upper 95 percent confidence limit if the maximum value of $Q_r$ is greater than its expected value $E(Q_r)$ under $H_0$, or (ii) the corresponding lower 95 percent confidence limit if the maximum value of $Q_r$ is less than $E(Q_r)$ under $H_0$. Analogously, the sample value of the sequential Chow test statistic $F_r$ reported for each equation is that for which the difference between $F_r$ and its corresponding 5 percent critical value attains a maximum over the interval $r = k+1, k+2, \ldots T$; this value is designated the maximum value of $F_r$, and is denoted by $\text{max} F_r$ in Table 5. The results in Table 5 indicate that neither the cusum-of-squares test nor the sequential Chow test implies rejection of the null hypothesis of parameter constancy for any of the six estimating equations at the 5 percent significance level. None of the sample values of $\text{max} Q_r$ exceed their respective upper 95 percent confidence limits, nor do any of the sample values of $\text{max} F_r$ exceed their respective 5 percent critical values. In fact, the non-rejection of parameter constancy would still obtain for all six equations even at the stricter 10 percent significance level. Thus, all six equations for $\Delta T_t$ can be considered to exhibit constant parameters over the 1925–1966 estimation period at both 5 and 10 percent significance levels.
5.3 **Tests for Non-Zero Error Means**

In a seminal paper on specification errors in the context of the classical linear regression model, Ramsey (1969) derived the distributions of the OLS residuals under a variety of different types of specification error and developed several procedures for testing for the presence of these errors. This section employs one of Ramsey's (1969) procedures -- the RESET procedure -- to test for the presence of specification errors in the models of Canadian union growth given by equations (1A) and (1B).

The types of specification error for which the RESET test is appropriate include omitted variables, incorrect functional form and simultaneity problems (correlation between the error term and one or more regressors). Ramsey (1969) demonstrated that, if the error terms in the linear regression model are normally and independently distributed, the effect of these specification errors is to yield OLS residuals that are still normally distributed but that do not have zero means.

More specifically, given the standard linear regression model
\[ y = X\beta + u, \]
the RESET procedure tests the null hypothesis \( H_0: u \sim N(0, \sigma_0^2 I_T) \) against the alternative hypothesis \( H_1: u \sim N(\xi, \sigma_1^2 I_T) \). In other words, the RESET procedure tests the null hypothesis that the mean of the error vector \( u \) conditional on \( X \) is zero -- i.e., \( E(u|X) = 0 \) -- against the alternative hypothesis that the conditional mean of \( u \) is non-zero -- i.e., \( E(u|X) = \xi \) -- on the assumption that \( u \) is normally distributed under both the null and the alternative hypotheses.
Ramsey and Schmidt's (1976) reformulation of the original Ramsey (1969) RESET test in terms of OLS residuals assumes that the conditional mean of the error vector $u$ under the alternative hypothesis $H_1$ can be approximated by a linear combination of observable variables; that is, it assumes that $\xi = Q\gamma$, where $Q$ is a $T \times s$ matrix of observations on $s$ test variables and $\gamma$ is a $s \times 1$ vector of constant coefficients. The reformulated RESET test amounts to performing a conventional $F$-test of the hypothesis $\gamma = 0$ against the alternative $\gamma \neq 0$ in the augmented linear regression equation\(^{53}\)

$$y = X\beta + Q\gamma + e$$

(5.7)

Under the null hypothesis $H_0$, the resulting test statistic $F(s, T-k-s)$ has the central $F$-distribution with $s$ and $(T-k-s)$ degrees of freedom in the numerator and denominator, respectively.

The major practical problem that arises in applying the RESET test involves choosing an appropriate matrix of test variables $Q$. In view of the results of Thursby and Schmidt's (1977) Monte Carlo experiments respecting the power of several alternative choices of test variables and the relatively small sample size in the present case, two restricted variants of the RESET procedure were employed to test for non-zero error means in equations (1A) and (1B). In accordance with Ramsey's (1969) original suggestion, the first specifies the matrix of test variables to include the second, third and fourth powers of $\gamma$, the OLS fitted values of $y$ under $H_0$, so that $Q = [\gamma^{(2)} \gamma^{(3)} \gamma^{(4)}]$. The second specifies the matrix
of test variables to include $X^{(2)}$, the squared values of the regressors in $X$, so that $Q = X^{(2)}$. In the case of the second variant, the squared values of some regressors are equal to other regressors originally included in equations (1A) and (1B), and hence must be excluded from $Q$ in order to avoid perfect collinearity among the regressors in the augmented regression equation (5.7). For this reason, the number of test variables $s$ in $Q$ is less than the number of original regressors $k$ in $X$ for the second variant of the RESET test of equations (1A) and (1B).

The results obtained by applying the two variants of the RESET procedure to equations (1A) and (1B) are given in Table 6. For both equations, neither variant of the RESET test indicates rejection of the null hypothesis of zero mean errors at the 5 percent significance level (or, for that matter, at the more stringent 10 percent significance level). Thus, the RESET tests fail to detect in equations (1A) and (1B) the presence of those types of specification error that result in normally-distributed OLS residuals which have non-zero means.
TABLE 6
Results of RESET Tests for Specification Errors in Equations (1A) and (1B): Estimation Period 1925-1966

<table>
<thead>
<tr>
<th>Equation</th>
<th>Test Variables: Powers of $\hat{y}$</th>
<th>Test Variables: Squares of $X$</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$Q = [\hat{y}^{(2)}, \hat{y}^{(3)}, \hat{y}^{(4)}]$</td>
<td>$Q = [X^{(2)}]$</td>
</tr>
<tr>
<td></td>
<td>Sample Value of $F(s, T-k-s)$</td>
<td>5 Percent Critical Value</td>
</tr>
<tr>
<td>(1A)</td>
<td>1.602</td>
<td>2.98</td>
</tr>
<tr>
<td></td>
<td>(3,26)</td>
<td></td>
</tr>
<tr>
<td>(1B)</td>
<td>1.228</td>
<td>2.98</td>
</tr>
<tr>
<td></td>
<td>(3,26)</td>
<td></td>
</tr>
</tbody>
</table>

Notes: Under the null hypothesis of no specification errors, (i.e., if the error terms of the equation are normally distributed with zero mean and constant variance), the RESET test statistic $F(s, T-k-s)$ is distributed as the central $F$-distribution with $s$ and $(T-k-s)$ degrees of freedom, where $s$ is the number of test variables in $Q$, $k$ is the number of regression parameters in the proposed estimating equation, and $T$ is the number of sample observations.
6. TESTS OF MODEL SPECIFICATION

The tests for model misspecification reported in the previous section appear to confirm the empirical appropriateness of both variants of the proposed model of Canadian union growth for the estimation period 1925-1966. However, by themselves, these tests indicate only whether each model is, in its own right, consistent with the estimation period sample data; they do not take into account the existence of alternative models, each of which purports to represent or explain the same data. The present section therefore supplements the tests for model misspecification by submitting the models represented by equations (1A) and (1B) to tests of model specification based on the presence of alternative models for $\Delta T_t$.

These tests of model specification are directed specifically towards obtaining empirical evidence on each of the following questions: (1) which, if either, of the two equations (1A) and (1B) provides a valid representation of the sample data for the estimation period 1925-1966?; (2) which of two alternative measures of aggregate labor market conditions — the rate of change of unionizable employment $\Delta E_t$, or the aggregate unemployment rate $U_t$ — is consistent with the sample data?; and (3) can the estimating equations (1A) and (1B) be considered valid representations of the sample data in view of the performance of the modified Ashenfelter-Pencavel model of Canadian union growth represented by equations (2A) and (2B)?

These questions are addressed empirically by means of both nested and non-nested hypothesis testing procedures. The non-nested hypothesis testing procedure employed here is the pairwise J-test, which is one of several procedures recently proposed by Davidson and MacKinnon (1981) for
testing non-nested (and possibly non-linear) regression models. Like the
tests for autoregressive errors, parameter non-constancy, and non-zero
error means considered in Section 5, non-nested hypothesis tests are also
tests of model specification; however, they differ from the more familiar
tests for model misspecification by virtue of their reliance on the
existence of non-nested alternative models. The Davidson-MacKinnon (1981)
tests are based on the Artificial Nesting Principle of non-nested hypothe-
sis testing, whereby the test statistic required to test the truth of one
model against a non-nested alternative model is computed from an artificial
regression equation in which the tested model and the alternative model are
embedded parametrically. This artificial regression equation is so for-
mulated that the truth of the tested model vis-a-vis the alternative model
corresponds to a parameter restriction on the artificial regression
equation. Tests of this parameter restriction on the artificial regression
equation therefore yield tests of the validity of the tested model against
the alternative model.

The essential features of Davidson and MacKinnon's (1981) pairwise
J-test may be outlined in terms of the two linear regression models

\[ H_0: \ y = \mathbf{X} \beta + \mathbf{u}_0, \ \mathbf{u}_0 \sim \mathcal{N}(0, \sigma_0^2 \mathbf{I}) \]  \hspace{1cm} (6.1)

and

\[ H_1: \ y = \mathbf{Z} \gamma + \mathbf{u}_1, \ \mathbf{u}_1 \sim \mathcal{N}(0, \sigma_1^2 \mathbf{I}_T) \]  \hspace{1cm} (6.2)
where: \( y \) is a \( T \times 1 \) vector of observations on the dependent variable; \( X \) and \( Z \) are, respectively, \( T \times k_0 \) and \( T \times k_1 \) regressor data matrices; \( \beta \) and \( \gamma \) are, respectively, \( k_0 \times 1 \) and \( k_1 \times 1 \) vectors of unknown regression parameters; and \( u_0 \) and \( u_1 \) are, respectively, \( T \times 1 \) error vectors, each of which is assumed to be identically and independently distributed as the normal distribution if the corresponding model is true. The models \( H_0 \) and \( H_1 \) are assumed to be non-nested in the sense that neither can be obtained by imposing suitable parameter restrictions on the other. The objective is to test the specification of the model \( H_0 \) against the alternative model \( H_1 \) by determining whether the performance of the latter on the sample data is consistent with the assumed truth of the former.

The pairwise \( J \)-test of the model \( H_0 \) in (6.1) against the alternative model \( H_1 \) in (6.2) involves estimation of the artificial regression equation

\[
y = (1-\alpha)X\beta + \alpha Z\hat{\gamma} + u \tag{6.3}
\]

where \( \alpha \) is a scalar nesting parameter, \( \hat{\gamma} = (Z'Z)^{-1}Z'y \) is the OLS estimate of \( \gamma \) under \( H_1 \), and \( Z\hat{\gamma} = \hat{\gamma}_1 \) is the \( T \)-vector of fitted values of \( y \) under \( H_1 \). From (6.3), it is obvious that, if \( H_0 \) is true, then the true value of \( \alpha \) is zero. Accordingly, given a consistent estimate \( \hat{\alpha} \) of \( \alpha \) in (6.3), a value of \( \hat{\alpha} \) significantly different from zero would imply rejection of the model \( H_0 \), while a value of \( \hat{\alpha} \) not significantly different from zero would imply non-rejection of \( H_0 \). Davidson and MacKinnon (1981) prove that, if \( H_0 \) is true, the conventional \( t \)-statistic \( t_\hat{\alpha} \) for the estimate \( \hat{\alpha} \) of \( \alpha \) obtained from OLS
estimation of (6.3) is distributed asymptotically as \( N(0,1) \), the unit normal distribution. Consequently, an asymptotically valid test of the truth of \( H_0 \) can be computed by merely estimating the artificial regression equation (6.3) and calculating the conventional t-ratio \( t_A \) corresponding to the hypothesis \( a = 0 \). Then, in order to obtain a consistent test of the truth of the model \( H_1 \) against the model \( H_0 \), the simplest procedure is simply to reverse the roles of the two models and repeat the calculations of the J-test.\(^{54}\) Thus, for any pair of non-nested hypotheses \( H_0 \) and \( H_1 \), four test outcomes are possible: (1) reject \( H_0 \) and accept \( H_1 \); (2) accept \( H_0 \) and reject \( H_1 \); (3) accept both \( H_0 \) and \( H_1 \); and (4) reject both \( H_0 \) and \( H_1 \).

An obvious alternative approach to testing the truth of models \( H_0 \) and \( H_1 \) is one based on standard procedures for testing nested hypotheses. Suppose that the two models of interest are again \( H_0 \) and \( H_1 \). These models may be rewritten as

\[
H_0: y = X_\star \beta_\star + W\delta_0 + u_0
\]

and

\[
H_1: y = Z_\star \gamma_\star + W\delta_1 + u_1
\]

where: \( X_\star \) and \( Z_\star \) are data matrices of regressors unique to \( H_0 \) and \( H_1 \), respectively; and \( W \) is a data matrix of regressors common to \( H_0 \) and \( H_1 \). Given this partitioning of the regressor set of each model, a compound model can be formed for which the regressor set is simply the union of the
regressor sets of the models $H_0$ and $H_1$: this compound model may be written as

$$H_C: y = X_\gamma \gamma + Z_\varphi \varphi + \mu$$

(6.4)

Both the original models are obviously nested in the compound model $H_C$:

- model $H_0$ corresponds to the parameter restrictions $\gamma_* = 0$ on $H_C$;
- model $H_1$ corresponds to the parameter restrictions $\varphi_* = 0$ on $H_C$. Thus, adopting the compound model $H_C$ as the maintained hypothesis, the model $H_0$ may be tested against $H_C$ in the presence of the alternative model $H_1$ by performing a conventional $F$-test of the hypothesis $\gamma_* = 0$ against the alternative $\gamma_* \neq 0$;

- similarly, the model $H_1$ may be tested against $H_C$ in the presence of the alternative model $H_0$ by performing an $F$-test of the hypothesis $\varphi_* = 0$ against the alternative $\varphi_* \neq 0$.

The $F$-test procedure has the same four possible outcomes as the non-nested $J$-test with respect to the validity of the two models $H_0$ and $H_1$; however, the interpretation given these outcomes is quite different. Moreover, since little is known about their small sample properties or their power, it is by no means obvious which of the two procedures is more appropriate in any particular application. Consequently, both the nested $F$-test and the non-nested $J$-test are employed here to perform pairwise specification tests of several alternative models of Canadian union membership growth.
The first question addressed is which of the two alternative forms of the union density effect — the linear form \( D_{t-1} \) incorporated in Model A, or the nonlinear (reciprocal) form \( D_{t-1}^{-1} \) incorporated in Model B — provides a valid representation of the sample data. Since equations (1A) and (1B) differ from each other only with respect to the formulation of the density effect, this question is addressed by testing (1A) and (1B) against one another. For comparative purposes, the linear and nonlinear variants of the modified Ashenfelter-Pencavel model given by equations (2A) and (2B), respectively, are also tested against each other. The results of both the J-test and F-test are tabulated in panel I of Table 7. They indicate that neither model (1A) nor model (1B) can be rejected against the other on either test, even at very high significance levels. Consequently, both specifications of the density effect appear to be valid, at least within the framework of the proposed model. In contrast, the test results for the modified Ashenfelter-Pencavel model imply rejection of the reciprocal form, but nonrejection of the linear form, of the union density effect. Specifically, the sample values of both the J-test statistic and the F-test statistic when equation (2A) is tested against equation (2B) are not statistically significant at even the 10 percent test level; however, when equation (2B) is tested against equation (2A), both tests yield sample values of their respective test statistics that are significant at the 5 percent (though not the 1 percent) test level. Thus, of the two variants of the modified Ashenfelter-Pencavel model, only the linear form of the density effect incorporated in equation (2A) can be accepted as being true.
TABLE 7
Results of Pairwise Tests of Model Specification:
Estimation Period 1925-1966

<table>
<thead>
<tr>
<th>Tested Hypothesis $H_0$</th>
<th>Alternative Hypothesis $H_1$</th>
<th>Non-Nested J-Test</th>
<th>Nested F-Test</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>Sample Value of $t^{a}$</td>
<td>Sample Value of $F(n_1,n_2)^{b}$</td>
</tr>
<tr>
<td>Panel I</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(1A)</td>
<td>(1B)</td>
<td>0.445</td>
<td>0.748 (2,27)</td>
</tr>
<tr>
<td>(1B)</td>
<td>(1A)</td>
<td>0.130</td>
<td>0.663 (2,27)</td>
</tr>
<tr>
<td>(2A)</td>
<td>(2B)</td>
<td>-1.443</td>
<td>2.081 (1,34)</td>
</tr>
<tr>
<td>(2B)</td>
<td>(2A)</td>
<td>2.377</td>
<td>5.649 (1,34)</td>
</tr>
<tr>
<td>Panel II</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(1A)</td>
<td>(3A)</td>
<td>0.845</td>
<td>0.486 (4,25)</td>
</tr>
<tr>
<td>(3A)</td>
<td>(1A)</td>
<td>6.547</td>
<td>10.383 (4,25)</td>
</tr>
<tr>
<td>(1B)</td>
<td>(3B)</td>
<td>0.823</td>
<td>0.500 (4,25)</td>
</tr>
<tr>
<td>(3B)</td>
<td>(1B)</td>
<td>6.678</td>
<td>10.812 (4,25)</td>
</tr>
<tr>
<td>Panel III</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(1A)</td>
<td>(2A)</td>
<td>-0.247</td>
<td>0.061 (1,28)</td>
</tr>
<tr>
<td>(2A)</td>
<td>(1A)</td>
<td>10.390</td>
<td>12.710 (7,28)</td>
</tr>
<tr>
<td>(1B)</td>
<td>(2B)</td>
<td>-0.351</td>
<td>0.123 (1,28)</td>
</tr>
<tr>
<td>(2B)</td>
<td>(1B)</td>
<td>11.097</td>
<td>14.514 (7,28)</td>
</tr>
</tbody>
</table>
TABLE 7 (Continued)

Notes: a/ Under the null hypothesis that the model $H_0$ is true, the $J$-test statistic $t_0$ is distributed asymptotically as $N(0, 1)$, the standard normal distribution.

b/ Conditional on the validity of the parameter restrictions on the compound model that yield the tested model $H_0$, the test statistic $F(n_1, n_2)$ is distributed as the central $F$-distribution with $n_1$ and $n_2$ degrees of freedom, where $n_1$ is the number of regression parameters that are unique to the alternative model $H_1$, and $n_2$ is the number of degrees of freedom for the compound model (i.e., the number of sample observations minus the number of regression parameters in the compound model).
in view of the performance of the other model on the estimation period sample data.

The second question addressed here is which of two alternative measures of aggregate labor market conditions — the percentage change in employment in the unionizable sector of the economy, \( \Delta E_t \), or the aggregate unemployment rate, \( U_t \) — provides a valid representation of the sample data. Within the context of the union growth model given by equations (1A) and (1B), one straightforward approach to this question entails alternately specifying a given equation for \( \Delta T_t \) to include a finite distributed lag on \( \Delta E_{t-1} \) and a finite distributed lag on \( U_{t-1} \), and then testing each specification against the other. Accordingly, the unrestricted distributed lag on \( \Delta E_{t-1} \) over \( i = 0, 1, 2, 3 \) in equations (1A) and (1B) is replaced by an unrestricted distributed lag on \( U_{t-1} \) over \( i = 0, 1, 2, 3 \). The estimating equations so derived are designated (3A) and (3B), respectively, and are estimated on the sample data for the estimation period 1925-1966. The empirical validity of the two alternative measures of aggregate labor market conditions is then assessed by performing pairwise J- and F-tests of equations (1A) and (3A) and of equations (1B) and (3B).

Although not reported here, the OLS estimates of equations (3A) and (3B) for the estimation period 1925-1966 indicate that the distributed lag on \( U_{t-1} \) yields a substantially worse fit to the data than the distributed lag on \( \Delta E_{t-1} \).\(^57\) The maximized log-likelihood \((\ln L)\) and the coefficient of determination \( (R^2) \) are considerably smaller, and the standard error of the regression \((s)\) is substantially larger, for equations (3A) and (3B) than for equations (1A) and (1B).\(^58\) In addition, the coefficient estimates of
(3A) and (3B) indicate a substantially higher frequency of statistically insignificant regression parameters than do the coefficient estimates of (1A) and (1B). The general impression conveyed by these comparisons is that the distributed lag on $\Delta E_{t-i}$ better accounts for the sample values of $\Delta T_t$ than does the distributed lag on $U_{t-i}$. This impression is reinforced by the results for the J- and F-tests of model specification displayed in panel II of Table 7. Neither the J-test nor the F-test indicates rejection of (1A) against (3A) or of (1B) against (3B), even at much higher than conventional significance levels. However, when (3A) is tested against (1A) and when (3B) is tested against (1B), both tests imply rejection of equations (3A) and (3B) at the 1 percent significance level. These findings suggest that, given the other features of the model of Canadian union growth advanced in this study, the equations for $\Delta T_t$ which specify a distributed lag on $\Delta E_{t-i}$ ($i=0,1,2,3$) can be considered valid, while those which specify a distributed lag on $U_{t-i}$ ($i=0,1,2,3$) must be considered invalid.

The final question investigated in this section is: can the modified Ashenfelter-Pencavel model given by equations (2A) and (2B) and/or the present model given by equations (1A) and (1B) be accepted as valid representations of the estimation period sample data? Given the substantially better fit of equations (1A) and (1B) compared with equations (2A) and (2B), it is reasonable to expect that (1A) will reject (2A) and that (1B) will reject (2B). What cannot be predicted in advance is whether (2A) and (2B) will reject (1A) and (1B), respectively, since the J-test of $H_0$ against $H_1$ may well reject $H_0$ even though $H_1$ fits the data much worse than
\( H_0 \). The results of the pairwise \( J \)- and \( F \)-tests of equations (1A) and (2A), and of equations (1B) and (2B), are tabulated in panel III of Table 7. As expected, both tests clearly reject (2A) against (1A) and (2B) against (1B) at the 1 percent significance level. In addition, neither the \( J \)-test nor the \( F \)-test can reject (1A) against (2A) or (1B) against (2B), even at extremely high significance levels. These findings yield the following implications: (1) the modified Ashenfelter-Pencavel model of Canadian union growth cannot be sustained as true in view of the performance of the present model on the sample data; and (2) the present model given by equations (1A) and (1B) cannot be rejected as false in view of the performance of the modified Ashenfelter-Pencavel model on the sample data for the estimation period 1925-1966.

The present section has examined the results of submitting six alternative models of Canadian union membership growth to two pairwise tests of model specification — the non-nested \( J \)-test procedure and the nested \( F \)-test procedure. The overall findings of these tests may be summarized as follows. First, for all pairwise tests performed, the inferences yielded by the \( J \)-test are in complete agreement with those yielded by the \( F \)-test. Second, four of the six estimating equations considered — namely equations (2A), (2B), (3A) and (3B) — appear to be false, since each was rejected by at least one of the alternative models against which it was tested. Third, neither of the remaining two estimating equations — namely equations (1A) and (1B) — was rejected by any of the alternative models against which each was tested. These findings suggest that, given
the sample data and the set of alternative models considered, both variants of the model of Canadian union membership growth herein proposed can be tentatively accepted as valid representations of $\Delta T_c$ over the estimation period 1925-1966. This apparent validity of equations (1A) and (1B) is properly interpreted to mean that the performance of the six alternative models on the given sample data is not inconsistent with the truth of both equations (1A) and (1B). However, since these findings are conditional on the particular six models considered here, they conceivably could be altered by the addition of other alternative hypotheses to the model choice set, and therefore must be regarded as tentative.

7. FORECASTING PERFORMANCE OF THE MODELS

Both the tests for model misspecification in Section 5 and the tests of model specification in Section 6 evaluate the performance of the proposed union membership growth models against the sample data for the 1925-1966 estimation period. This section, in contrast, evaluates the performance of the estimated models against the post-sample data for the years 1967-1972, which correspond to the first six years following the federal government's enactment of the Public Service Staff Relations Act (PSSRA) in 1967. In addition to providing further evidence on the empirical performance of the models, this evaluation serves to provide some preliminary evidence on how the provisions of the PSSRA respecting union certification and collective bargaining in the federal public service may have affected the growth rate of aggregate union membership in Canada.
The forecasting performance of each model is investigated by assessing the accuracy of the unconditional \textit{ex post} forecasts of $\Delta T_t$ for the years 1967-1972 that are generated by the parameter estimates of each model for the estimation period 1925-1966 and the realized values of the regressors for the forecast period 1967-1972. Accordingly, if the estimation period model for $t = 1, \ldots, T$ is written as $y_t = X_1 \beta_1 + u_t$ and the forecast period model for $t = T + 1, \ldots, T + T_2$ as $y_t = X_2 \beta_2 + u_t$, the $T_2$-vector of \textit{ex post} forecasts is given by $\hat{y}_2 = X_2 \hat{\beta}_1 = X_2 (X_1' X_1)^{-1} X_1' y_1$, and the corresponding $T_2$-vector of forecast errors by $e_2 = y_2 - y_2 = X_2 \hat{\beta}_1 - y_2$, where $\hat{\beta}_1 = (X_1' X_1)^{-1} X_1' y_1$ is the $k$-vector of OLS coefficient estimates for the estimation period. One way of assessing forecasting performance is to compute various tests of the \textit{relative} predictive accuracy of each model. These tests are "relative" in the sense that they identify predictive accuracy with parameter equality over the estimation and forecast periods. Thus, "predictive accuracy" is associated with the null hypothesis $H_0: \beta_1 = \beta_2$ and $\sigma_1^2 = \sigma_2^2$; "predictive failure", on the other hand, is associated with the alternative hypothesis $H_1: \beta_1 \neq \beta_2$ and/or $\sigma_1^2 \neq \sigma_2^2$.

Both sequential and split-sample tests are employed to test the relative predictive accuracy of the union membership growth models in Tables 1 and 2. Both the sequential Chow prediction interval test and the cusum-of-squares test outlined in Section 5.2 are used to test for parameter constancy over the full sample for estimation and forecast periods. Each model is estimated on the combined sample of $T + T_2$ observations, and sample values of the sequential Chow test statistic $F_T$ in (5.4) and of the cusum-
of-squares test statistic $Q_r$ in (5.5) are calculated for all $r = k+1, \ldots, T, T+1, \ldots, T+T_2$. Sample values of $P_r$ and/or $Q_r$ that exceed their respective critical values indicate predictive failure of the model, i.e., parameter non-constancy over the estimation and forecast periods.

In addition to the two sequential test procedures, three split-sample tests are used to test the relative predictive accuracy of each model. The first is the conventional Chow (1960) prediction interval test based on the test statistic (5.4) with $r = T$ and with $T + T_2$ in place of $T$. On the assumption that the parameters of the model are the same for the estimation and forecast periods, the resulting test statistic $P_T$ is distributed as a central $F$ variate with $T_2$ and $T-k$ degrees of freedom; i.e., under $H_0$, $P_T \sim F(T_2, T-k)$. The second split-sample test is a large sample variance ratio test based on the test statistic

$$f(T_2) = \frac{\hat{\sigma}_2^2}{\hat{\sigma}_1^2}$$

(7.1)

where $\hat{\sigma}_1^2 = \hat{\sigma}_1 / (T-k)$ is the degrees-of-freedom-adjusted estimate of the error variance $\sigma_1^2$ for the estimation period model, and $e_2^T e_2$ is the sum of squared ex post forecast errors $e_2$ over the forecast period. Under the assumption that the null hypothesis of relative predictive accuracy is true, $f(T_2)$ is distributed asymptotically as a chi-square variate with $T_2$ degrees of freedom; i.e., $f(T_2) \sim \chi^2(T_2)$ under $H_0$. The third split-sample test of relative predictive accuracy is the conventional $t$-test for equality of the intercept parameter between the estimation and forecast periods. It is performed in the usual way by including in the original estimating equation a forecast period dummy variable that equals 1 for each
year of the forecast period and 0 for each year of the estimation period, and then re-estimating the augmented equation on the full sample of $T + T_2$ observations for both estimation and forecast periods. The test statistic is the conventional $t$-ratio for the hypothesis that the coefficient of the forecast period dummy variable is zero.

Table 8 presents the results for both the sequential and the split-sample tests of relative predictive accuracy. The results for the two sequential tests reveal a notable disparity between the findings for the sequential Chow prediction interval test and those for the cusum-of-squares test. For the cusum-of-squares test, not a single sample value of the test statistic $Q_T$ falls outside the interval defined by its upper and lower 95 percent confidence limits; thus, the cusum-of-squares test does not reject, at the 5 percent significance level, the null hypothesis of parameter constancy over the full sample period 1925-1972 for any of the six estimating equations. For the sequential Chow test, in contrast, almost all sample values of the test statistic $F_T$ for the years 1967-1971 exceed their corresponding 5 percent or 1 percent critical values, while none of those for the years prior to 1967 are significant at the 5 percent test level. These findings obtain both for the unrestricted equations (1A) and (1B) and for the restricted equations (1A.1), (1A.2) and (1B.1). Only for equation (1B.2) do none of the sample values of $F_T$ exceed their respective 5 percent critical values, either for years before or after 1967. Thus, the sequential Chow prediction interval test clearly indicates predictive failure for all equations other than equation (1B.2) by revealing significant departures from parameter constancy during the forecast period 1967-1972.
TABLE 8
Results of Tests of Relative Forecast Accuracy for Equations (1A) and (1B):
Estimation Period 1925-1966; Forecast Period 1967-1972

<table>
<thead>
<tr>
<th>Equation</th>
<th>Sequential Tests</th>
<th>Split-Sample Tests</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Chow Prediction</td>
<td>Chow Test: Sample</td>
</tr>
<tr>
<td></td>
<td>Interval Test$^a$</td>
<td>Test of Sample</td>
</tr>
<tr>
<td></td>
<td>Sample Value</td>
<td>Value of F$_T$($T_2$,$T-k$)$^c$/</td>
</tr>
<tr>
<td></td>
<td>of max$U_H$</td>
<td>Value of f(T$_2$)$^d$/</td>
</tr>
<tr>
<td></td>
<td>5 Percent</td>
<td>(Estimated</td>
</tr>
<tr>
<td></td>
<td>Critical Value</td>
<td>Coefficient)$^c$/</td>
</tr>
<tr>
<td></td>
<td>Year</td>
<td></td>
</tr>
<tr>
<td>(1A)</td>
<td>7.327 (2,33)</td>
<td>2.256 (6,29)</td>
</tr>
<tr>
<td></td>
<td>3.29 1970</td>
<td>16.614</td>
</tr>
<tr>
<td>(1A.1)</td>
<td>7.639 (2,34)</td>
<td>2.359 (6,30)</td>
</tr>
<tr>
<td></td>
<td>3.28 1970</td>
<td>16.774</td>
</tr>
<tr>
<td>(1A.2)</td>
<td>4.665 (2,34)</td>
<td>1.533 (6,30)</td>
</tr>
<tr>
<td></td>
<td>3.28 1970</td>
<td>9.950</td>
</tr>
<tr>
<td>(1B)</td>
<td>6.677 (2,33)</td>
<td>2.047 (6,29)</td>
</tr>
<tr>
<td></td>
<td>3.29 1970</td>
<td>14.566</td>
</tr>
<tr>
<td>(1B.1)</td>
<td>2.817 (2,34)</td>
<td>1.056 (6,30)</td>
</tr>
<tr>
<td></td>
<td>3.28 1970</td>
<td>6.574</td>
</tr>
<tr>
<td>(1B.2)</td>
<td>5.502 (2,34)</td>
<td>1.734 (6,30)</td>
</tr>
<tr>
<td></td>
<td>3.28 1970</td>
<td>11.401</td>
</tr>
</tbody>
</table>
Notes: a/ max\( F_{\kappa} \) denotes the sample value of the sequential Chow test statistic \( F_{\kappa} \) for which the difference between \( F_{\kappa} \) and its corresponding 5 percent critical value is a maximum over \( r = k+1, k+2, \ldots, T+T_2 \), where \( k \) is the number of regression parameters in the estimating equation, \( T \) is the number of sample observations in the estimation period, and \( T_2 \) is the number of observations in the forecast period. The figures in parentheses are the numerator and denominator degrees of freedom for \( \text{max} F_{\kappa} \), and are equal to \((T+T_2-k)\) and \((k-k)\), respectively. The column headed 'Year' identifies the year corresponding to \( \text{max} F_{\kappa} \).

b/ max\( Q_{\kappa} \) denotes the sample value of the cusum-of-squares test statistic \( Q_{\kappa} \) for which the absolute deviation of \( Q_{\kappa} \) from \( E(Q_{\kappa}) \) is a maximum over \( r = k+1, k+2, \ldots, T+T_2 \), where \( E(Q_{\kappa}) = (k-k)/(T+T_2-k) \) is the expectation of \( Q_{\kappa} \) under the null hypothesis of parameter constancy. An (L) after the 5 percent critical value indicates the lower 95 percent confidence limit for \( \text{max} Q_{\kappa} \), and a (U) the upper 95 percent confidence limit for \( \text{max} Q_{\kappa} \), under the null hypothesis of parameter constancy. Sample values of \( \text{max} Q_{\kappa} \) that are either greater than the corresponding lower 95 percent confidence limit or less than the corresponding upper 95 percent confidence limit imply nonrejection of the null hypothesis of parameter constancy at the 5 percent significance level. The column headed 'Year' identifies the year corresponding to \( \text{max} Q_{\kappa} \).

c/ Under the null hypothesis that the parameters of the estimating equation are equal for the estimation and forecast periods, the Chow test statistic \( F_{\kappa} \) is distributed as the central F-distribution with \( T_2 \) and \((T-k)\) degrees of freedom, where \( T_2 \) is the number of observations in the forecast period, \( T \) is the number of observations in the estimation period, and \( k \) is the number of regression parameters in the estimating equation. The figures in
TABLE 8 (Continued)

parentheses are the numerator and denominator degrees of freedom for \( F_T \), and are equal to \( T_2 \) and \((T-k)\), respectively. Here, \( T_2 = 6 \) and \((T-k) = 29 \) or \( 30 \). Selected 100\( \alpha \) percent critical values of \( F_{\alpha} (6,30) \) are: \( F_{0.10} (6,30) = 1.93 \); \( F_{0.05} (6,30) = 2.42 \); and \( F_{0.01} (6,30) = 3.47 \).

d/ Under the null hypothesis that the parameters of the estimating equation are equal for the estimation and forecast period, the variance ratio test statistic \( f(T_2) \) is distributed asymptotically as \( \chi^2 (T_2) \), where \( T_2 \) is the number of sample observations in the forecast period. Here, \( T_2 = 6 \). Selected 100\( \alpha \) percent critical values of \( \chi^2_{\alpha} (6) \) are: \( \chi^2_{0.10} (6) = 10.6 \); \( \chi^2_{0.05} (6) = 12.6 \); and \( \chi^2_{0.01} (6) = 16.8 \).

e/ The intercept shift test statistic is the absolute value of the conventional \( t \)-ratio for testing the null hypothesis that the coefficient of the forecast period dummy variable (which equals 0 for each year of the estimation period and 1 for each year of the forecast period) equals zero. Under the null hypothesis, this \( t \)-statistic is distributed as the Student's \( t \)-distribution with \((T+T_2-k-1)\) degrees of freedom, where \( T \) is the number of sample observations in the estimation period, \( T_2 \) is the number of sample observations in the forecast period, and \( k \) is the number of regression parameters other than the intercept-shift parameter that are included in the estimating equation. The figure in parentheses below each \( t \)-ratio is the estimated coefficient on the intercept-shift dummy variable.
The split-sample test results in Table 8 also provide some evidence of predictive failure. None of the sample values of the Chow prediction interval test statistic $F_T$ exceed their respective 5 percent critical values; however, for three of the six equations — namely (1A), (1A.1) and (1B) — the values of $F_T$ are sufficiently large to imply rejection of parameter equality between estimation and forecast periods at the 10 percent significance level. Only for the restricted models given by equations (1A.2), (1B.1) and (1B.2) does the prediction interval test clearly indicate non-rejection of the hypothesis of relative forecast accuracy at the 10 percent significance level. The large sample variance ratio test yields inferences similar to those implied by the prediction interval test. The sample value of $f(T_2)$ exceeds the corresponding 2.5 percent critical value for equations (1A), (1A.1) and (1B), and the 10 percent critical value for equation (1B.2). Only for equations (1A.2) and (1B.1) does the variance ratio test indicate non-rejection of the hypothesis of relative forecast accuracy at the 10 percent significance level. Finally, the forecast period dummy variable test yields positive coefficient estimates for the intercept shift dummy variable in all six equations. However, none of those coefficients is statistically significant on a two-tailed test at the 5 percent level; in fact, only for equation (1A.1) is the forecast period dummy variable statistically significant at the 10 percent level, and even then only marginally so. In summary, therefore, the forecast period dummy variable test yields virtually no evidence of predictive failure for any of the six estimating equations. The Chow prediction interval test and the variance ratio test, in contrast, produce fairly
strong evidence of predictive failure for equations (1A), (1A.1) and (1B), but little or no evidence of predictive failure for equations (1A.2), (1B.1) and (1B.2).

Table 9 supplements the tests of relative predictive accuracy by presenting some standard descriptive summary measures of the estimated equations' ex post forecasting accuracy. The first three summary measures in Table 9 are the mean prediction error (ME), the root-mean-square prediction error (RMSE), and Theil's (1966) revised inequality coefficient (U_1); each in effect compares the accuracy of the models' ex post forecasts with that of the perfect predictor, for which all three measures of forecast accuracy equal zero. The values of ME in Table 9 are all negative, indicating that all six models on average underpredict the annual percentage change in Canadian union membership during the forecast period 1967-1972. All six estimating equations yield values of the RMSE substantially greater than zero. Moreover, the values of U_1 exceed unity for all equations except (1B.1), a finding which indicates that five of the six estimated models forecast worse over the period 1967-1972 than a naive "no-change" forecasting rule (according to which the predicted change in union membership in each period is equal to last period's observed change). Finally, all three descriptive measures yield the same ranking of the six models in terms of the accuracy of their ex post forecasts: equation (1B.1) generates the most accurate forecasts, followed in descending order by equations (1A.2), (1B.2), (1A), (1A.1) and (1A). Paradoxically, the most accurate forecasts are given by those two equations that impose an invalid parameter restriction on their corresponding unrestricted equations.
TABLE 9

Descriptive Summary Measures of Forecast Accuracy for Equations (1A) and (1B): Estimation Period 1925-1966; Forecast Period 1967-1972

<table>
<thead>
<tr>
<th>Equation</th>
<th>Mean Prediction Error ME</th>
<th>Root-Mean-Square Prediction Error RMSE</th>
<th>Theil's Inequality Coefficient $U_1$</th>
<th>Theil's Inequality Proportions</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1A)</td>
<td>-2.224</td>
<td>3.819</td>
<td>1.253</td>
<td>0.339</td>
</tr>
<tr>
<td>(1A.1)</td>
<td>-2.155</td>
<td>3.775</td>
<td>1.238</td>
<td>0.326</td>
</tr>
<tr>
<td>(1A.2)</td>
<td>-1.144</td>
<td>3.169</td>
<td>1.039</td>
<td>0.130</td>
</tr>
<tr>
<td>(1B)</td>
<td>-2.071</td>
<td>3.565</td>
<td>1.169</td>
<td>0.337</td>
</tr>
<tr>
<td>(1B.1)</td>
<td>-0.470</td>
<td>2.784</td>
<td>0.913</td>
<td>0.029</td>
</tr>
<tr>
<td>(1B.2)</td>
<td>-1.458</td>
<td>3.216</td>
<td>1.055</td>
<td>0.205</td>
</tr>
</tbody>
</table>
The last three columns of Table 9 contain the values of Theil's (1966) inequality proportions for each equation. These inequality proportions are derived from Theil's (1966) demonstration that the mean square prediction error MSE can be decomposed into three additive components: a bias proportion \( U_M \), a regression proportion \( U_R \), and a disturbance proportion \( U_D \). Since the bias and regression proportions \( U_M \) and \( U_R \) reflect systematic forecast errors that are potentially avoidable, while the disturbance proportion \( U_D \) reflects unavoidable random errors, the closer to zero are \( U_M \) and \( U_R \) and the closer to one is \( U_D \), the better is the proposed forecasting model relative to the optimal predictor for which \( U_M = U_R = 0 \) and \( U_D = 1 \).

The values of the inequality proportions in Table 9 indicate that systematic sources of forecast error, as a proportion of MSE, are 43 percent for equation (1B.1), about 59 percent for equations (1A.2) and (1B.2), 68 percent for equation (1B), and nearly 75 percent for equations (1A) and (1A.1). These relatively high fractions of systematic forecast error reflect both regression proportions that are uniformly in excess of \( 1/3 \), and bias proportions that range from \( 1/5 \) to \( 1/3 \) for all equations except (1B.1) and (1A.2). Thus, the results of Theil's MSE decomposition tend to reinforce the results obtained for the previously reported descriptive measures of forecast accuracy: namely, that the estimated union growth equations for the period 1925-1966 substantially and systematically underpredict the observed growth rate of Canadian trade union membership during the forecast period 1967-1972. Together with the admittedly mixed results for the tests of relative ex post forecasting accuracy, these findings suggest that the annual percentage change in Canadian union mem-
bership was somewhat higher immediately following the enactment of the
PSSRA in 1967 than otherwise would have been expected.

8. CONCLUSIONS AND LIMITATIONS

Building on several previously proposed econometric models of union
growth, the present study has formulated a single equation model of aggrega-
gate trade union membership growth in Canada. Several variants of the
model were estimated and tested on annual Canadian time series data for the
period 1925-1966. The ex post forecasting performance of the estimated
models was then investigated using data for the period 1967-1972.

The model is specified with a view to testing several hypotheses
that have been advanced in the literature respecting the economic and
institutional determinants of intertemporal variations in the rate of
change of aggregate trade union membership in Canada. The parameter esti-
mates of the model were found to be broadly consistent with many of these
hypotheses. More specifically, strong empirical support was obtained for
the presence of a positive ceteris paribus effect on \( \Delta T_t \), the percentage
change in Canadian union membership, of each of the following explanatory
variables: (1) the current rate of change of aggregate money prices, the
effect of which appears to decrease linearly with increases in the rate of
price inflation; (2) current and recent past rates of change in aggregate
unionizable employment; (3) a variable representing the occurrence and
severity of business cycle contractions, which is tentatively interpreted
as a proxy measure of worker discontent; and (4) the current percentage
change in aggregate trade union membership in the United States. In addition, the parameter estimates confirmed the existence of: (5) a negative effect on $\Delta T_\ell$ of union density, the level of union membership as a percentage of unionizable employment, as postulated by the saturation hypothesis; and (6) a negative effect on $\Delta T_\ell$ of the two-year lagged value of the annual percentage change in strike starts, in direct contradiction of the interpretation of this variable as a proxy measure of union militancy. Finally, the results also indicated significant differences in either the intercept parameter or the coefficient of the union density variable between the sub-periods 1925-1943 and 1944-1966; they therefore provide indirect evidence that the growth rate of Canadian union membership was higher after the rights of private-sector workers to organize and bargain collectively were legally guaranteed and enforced by the provisions of Order-in-Council P.C. 1003 in 1944, but leave unresolved the question of how precisely these parameter shifts should be modelled and interpreted.

The tests for various types of specification error reported in Section 5 generally corroborate the specification of the proposed model; in particular, they fail to yield any evidence of autoregressive errors, non-constant parameters or non-zero error means. The tests of model specification reported in Section 6 imply that both principal variants of the proposed model provide valid representations of the estimation period data when compared with the empirical performance on the same data of two alternative models suggested by earlier econometric studies of union growth. The investigation of the ex post forecasting performance of the estimated models, the findings of which were reported in Section 7, casts considerable doubt on the post-sample predictive accuracy of the estimating
equations. Despite their evident success in adequately representing the growth rate of Canadian union membership over the estimation period 1925-1966, the variants of the proposed model do not appear to explain satisfactorily the data for the forecast period 1967-1972. Thus, at least insofar as post sample predictive accuracy is concerned, the conclusion of the present study might well be similar to that arrived at by Moore and Pearce (1976) in their analysis of Ashenfelter and Pencavel's (1969) original model of U.S. union growth — that "students of the labor movement interested in predicting future changes in union membership would be advised to find an alternative model" (Moore and Pearce, (1976, p. 247)).

With the notable exception of the findings pertaining to its post-sample forecasting accuracy, the estimates and tests of the present model of Canadian union growth for the period 1925-1966 are generally quite supportive of the model's empirical validity. Nevertheless, there are several limitations that warrant mention in closing; these pertain to certain aspects of the present model's specification and to the nature and interpretation of the empirical results.

1. The evidence of post-sample predictive failure for the period 1967-1972 raises the question of how to model more satisfactorily recent Canadian union growth experience during the late 1960's and the 1970's, the period in which union recognition and collective bargaining in the public sector was governed by the federal Public Service Staff Relations Act of 1967 and by roughly comparable provincial statutes. In this connection, further analysis of the performance of this and other models on data for more recent years would clearly be worthwhile.
2. As mentioned earlier in Section 4, the implausibly large estimates of the intercept shift coefficient in equations (1A) and (1A.1) raise doubts about both the appropriateness of the dummy variable specification of the intercept shift and the validity of its interpretation as an indicator of the effects on Canadian union growth of the labor relations framework established by P.C. 1003 in 1944. Alternative specifications of the intercept shift, such as those suggested by the transition function approach to modeling structural change, would seem to offer potentially fruitful alternatives to the dummy variable specification employed here.

3. Like most previous econometric investigations of union growth, the present study attempts to explain year-over-year percentage changes in the total stock of union members. Yet several of the hypotheses incorporated in this model may reasonably be interpreted as pertaining to the gross inflow of new members to this stock and/or the gross outflow of existing members from this stock. It would therefore be useful to test the ability of the present model to explain annual gross flows into and out of union membership in Canada. Such undertakings might attempt to model time series variations in measures of organizing activity (e.g., the number of employees covered by applications for new certification or re-certification), measures of organizing success (e.g., the number of employees covered by new certifications), and measures of organizing failure (e.g., the number of employees covered by successful applications for de-certification). Efforts to econometrically model the determinants of the certification/de-certification process have been undertaken recently on U.S. data by Adams and Krislov (1974), Roomkin and Juris (1978) and Anderson, O'Reilly and Busman (1980). Although comprehensive data, either
time series or cross section, on certifications and de-certifications are not yet presently available for Canada, there are indications that such data can be assembled in a form suitable for econometric analysis. The availability of such data would afford an opportunity to subject to more rigorous empirical testing various hypotheses about the determinants of annual flows into and out of union membership, and hence would constitute a potentially useful contribution to understanding the dynamics of the union growth process.

4. In order to test hypotheses respecting the effects on union membership growth of alleged causal factors such as worker discontent and union militancy, it has been necessary to specify proxy variables to represent indirectly these factors in the estimating equations. Unfortunately but unavoidable, the relationship of many of these proxy variables to the underlying factor each is intended to measure is at best rather tenuous and speculative. Obviously, alternative proxy measures could always be devised and, despite the fact that non-nested hypothesis testing procedures hold out some promise of permitting investigators to distinguish empirically among alternative proxy measures of a given factor, there inevitably will remain considerable ambiguity over the interpretation to be given to estimated relationships between union membership growth on the one hand, and these proxy variables on the other. Thus, for example, it is a very long leap from the finding that the proxy variable \( DD_t(U_t - U^*) \) is empirically related to \( \Delta T_t \) to the conclusion that worker discontent is an important determinant of variations in union membership growth in Canada.
5. Like most previous econometric investigations of union growth, the present analysis has been conducted at a very high level of aggregation. It would be desirable for several reasons to extend this type of analysis of the sources of union growth to the level of the industry division and/or geographic region. For one thing, the process of union organization has proceeded at very different rates in different sectors and regions of the Canadian economy. For another, the existence of marked disparities in economic structure and growth among various industries and regions makes it highly unlikely that the responses of union membership growth to changes in its measurable determinants are constant across industrial sectors and geographic regions. Furthermore, investigations undertaken at the level of the industry division or economic region are more likely to be able to account for economic, institutional, legal and political factors that are specific to individual sectors than are investigations conducted at higher levels of aggregation.\textsuperscript{53} Finally, some of the hypotheses about the determinants of $\Delta T_r$, such as the saturation hypothesis, are probably more appropriately tested at the industry or regional level. Needless to say, one's confidence in the findings of this and other studies of aggregate union membership growth would be strengthened if these findings could be substantiated (at least qualitatively) at the level of the individual industrial sector or economic region.

6. In view of the imprecision with which ad hoc "theories" of union growth have been formulated in the literature, there is bound to be considerable scope for disagreement among econometric investigators over the precise empirical specifications to be given these theories for purposes of empirical testing. Indeed, the likelihood of such disagreement emphasizes
the tentative nature of the findings gleaned from the present study.

These limitations and shortcomings of the present study underscore the fact that much more work remains to be done before there can be any hope of reaching a consensus concerning the empirical formulation, interpretation and importance of the measurable determinants of Canadian trade union growth. Attempts to remedy these limitations constitute a substantial prospective agenda for future research.
FOOTNOTES

1. For an excellent summary and critique of much of this literature, see Blum (1968).

2. At least two other categories of econometric investigations of union membership can be identified. The first comprises econometric studies that use primarily cross-sectional data to model variations across industries, establishments or individuals in the level or frequency of union membership; included in this category are the studies by Blinder (1972), Moore and Newman (1975), Freeman and Medoff (1976) and Bain and Elsheikh (1979). The second includes econometric analyses that use time series and/or cross section data in an effort to model various measures of the results of the union representation process; examples of recent studies in this category are those by Adams and Krislov (1974), Roomkin and Juris (1978), and Anderson, O'Reilly and Busman (1980), as well as the innovative analysis by Farber and Saks (1980), all of which employ U.S. data on certifications and decertifications resulting from union representation votes.

3. Bain and Elsheikh have also applied variants of their general model to aggregate time series data for the United States, Australia and Sweden; see Bain and Elsheikh (1976, 1978, 1979).

4. In this connection, see the criticisms of the Swidinsky (1974) model by Bain and Elsheikh (1976), the note on the original Ashenfelter-Pencavel (1969) model by Moore and Pearce (1976), and the comments on the Bain-Elsheikh (1976) model by Richardson (1977, 1978), Pederson (1978) and the original authors (1978).

5. See Labour Canada (1975), Labor Organizations in Canada, Table 1.

6. Although this sample period completely contains the era in which industrial unionism became a major presence in Canada, it excludes one period of extremely rapid union growth — namely, the years 1916 to 1919 — as well as the immediately ensuing period of membership decline — the years 1920-1924. Regrettably as these exclusions are, it was judged better to confine the analysis to the post-1920 period than to resort to the questionable practice of using either proxy data measures or interpolation methods in order to extend the available time series for other variables back to 1911, the first year for which official union membership figures are available.

8. Woods (1973, pp. 344-345) succinctly summarizes the major innovative provisions of Order-in-Council P.C. 1003 as follows:

[P.C. 1003] added a clear-cut policy that unions had a right to exist, that employees had a right to join unions, that unions should be free from employer domination or influence, that the denial of such rights was an unfair practice, and that a union with the support of a majority of workers in a unit appropriate for collective bargaining could impose collective bargaining on employers without the need to resort to the strike. Conversely, the employer was deprived of the former freedom to use his control over jobs to prevent employees from joining unions. Dismissal for union activity—a threat to the individual workman—and the lockout—a threat to the group as a whole—were no longer legally available to keep the employee out of unions or to destroy unions once they were established.

9. Moreover, by 1948, the application of this policy was, with few exceptions, more or less universal across the country; see Woods (1973, p. 347).

10. As Woods has observed of the PSSRA, "it was intended that this law should do for the public sector what the Industrial Relations and Disputes Investigation Act was intended to do for the private sector." (Woods (1973, p. 297)).


12. For example, to quote Ulman (1955a) "One is led to expect that membership changes would be responsive to cyclical variations in business activity, other things being equal, since such cyclical swings have affected both the propensity and the opportunity to form or join unions" (p. 237). Ulman (1955a) goes on to say that "if one observes movements in the rate, as well as the direction, of change, variations in union membership do conform to (at least) major fluctuations in business activity." (p. 239). J. T. Montague (1970), in his textbook on Canadian labor economics, emphasizes the universality of business cycle theories of union growth: "Throughout all the studies of union growth is the recurring theme that union membership is bound to vary according to the course of the business cycle and—as it was expressed at an early date—union membership would be the plaything of the business cycle" (p. 157).

13. Indeed, most of the differences among various econometric models of union growth center not on the identity of those general factors thought to influence the pattern of union membership changes, but rather on the selection and specification of measurable proxy variables to represent these factors empirically.
14. On the one hand, the rate of change of unionizable employment serves principally as a direct measure of cyclical variations in the overall employment conditions facing present and prospective union members; in this role, its economic meaning and interpretation would seem to have changed little over time. The aggregate unemployment rate, on the other hand, serves primarily as an inverse index of cyclical variations in labor market tightness. But pronounced secular changes in the demographic composition of the labor force, as well as the growth of public income transfer programs (e.g., unemployment insurance, welfare programs, and minimum wages) appear to have substantially altered the economic meaning of the officially measured aggregate unemployment rate as an indicator of the degree of labor market tightness, presumably because these demographic and public policy changes have mainly affected the structural, rather than the cyclical, component of the official unemployment rate. There now exists a substantial body of empirical evidence indicating that a given degree of labor market tightness has been associated with higher values of the measured unemployment rate during the 1960's and 1970's than during the earlier decades of this century; see, for example, the U.S. studies by Perry (1970) and Wachter (1976) and the evidence for Canada given in Fortin and Phaneuf (1981). This evidence suggests that the economic interpretation of movements in the measured aggregate unemployment rate as an indicator of cyclical variations in labor market tightness has changed rather substantially over sample periods as long as those on which this and other econometric studies of union growth have been based. It is therefore tempting to conclude that the measured aggregate unemployment rate is an empirically less reliable indicator over time of cyclical variations in labor market tightness than the percentage change in total unionizable employment is of cyclical variations in overall employment conditions. At the very least, the evidence tends to support the following observations by D. C. Smith (1980) on the Bain-Elsheikh (1976) model:

...Bain and Elsheikh's reliance on the unemployment rate as the indicator of employment conditions is worrisome. Questions can be raised both about the reliability of the data on unemployment rates for the early decades of this century and about changes over time in their economic significance as an indicator of labor market tightness. Moreover, since it is the rate of change of union membership that is being investigated, the inclusion of the rate of change of employment as an additional variable merits further consideration (pp. 105-106).

However, the rate of change of employment may not be as reliable an index of labor market conditions as the foregoing observations would seem to suggest, primarily because the same demographic changes in the age-sex composition of the labor force that have altered the interpretation of the unemployment rate have also affected the composition of aggregate non-agricultural employment. In particular, the
substantial growth in the employment and labor force participation of younger persons and married women has increased the fraction of part-time workers in total non-agricultural employment. Hence, a given percentage change in total non-agricultural employment has likely corresponded to a smaller percentage change in full-time employment (and in total manhours worked) during recent years than it did during the years before and immediately after World War II. Thus, the same demographic compositional changes that have altered the relationship between the aggregate unemployment rate and labor market tightness have also altered somewhat the economic interpretation of percentage changes in unionizable employment as an indicator of variations in the employment of present and potential union members.

15. In their original study of American trade union growth, Ashenfelter and Pencavel (1969, p. 443) define unionizable employment to include employment in mining, manufacturing, construction, and transport and utilities. There are basically two reasons for adopting non-agricultural paid employment as the measure of unionizable employment in the present study. First, as Kreuger (1971, p. 952) observes, "figures on non-agricultural paid employment provide the best available approximation of the potentially organized." Second, a broader and more comprehensive measure of unionizable employment than that used by Ashenfelter and Pencavel seemed appropriate in view of the inclusion in the sample period of the late 1960's and early 1970's, when much of the increase in union membership in Canada occurred in the public and private service sectors, areas in which unionization has historically been very limited.

16. This hypothesis of a positive relationship between price changes and union growth was advanced (and empirically documented) by H. B. Davis (1941), who suggested that the positive correlation between price changes and union growth could be explained by the fact that joining or forming a union is in part a defensive response by workers to perceived threats to or reductions in their real wages arising from increases in consumer prices.

Davis (1941) contended that "(m)ost unions seem to have come into being originally as defensive organizations, to preserve a standard already enjoyed" (p. 611). He cited evidence "which seems to indicate that changes in union membership correlate more closely with sharp changes in prices than with 'prosperity'" (p. 617). Davis offered the following reasons for expecting a positive relationship between price changes and union growth: "In the first place, increases in wage rates tend to lag behind price rises, so that fully employed workers must organize if they are to avoid a drop in their standard of living. To that extent, ...unionization moves in times of rising prices are defensive movements. In the second place, a period of rapidly rising prices is nearly (but not quite) always a period of increasing production and decreasing unemployment, so that the favorable factors of prosperity will apply. ...Finally, a period of rising prices is a favorable time for employers to pass
along increases in cost to the customer in the form of higher prices, and to that extent the antagonism between capital and labor decreases in intensity" (pp. 617-618).

17. On the one hand, if, as Davis (1941) contended, unions are defensive organizations which attempt to protect the real wages of workers by increasing their recruiting efforts in response to increases in the rate of price inflation, and if workers themselves are more receptive to such efforts when prices are rising rapidly, then one might expect the rate of union growth to accelerate with increases in the rate of price inflation. This line of reasoning would lead one to expect a positive coefficient for $\Delta P_t^2$. On the other hand, however, there exists substantial evidence that high rates of inflation tend to reduce the average relative wage advantage of unionized workers. For example, Lewis' (1963, pp. 5-6) seminal analysis of U.S. data led him to conclude that "during 1920-1958, abnormally high rates of inflation strongly tended to reduce—and abnormally low rates of inflation to increase—the effect of unionism on the average union/non-union relative wage"; see also Chapter VI of Lewis (1963, pp. 195-222). To the extent that unorganized workers perceive the erosion of the relative union/nonunion wage as a reduction in one of the potential benefits of union membership, they may not be as inclined to join a union as they would be if prices were rising less rapidly and the real wage gains from union membership were seen to be greater. This latter argument implies that, ceteris paribus, the effect on $\Delta T_t$ of $\Delta P_t$ will be smaller the higher the rate of inflation, and therefore that the coefficient on $\Delta P_t^2$ should be negative. On balance, therefore, it is not possible to unambiguously predict the sign of the coefficient on $\Delta P_t^2$.

A somewhat similar argument to that given here for a quadratic relationship of $\Delta P_t$ with $\Delta T_t$ is advanced by Bain and Elsheikh (1976) to justify the specification of a threshold-type step discontinuity in the effect of $\Delta P_t$ on the rate of union membership growth. In particular, Bain and Elsheikh define a variable $\Delta PS_t = \Delta P_t \cdot S_t$, where $S_t = 0$ for all years in which $\Delta P_t < 4$ percent and $S_t = 1$ for all years in which $\Delta P_t \geq 4$ percent. They include the variable $\Delta PS_t$ in their model on the expectation that its coefficient will be negatively-signed—i.e., that the effect of $\Delta P_t$ on $\Delta T_t$ will be smaller when $\Delta P_t$ equals or exceeds 4 percent per year than when $\Delta P_t$ is less than 4 percent per year. Although this expectation is supported by their empirical results, the present model opts for a second degree polynomial specification of the relationship between $\Delta P_t$ and $\Delta T_t$, for two reasons. First, like the threshold specifications, the quadratic specification allows the effect of $\Delta P_t$ on $\Delta T_t$ to vary with the value of $\Delta P_t$, though in a continuous rather than a discontinuous manner. Second, the choice of one particular threshold value for $\Delta P_t$ is necessarily somewhat arbitrary, and suffers from the additional disadvantage that perceptions of what constitutes a "high" rate of inflation are likely to change over sample periods as long as those for which models of union growth are typically estimated.
18. As Ashenfelter and Pencavel (1969, p. 438) point out, this hypothesis has been succinctly stated by Hines (1964, p. 229): "As membership increases, there is a diminishing response to a given intensity of recruiting effort."

19. The function $\lambda(t-\theta) U^P_t$ assumes a new (higher) value in the trough of each recession when $(t-\theta) = 0$. For $\lambda = 1$, the function $\lambda(t-\theta) U^P_t = U^P_t$ for all $t$, and thus is a conventional step function which assumes a new value in the trough of each contraction. For values of $\lambda$ such that $0 < \lambda < 1$, the value of the function $\lambda(t-\theta) U^P_t$ declines monotonically between cyclical troughs. Finally, for $\lambda = 0$, the function $\lambda(t-\theta) U^P_t$ equals $U^P_t$ when $(t-\theta) = 0$, and zero when $(t-\theta) > 0$.

20. An incidental advantage of using the regressor $\Delta U_t (U_t - U^*)$ rather than the function $\lambda(t-\theta) U^P_t$ to proxy worker discontent is that the former avoids the nonlinearity in parameters associated with the parameter $\lambda$.

21. In his classic work on the history of the Canadian union movement, H. A. Logan (1948, pp. 4-5) artfully expresses this argument as follows:

Very important has been the effect of the United States. Not only has she been a foremost investor but has in many instances sent control along with capital. With her great population so near at hand, her export of books, magazines, newspapers, movie films and radio programmes, the interplay of migrants and tourists, and finally the organizational tie up of labour itself, she has permitted Canada to escape little of what interests and engages her. She shares her 'isms', her phobias, her price swings, her prosperity and her depressions... Canada also is stimulated by her labour legislation... and tends to follow it at a distance—though in some important instances Canada has led the way.

22. For reasons similar to those given here, Swidinsky (1974, p. 441) also includes the rate of change of U.S. union membership in his model of Canadian union growth. However, it is not clear from Swidinsky's paper whether he includes or excludes Canadian membership in American unions in constructing the series he uses to measure percentage changes in American union membership.

An additional point about the role of American union membership growth as a regressor in the present model warrants mention. The
notion of economic and institutional interdependence between Canada and the United States suggests several alternative approaches to econometric modelling of union growth in the two countries. The present single-equation model of Canadian union growth might be embedded in a two-equation recursive model in which both $\Delta T_U$ and $\Delta T_{US},_t$ are treated as endogenous variables, with $\Delta T_{US},_t$ appearing as a predictor in the $\Delta T_U$ equation but not vice versa. Alternatively, $\Delta T_U$ and $\Delta T_{US},_t$ might be treated as jointly-determined variables in a fully simultaneous, two-equation model; however, the plausibility of this approach is impaired somewhat by the appearance of $\Delta T_U$ as an endogenous regressor in the equation determining $\Delta T_{US},_t$, since such a specification would be akin to "the tail wagging the dog."

23. It might be noted in this connection that D. C. Smith (1980), in his recent re-evaluation of trade union growth in the United Kingdom, strongly suggests that government action has had an appreciable effect on British union growth, especially after the mid-1960's, but also perhaps during the years of the World War I and World War II periods. Smith's (1980, pp. 122-123) estimates of a modified model of U.K. union growth appear to provide empirical support for the proposition that government action has at times had a positive independent effect on union membership growth in the United Kingdom.

24. See, for example, Bain (1978, pp. 25-26).

25. In their original study, Ashenfelter and Pencavel (1969) use a variable measuring the percentage of major party membership in the House of Representatives which is affiliated with the National Democratic Party, on the grounds that the long-standing association between organized labor and the Democratic Party in the United States makes it "sensible to suppose that periods in which the Democratic Party predominates should reflect a high degree of pro-labor sentiment" (Ashenfelter and Pencavel (1969, p.439)). In a comment on the Ashenfelter-Pencavel study, Manke (1971) argued that a more direct way to allow for government attitudes and actions respecting union recognition is to employ intercept shift dummy variables to capture the effect on union growth of the enactment of major pieces of labor legislation — in particular, the effect of the National Labor Relations Act (or Wagner Act) of 1935 and its subsequent constitutional validation by the United States Supreme Court in 1937. In Manke's (1971, p. 190) words, "a much better way to take account of the political factor is by explicitly introducing into the model that is to be estimated dummy variables denoting the major 'political' events thought to have benefitted or harmed union membership." Manke's re-estimation of the Ashenfelter-Pencavel model provides strong support for the proposition that American union membership growth was higher after the Wagner Act passed its federal court test in 1937 than it otherwise would have been.

26. It is worth pointing out that an index of public and government attitudes towards unions similar to Ashenfelter and Pencavel's (1969)
variable (the percentage of members in the House of Representatives that is affiliated with the National Democratic Party) is clearly inappropriate in the Canadian context. In contrast to the situation that prevails in both the United States and the United Kingdom, organized labor in Canada has not consistently supported any of the three major federal political parties, including the New Democratic Party (NDP) and its predecessor, the Cooperative Commonwealth Federation (CCF). In his account of the history of the Canadian labor movement's political relations, Miller (1971, pp. 229-230) offers the following summary observations on Canadian labor's political alliances: "Canadian labour has by no means restricted its favours to a single party such as the NDP. On the contrary, both trade union leaders and rank-and-file members have frequently aligned themselves with the Liberals, the Progressive Conservatives, and any number of minor parties. While this support has generally fallen short of direct union affiliation, participation on an individual basis has been open and widespread."

27. For equations (1A) and (1B), the appropriate 1 percent critical value of the F-distribution is $F_{0.01}(12, 29) = 2.87$; for equations (1A.1), (1A.2), (1B.1) and (1B.2), the appropriate 1 percent critical value is $F_{0.01}(11, 30) = 2.90$.

28. About 85 percent of the parameter estimates in Tables 1 and 2 are statistically significant on a two-tailed test at the 1 percent test level, while an additional 10 percent are significant at the 5 percent test level.

29. For example, when the lag coefficients on $\Delta E_{t-1} (i=0, 1, 2, 3)$ were constrained to lie on a second-degree Almon polynomial, a conventional F-test of the implied parameter restriction indicated decisive rejection at the 1 percent significance level for both equations (1A) and (1B). The sample values of the F-ratio for the null hypothesis that the lag coefficients on $\Delta E_{t-1} (i=0, 1, 2, 3)$ conform to a second-degree Almon polynomial are $F(1, 29) = 24.843$ for equation (1A) and $F(1, 29) = 25.241$ for equation (1B), both of which far exceed the 1 percent critical value $F_{0.01}(1, 29) = 7.60$.

30. This finding differs from the formulation adopted by Ashenfelter and Pencavel (1969) in their original model of U.S. union growth, which specified a four-period, second-degree Almon polynomial lag with the far endpoint restriction imposed.

31. In addition, the parameter estimates for $\Delta P_t$ and $\Delta P_t^2$ in Tables 1 and 2 suggest the existence of a threshold value of the inflation rate equal to about 5.0 percent: for values of the inflation rate above this threshold, the effect of $\Delta P_t$ on $\Delta T_t$ is statistically negligible, while for inflation rates below this threshold, the effect of $\Delta P_t$ on $\Delta T_t$ is positive and statistically significant. But unlike the step discontinuity in the effect of $\Delta P_t$ on $\Delta T_t$ specified in the Bain-Elsheikh (1976) model, the quadratic specification of the relationship between $\Delta T_t$ and $\Delta P_t$ avoids the need to select the threshold value of
\( \Delta P_t \) a priori, and instead yields the threshold value as an implication of the estimated quadratic relationship.

32. In equation (1A), which incorporates the linear specification of the union growth-union density relationship, the effect of \( D_{t-1} \) on \( \Delta D_t \) is given by \((\partial \Delta T_t / \partial D_{t-1})_{1925-43} = a_4 \) for the 1925-1943 sub-period, and by \((\partial \Delta T_t / \partial D_{t-1})_{1944-66} = a_4 + a_9 \) for the 1944-1966 sub-period. Thus, the coefficient estimates (and corresponding t-ratios) for \( D_{t-1} \) implied by the parameter estimates in Table 1 are as follows:

<table>
<thead>
<tr>
<th></th>
<th>1925-1943</th>
<th>1944-1966</th>
</tr>
</thead>
<tbody>
<tr>
<td>Equation (1A)</td>
<td>-0.801</td>
<td>-0.858</td>
</tr>
<tr>
<td></td>
<td>(-3.334)</td>
<td>(-4.688)</td>
</tr>
<tr>
<td>Equation (1A.1)</td>
<td>-0.837</td>
<td>-0.837</td>
</tr>
<tr>
<td></td>
<td>(-5.834)</td>
<td>(-5.834)</td>
</tr>
<tr>
<td>Equation (1A.2)</td>
<td>-1.091</td>
<td>-0.497</td>
</tr>
<tr>
<td></td>
<td>(-4.936)</td>
<td>(-4.713)</td>
</tr>
</tbody>
</table>

33. In equation (1B), which incorporates the nonlinear (i.e., reciprocal) specification of the union growth-union density relationship, the effect of \( D_{t-1} \) on \( \Delta T_t \) is given by \((\partial \Delta T_t / \partial D_{t-1})_{1925-43} = \beta_4 \) for the 1925-1943 sub-period, and by \((\partial \Delta T_t / \partial D_{t-1})_{1944-66} = \beta_4 + \beta_9 \) for the 1944-1966 sub-period. Thus, the estimates of \( \partial \Delta T_t / \partial D_{t-1} \) (and corresponding t-ratios) implied by the coefficient estimates in Table 2 are as follows:

<table>
<thead>
<tr>
<th></th>
<th>1925-1943</th>
<th>1944-1966</th>
</tr>
</thead>
<tbody>
<tr>
<td>Equation (1B)</td>
<td>178.827</td>
<td>694.377</td>
</tr>
<tr>
<td></td>
<td>(3.247)</td>
<td>(4.859)</td>
</tr>
<tr>
<td>Equation (1B.1)</td>
<td>242.083</td>
<td>242.083</td>
</tr>
<tr>
<td></td>
<td>(4.018)</td>
<td>(4.018)</td>
</tr>
<tr>
<td>Equation (1B.2)</td>
<td>230.097</td>
<td>537.099</td>
</tr>
<tr>
<td></td>
<td>(5.273)</td>
<td>(5.520)</td>
</tr>
</tbody>
</table>

34. In particular, equation (1B) implies that the ceteris paribus effect of \( D_{t-1} \) on \( \Delta T_t \) is given by \((\partial \Delta T_t / \partial D_{t-1})_{1925-43} = -\beta_4 / D_{t-1}^2 \) for the 1925-1943 sub-period, and by \((\partial \Delta T_t / \partial D_{t-1})_{1944-66} = -(\beta_4 + \beta_9) / D_{t-1}^2 \) for the 1944-1966 sub-period.
35. Similar evidence of a statistically significant positive effect of U.S. union growth on Canadian union growth is obtained by Swidinsky (1974) and by Bain and Elsheikh (1976).

36. In order to investigate the hypothesis of a finite distributed lag effect of $\Delta TU_S$ on $\Delta T$, equations (1A) and (1B) were re-specified to include an unrestricted four-period distributed lag on $\Delta TU_S, t-1$ for $i = 0,1,2,3$. However, the estimates of these re-specified equations for the period 1925-1966 indicated that in neither case was it possible to reject the null hypothesis that the lag coefficients on $\Delta TU_S, t-1, \Delta TU_S, t-2$ and $\Delta TU_S, t-3$ jointly equal zero: the sample values of the conventional F-test statistic under the null hypothesis were $F(3,26) = 0.396$ for equation (1A) and $F(3,26) = 0.371$ for equation (1B), both of which are much less than the 5 percent critical value $F_{0.05}(3,26) = 2.98$. Moreover, none of the estimated lag coefficients on $\Delta TU_S, t-1, \Delta TU_S, t-2$ and $\Delta TU_S, t-3$ in either equation were individually significant on a two-tailed test at even the 50 percent significance level.

37. When tested against an unrestricted three-year distributed lag on $\Delta S_{t-1}$ for $i = 1,2,3$, the null hypothesis that the lag coefficients on $\Delta S_{t-1}$ and $\Delta S_{t-3}$ are jointly zero yields sample values of the F-test statistic of $F(2,27) = 0.0322$ for equation (1A) and of $F(2,27) = 0.0374$ for equation (1B); since the 5 percent critical value is $F_{0.05}(2,27) = 3.35$, the null hypothesis obviously cannot be rejected even at very high significance levels. When equation (1A) is estimated with an unrestricted three-year distributed lag on $\Delta S_{t-1}$ for $i = 1,2,3$, the coefficient estimates and (t-ratios) are $-0.00145$ (-0.0927) for $\Delta S_{t-1}$, $-0.0440$ (-3.297) for $\Delta S_{t-2}$ and $-0.00328$ (-0.246) for $\Delta S_{t-3}$. Similar results are obtained when equation (1B) is estimated with a distributed lag on $\Delta S_{t-1}$ for $i = 1,2,3$: the coefficient estimates (and t-ratios) are $-0.00267$ (-0.170) for $\Delta S_{t-1}$, $-0.0449$ (-3.377) for $\Delta S_{t-2}$ and $-0.00308$ (-0.231) for $\Delta S_{t-3}$. Thus, the estimated lag coefficients on $\Delta S_{t-1}$ and $\Delta S_{t-3}$ are all individually insignificant, while the estimates of the lag coefficient on $\Delta S_{t-2}$ are virtually identical, both in magnitude and in relation to their respective standard errors, to those in Tables 1 and 2. Thus, the fixed two-year lag specification appears to capture the dependence of Canadian union growth on past percentage changes in aggregate strike starts.

38. A conventional F-test of the restrictions that the coefficients of $D_{4t}$ and $D_{4t} \cdot D_{t-1}$ in equation (1A) jointly equal zero yields a test statistic sample value of $F(2,29) = 15.690$; a similar test of the restrictions that the coefficients of $D_{4t}$ and $D_{4t} \cdot D_{t-1}$ in equation (1B) jointly equal zero yields a test statistic sample value of $F(2,29) = 16.054$. Since the 1 percent critical value of the test statistic is $F_{0.01}(2,29) = 5.42$, the coefficients of the intercept shift dummy variable and of the dummy variable-union density interaction term are easily jointly significant at the 1 percent test level in both equations.
39. The estimates of $\frac{\partial \Delta T_t}{\partial D_{t-1}}$ from equations (1A) and (1A.1) are, of course, independent of the value of $D_{t-1}$; they neither confirm nor refute the hypothesis of an algebraically larger value of $\frac{\partial \Delta T_t}{\partial D_{t-1}}$ in the 1944–1966 period compared with the 1925–1943 period. Note too that both equations (1A.2) and (1B.1) yield significantly larger mean estimates of $\frac{\partial \Delta T_t}{\partial D_{t-1}}$ for the later as compared with the earlier period; however, these estimates are here ignored because they are derived from estimating equations that incorporate an invalid parameter restriction.

40. Such an alternative approach is not pursued here, but would seem in principle to warrant further investigation; see, for example, the paper by McAleer and Gregory (1979).

41. The far endpoint constraint restricts the lag coefficient on $\Delta \varepsilon_{t-4}$ in equations (2A) and (2B) to equal zero. Together with the specification of a second-degree polynomial lag, this lag structure imposes two independent linear restrictions on the parameters of the unrestricted four-period distributed lag on $\Delta \varepsilon_{t-1}$ for $i = 0, 1, 2, 3$. In contrast to the findings for equations (1A) and (1B), these Almon restrictions on the parameters of the distributed lag on $\Delta \varepsilon_{t-1}$ in equations (2A) and (2B) are not rejected by the data. A conventional F-test of the two independent Almon restrictions yields sample values of $F(2, 33) = 1.142$ for equation (2A) and $F(2, 33) = 0.929$ for equation (2B), both of which are much less than the 5 percent significance value $F_{0.05}(2, 33) = 3.29$.

42. Ashenfelter and Pencavel's (1969) proxy measure is DEM$_t$, the percentage of major party membership in the U.S. House of Representatives that is affiliated with the National Democratic Party.

43. To handle the nonlinearity in parameters introduced by the function $\lambda(t-\omega)u^p$, a grid search procedure is employed whereby the sum of squared OLS residuals is calculated for alternative values of $\lambda$ in the interval $0 < \lambda < 1$; the chosen value of $\lambda$ is that which minimizes the sum of squared OLS residuals over the closed unit interval.

44. The significance points for the lower and upper bounds of $d_1$ are given in Savin and White (1977).

45. Because they require only estimation of the postulated model under the null hypothesis, both procedures may be interpreted as applications of the Lagrange Multiplier Principle of hypothesis testing; they are therefore computationally easy to perform and readily permit the null hypothesis to be tested against higher-order autoregressive error specifications without having to estimate the model under any of these alternative error specifications.

46. This practice amounts to setting the pre-sample values of the OLS residuals to zero, a convention which leaves unaltered the asymptotic properties of the tests.
47. In particular, the sample values of the two test statistics are related according to the equality

\[ F^*(n) = \frac{(T-k-n)}{T} \cdot \frac{\bar{\alpha}^t \bar{\alpha}}{\bar{\phi}^t \bar{\phi}} \cdot M(n) \]

where \((T-k-n)/T < 1\) and \(\bar{\alpha}^t \bar{\alpha}/\bar{\phi}^t \bar{\phi} > 1\).

48. Note, however, that the relatively small sample size of 42 observations (of which about 1/3 are used up in estimating the artificial regression equations corresponding to (1A) and (1B)) and the asymptotic nature of the tests suggest that the true size of each test may well exceed, perhaps substantially, the nominal size. Since the \(F^*(n)\) statistic incorporates a degrees-of-freedom correction while the \(M(n)\) statistic does not, the Durbin test based on \(F^*(n)\) may have true size closer to its nominal size than the LM test based on \(M(n)\) when the sample size is as small as it is in the present case.

49. Incidentally, the same tests applied to the modified Ashenfelter-Pencavel equations (2A) and (2B) also indicate non-rejection of the null hypothesis of AR(0) errors against both alternative autoregressive error structures at nominal significance levels of 10 percent or higher.

50. For details on the definition and properties of recursive residuals, see Brown, Durbin and Evans (1975).

51. The sequential Chow test and the cusum-of-squares test are distinctly different test procedures, and therefore may yield different inferences concerning the truth of \(H_0\) in any particular application. However, it is straightforward to demonstrate that the test statistics of the two procedures are related according to the equality

\[ F_r = \frac{(r-k)}{(T-r)F_T + (r-k)} \quad , \quad r = k+1, k+2, \ldots, T \]

See, for example, Fisher and McAleer (1980, pp. 14–15).

52. One important limitation of the RESET test is that it is not well defined for cases where the misspecification induces non-normality of the error vector in the misspecified model; see Ramsey (1969, p. 369).

53. Note that \(\gamma = 0\) in equation (5.7) corresponds to the null hypothesis \(H_0: E(u|X) = 0\), while \(\gamma \neq 0\) in equation (5.7) corresponds to the alternative hypothesis \(H_1: E(u|X) = \xi = 0\).

54. One important point concerning the application and interpretation of the J-test is that the artificial regression equation (6.3) is for-
mulated for the purpose of testing the truth of model $H_0$ and cannot be used directly to test the truth of the alternative model $H_1$, despite the fact that a value of $\alpha = 1$ in equation (6.3) implies that $H_1$ is true and $H_0$ is false. The reason for this is that the estimate of $\alpha$ and its variance are computed in (6.3) conditional on the assumed truth of $H_0$, not on the truth of $H_1$. Consequently, the estimated variance of $\hat{\alpha}$ from (6.3) will be inconsistent if $H_1$ is true. Although methods are available for obtaining a consistent estimate of the variance of $\hat{\alpha}$ under $H_1$, the simplest way to compute a consistent test of the truth of $H_1$ against $H_0$ is in general to reverse the roles of the two models and repeat the calculations of the J-test.

55. The differing interpretations given the outcomes of the J-test and F-test procedures stem from the fact that the former involves no maintained hypothesis, whereas the latter adopts the compound model $H_C$ as the maintained hypothesis. The F-test procedure thus in effect expands the set of alternative models under consideration from two to three, one of which—the compound model—is true by assumption. Nevertheless, when the parameter restrictions on this compound model implied by each of the original models $H_0$ and $H_1$ are tested in turn, the same four outcomes are available as in the case of the J-test: (1) reject $H_0$, but not $H_1$; (2) reject $H_1$, but not $H_0$; (3) reject neither $H_0$ nor $H_1$; (4) reject both $H_0$ and $H_1$. This means that, although the F-test and the J-test have similar possible outcomes respecting the validity of $H_0$ and $H_1$, the interpretation given these outcomes is quite different. For example, if $H_0$ is rejected but $H_1$ is not, then both $H_1$ and $H_C$ are accepted as being valid in the case of the F-test, whereas obviously only $H_1$ is accepted as being true in the case of the J-test. A decision to choose $H_1$ over $H_C$ on the basis of the F-test results does not follow from the truth of one and the falsity of the other, but rather from considerations relating to relative efficiency in estimation and the principle of parsimony in model selection.

56. On the one hand, if the error terms in equation (6.4) are normal and the regressors non-stochastic, then the F-test is exact in finite samples; the J-test, in contrast, is only asymptotically valid and so may be biased in small samples. On the other hand, the J-test, like all non-nested tests, is designed to have good power against true alternative models, although the F-test may have better power when both $H_0$ and $H_1$ are false. However, if the regressors in $X_*$ and $Z_*$ are highly collinear, as they frequently are in many applications, the F-test may have low power against either a true or a false alternative model.

57. A table of parameter estimates and summary statistics for equations (3A) and (3B) for the estimation period 1925-1966 is available from the author upon request.

58. The values of the maximized log-likelihood ($\ln L$), the coefficient of determination ($R^2$) and the standard error of the regression (s)
corresponding to the OLS estimates of equations (3A) and (3B) for the estimation period 1925-1966 are as follows:

<table>
<thead>
<tr>
<th></th>
<th>(\ln L)</th>
<th>(R^2)</th>
<th>(s)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(3A)</td>
<td>-105.693</td>
<td>0.783</td>
<td>3.607</td>
</tr>
<tr>
<td>(3B)</td>
<td>-106.057</td>
<td>0.779</td>
<td>3.638</td>
</tr>
</tbody>
</table>

59. A finding of predictive failure is sometimes, though by no means always, symptomatic of model misspecification, whereas a finding of predictive accuracy is perfectly consistent with the existence of model misspecification. Consequently, the results of tests of relative forecast accuracy are probably best viewed as indicators of a model's usefulness for the specific purpose of predicting out of sample, rather than as tests for model misspecification per se.

60. Let \(\hat{y}_t\) denote the predicted value, and \(y_t\) the actual or realized value, of the dependent variable \(\Delta T_t\) for year \(t\) of the forecast period. The mean prediction error (ME), the root-mean-square prediction error (RMSE), and Theil's (1966) revised inequality coefficient \((U_1)\) are defined as follows:

\[
ME = (1/T_2) \sum_{t=T+1}^{T+T} (\hat{y}_t - y_t)
\]

\[
= \bar{\hat{y}}_p - \bar{y}_A
\]

where \(\bar{\hat{y}}_p = (1/T_2) \sum_{t=T+1}^{T+T} \hat{y}_t\) and \(\bar{y}_A = (1/T_2) \sum_{t=T+1}^{T+T} y_t\) are, respectively, the forecast period sample means of the predicted and actual values of the dependent variable;

\[
RMSE = [(1/T_2) \sum_{t=T+1}^{T+T} (\hat{y}_t - y_t)^2]^{1/2}
\]

and
\[
U_1 = \left[ \frac{1}{T_2} \sum_{t=T+1}^{T+T_2} (\bar{y}_t - \bar{y}_t)^2 \right]^{1/2}
\]

For the perfect predictor, all three of these measures of forecast accuracy equal zero, i.e., \( ME = RMSE = U_1 = 0 \), while for any forecasting model other than the perfect predictor, \( ME \neq 0 \), \( RMSE > 0 \), and \( U_1 > 0 \).

6. The mean square prediction error \( MSE \) is defined as

\[
MSE = RMSE^2 = \frac{1}{T} \sum_{t=T+1}^{T+T_2} (\bar{y}_t - y_t)^2,
\]

where \( \bar{y}_t \) is the predicted value and \( y_t \) is the realized value of the dependent variable \( \Delta T_t \) in year \( t \) of the forecast period. Theil (1966) demonstrated that \( MSE \) could be decomposed into three additive components such that

\[
MSE = (\bar{y}_p - \bar{y}_A)^2 + (S_p - rS_A)^2 + (1-r^2)S_A^2
\]

where

\[
S_p = \frac{1}{(1/2)} \sum_{t=T+1}^{T+T_2} (\bar{y}_t - \bar{y}_p)^2 \]

and

\[
S_A = (1/2) \sum_{t=T+1}^{T+T_2} (y_t - \bar{y}_A)^2 \]

are the standard deviations of the predicted and actual values of \( y \), respectively; and

\[
r = \frac{[1/(1/2)] \sum_{t=T+1}^{T+T_2} (\bar{y}_t - \bar{y}_p)(y_t - \bar{y}_A)]}{S_p S_A}
\]

is the simple correlation coefficient between \( \bar{y}_t \) and \( y_t \) over the forecast period. The set of inequality proportions is obtained by dividing the above equation by \( MSE \):

\[
1 = \frac{(\bar{y}_p - \bar{y}_A)^2}{MSE} + \frac{(S_p - rS_A)^2}{MSE} + \frac{(1-r^2)S_A^2}{MSE}
\]

\[
= U_M + U_R + U_D
\]

\( U_M = (\bar{y}_p - \bar{y}_A)^2/MSE \) is denoted the bias proportion, and represents the proportion of the \( MSE \) that is attributable to the tendency of the forecasting model to overpredict or underpredict \( y_t \). \( U_R = \)
\( (S_p - rS_A)^2 / \text{MSE} \) is termed the regression proportion, and reflects the extent to which \( \partial y_t / \partial \theta_t \) systematically deviates from unity. \( U_0 = (1-r^2)S_A^2 / \text{MSE} \) is designated the disturbance proportion: it represents the fraction of the MSE that is attributable to non-systematic random errors.


63. This point has particular relevance in the Canadian context to the analysis of union growth on a provincial or regional basis, for at least two reasons. First, in accordance with the provisions of Sections 91 and 92 of the British North America Act of 1867, and with subsequent judicial interpretations of these constitutional provisions, principal authority to legislate in the sphere of labor relations has devolved to the individual provinces, which have exercised this authority in such a way as to produce far from uniform labor legislation. Second, largely owing to the unique and integral part played by the Roman Catholic Church, Quebec has long been regarded by students of the Canadian labor movement as a special case deserving of separate study. See, for example, the paper by Isbester (1971), in which he offers the following observations: "As ever, Quebec is different; while its unions, in common with unions in the rest of Canada, may be narrowly pragmatic, the milieu in which they have sought their ends, and the resultant method, has been at variance from the Canadian pattern" (Isbester (1971, p. 241)).
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DATA APPENDIX

A.1 Definitions and Sources of Data Series

1. \( T_t \) = total trade union membership in Canada in year \( t \) (in thousands).

Source: For all years 1921-1972, the data values for \( T_t \) are taken from Labour Canada, Labour Organizations in Canada 1974/1975, Catalogue No. L2-2/1975, Table 1.

2. \( P_t \) = all items Consumer Price Index for Canada in year \( t \) (1971 = 100).

Source: For all years 1921-1972, the data values for \( P_t \) are annual averages of the monthly figures in CANSIM database series B484000.

3. \( NAPWK_t \) = total number of non-agricultural paid workers in Canada in year \( t \) (in thousands).

Source: (1) For the years 1921-1930, the figures for \( NAPWK_t \) are those in series C52 of Urquhart and Buckley, Historical Statistics of Canada, p. 61.

(2) For the years 1931-1971, the figures for \( NAPWK_t \) are those given in Section 4, Table 3 of Statistics Canada, Canadian Statistical Review: Historical Summary 1970, Catalogue No. 11-505, p. 48.

(3) For the year 1972, the figure for \( NAPWK_t \) is that given in Section 4, Table 3 of Statistics Canada, Canadian Statistical Review, Catalogue No. 11-003 (January, 1975), p. 42.

Notes: (1) The figures for \( NAPWK_t \) for the years 1921-1945 refer to June 1st of each year; those for the years 1946-1952 are an average of 4 quarterly surveys; and those for 1953-1972 are an average of 12 monthly surveys.

(2) Estimates of non-agricultural paid employment are not available for years prior to 1931. Consequently, the figures for \( NAPWK_t \) for the years 1921-1930 are for total non-agricultural employment \( (NAGEM_t) \), which includes not only non-agricultural employees but also the number of self-employed persons and unpaid family workers in non-agricultural establishments.
4. \( NACEM_t \) = total non-agricultural employment in Canada in year \( t \) (in thousands).

Sources: Same as those given above for \( NAPWK_t \).

Notes: The figures for \( NACEM_t \) for the years 1921-1945 refer to June 1st of each year; those for the years 1946-1952 are an average of 4 quarterly surveys; and those for 1953-1972 are an average of 12 monthly surveys.

5. \( U_t \) = the aggregate civilian unemployment rate in Canada in year \( t \) (in percentage points); i.e., the total number of unemployed persons as a percentage of the total civilian labor force in year \( t \).

Sources: (1) For the years 1921-1930, the data values for \( U_t \) are equal to the ratio (in percentage points) of series C54, "persons without jobs and seeking work," to series C50, "total civilian labour force," as given in Urquhart and Buckley, Historical Statistics of Canada, p. 61.

(2) For the years 1931-1971, the data values for \( U_t \) are those given in Section 4, Table 3 of Statistics Canada, Canadian Statistical Review: Historical Summary 1970, Catalogue No. 11-505, p. 48.

(3) For the year 1972, the data value for \( U_t \) is that given in Section 4, Table 3 of Statistics Canada, Canadian Statistical Review, Catalogue No. 11-003 (January, 1975), p. 42.

Notes: (1) For the years 1921-1945, the estimates of \( U_t \) were made annually and refer to June 1st of each year; for the years 1946-1952, the annual figures for \( U_t \) are an average of 4 quarterly surveys; for the years 1953-1972, the annual figures for \( U_t \) are the average of 12 monthly surveys.

(2) The estimates of \( U_t \) for the years 1921-1945 exclude persons on temporary lay-off, while those for the years 1946-1972 include persons on temporary lay-off.
6. $DD_t$ = a binary dummy variable denoting years of cyclical contraction, which takes the value 1 for years of cyclical contraction and the value 0 for years of cyclical expansion.

Sources: (1) For the period 1921–1961, the reference cycle dates used to construct the series for $DD_t$ are taken from Table A-1.1 of D. A. White, *Business Cycles in Canada*, p. 236.

(2) For the period 1962–1972, the reference cycle dates used to specify the values of $DD_t$ are those given in C. Schwartz, *The Cyclical Momentum of Economic Activity in Canada 1953–1973*.

Notes: (1) Years during which cyclical peaks occurred are classified as years of cyclical expansion, so that $DD_t = 0$ if a reference cycle peak occurred during year $t$. Similarly, years during which cyclical troughs occurred are classified as years of cyclical contraction, so that $DD_t = 1$ if a reference cycle trough occurred during year $t$.

(2) The reference cycle dates provided in the above sources, together with the aforementioned convention for classifying reference cycle peaks and troughs, imply that $DD_t$ is coded to equal 1 for each of the following years in the 1921–1972 period: 1921, 1924, 1930–1933, 1938, 1946, 1949, 1954, 1958, 1961, 1970.

7. $U^*_t$ = the aggregate civilian unemployment rate in Canada in the year of the most recent business cycle peak prior to year $t$ (in percentage points).

Sources: (1) The years in which cyclical peaks occurred are identified using the reference cycle dates obtained from the sources given above for the cyclical contraction dummy variable $DD_t$.

(2) The values of $U^*_t$ are obtained from the same sources as those for $U_t$.

8. TOTUSₜ = total trade union membership in the United States in year t, including Canadian membership in U.S. unions (in thousands).

Sources: (1) The figures for TOTUSₜ for the years 1921-1929 are those contained in series D735 of U.S. Department of Commerce, Bureau of the Census, Historical Statistics of the United States: Colonial Times to 1957, p. 97.


9. INTMEMₜ = Canadian membership in international (U.S.-based) unions in year t (in thousands).

Sources: (1) For the years 1921-1966, the figures for INTMEMₜ are taken from Table IX.C of Labour Canada, Union Growth in Canada 1921-1967, p. 96.

(2) For the years 1967-1972, the figures for INTMEMₜ are taken from various annual issues of Labour Canada, Labour Organizations in Canada, Catalogue No. L2-2/19xx.

10. TUSₜ = trade union membership in the United States, exclusive of Canadian membership in international (U.S.-based) union, in year t (in thousands).

Sources: For each year 1921-1972, TUSₜ is measured by the difference between TOTUSₜ, total trade union membership in the United States in year t, and INTMEMₜ, Canadian membership in international (U.S.-based) unions in year t; i.e.,

\[ TUSₜ = TOTUSₜ - INTMEMₜ \]

11. Sₜ = total number of strikes and lockouts in Canada beginning during year t.

Source: For all years 1921-1972, the data values for Sₜ are taken from Table 1 of Labour Canada, Strikes and Lockouts in Canada, 1972, Catalogue No. L2-1/1972, pp. 28-29.
12. $D_{44}^t$ is a binary dummy variable denoting the years following enactment by the federal government of Order-in-Council P.C. 1003 in 1944; it is specified to equal 0 for all years 1921-1943 and 1 for all years 1944-1972.

13. $U^p_t$ is the aggregate civilian unemployment rate in Canada in the year of the most recent business cycle trough prior to year $t$ (in percentage points).

Sources: (1) The years in which cyclical troughs occurred are identified using the reference cycle dates given in the sources indicated above for the cyclical contraction dummy variable $D_{90}^t$.

(2) The values for $U^p_t$ are the appropriate values of $U_t^p$, the aggregate civilian unemployment rate in year $t$.


A.2 Definitions of Variables Used in Estimating Equations

1. $\Delta T_t = \frac{T_t - T_{t-1}}{T_t} \cdot 100$  
   $\Delta T_t$ is the annual percentage rate of change of total Canadian trade union membership in year $t$.

2. $\Delta P_t = \frac{P_t - P_{t-1}}{P_t} \cdot 100$  
   $\Delta P_t$ is the annual percentage rate of change of the all items Consumer Price Index for Canada in year $t$.

3. $\Delta E_t = \frac{E_t - E_{t-1}}{E_t} \cdot 100$  
   $\Delta E_t$ is the annual percentage rate of change of unionizable employment in Canada in year $t$, where:
(1) for the years 1922-1931, \( \Delta E_t = \frac{NAGEM_t}{NAGEM_{t-1}} \times 100 \), the annual percentage rate of change of total non-agricultural employment in Canada in year \( t \);

(2) for the years 1932-1972, \( \Delta E_t = \frac{NAPWK_t}{NAPWK_{t-1}} \times 100 \), the annual percentage rate of change of non-agricultural paid employment in Canada in year \( t \).

4. \( D_{t-1} = \left( \frac{T_{t-1}}{E_{t-1}} \right) \times 100 \)

= union density in Canada in year \( t-1 \), i.e., total Canadian trade union membership in year \( t-1 \) as a percentage of unionizable employment in year \( t-1 \), where:

(1) for the years 1922-1931, \( D_{t-1} = \left( \frac{T_{t-1}}{NAGEM_{t-1}} \right) \times 100 \), total Canadian trade union \( t-1 \) membership in year \( t-1 \) as a percentage of total non-agricultural employment in year \( t-1 \);

(2) for the years 1932-1972, \( D_{t-1} = \left( \frac{T_{t-1}}{NAPWK_{t-1}} \right) \times 100 \), total Canadian trade union \( t-1 \) membership in year \( t-1 \) as a percentage of non-agricultural paid employment in year \( t-1 \).

5. \( DD_{t} (U_{t} - U^{*}_{t}) \) = the difference for years of cyclical contraction between the current period unemployment rate \( U_{t} \) and the unemployment rate in the year of the most recent business cycle peak prior to year \( t \) \( (U^{*}_{t}) \), where \( DD_{t} \) is a binary dummy variable which equals 1 for years of cyclical contraction and 0 for years of cyclical expansion.

6. \( T_{US,t} = \left( \frac{T_{US,t} - T_{US,t-1}}{T_{US,t}} \right) \times 100 \)

= the annual percentage rate of change in year \( t \) of U.S. trade union membership exclusive of Canadian membership in international (U.S.-based) unions.
7. \[ \Delta S = \left( \frac{S_t - S_{t-1}}{S_{t-1}} \right) \cdot 100 \]

This is the annual percentage rate of change of the total number of strikes and lockouts beginning in Canada during year \( t \).