SOME EMPIRICAL AND METHODOLOGICAL PROBLEMS WITH THE ESTIMATION OF THE IMPACT OF UNEMPLOYMENT INSURANCE LIBERALIZATION ON THE MEASURED RATE OF UNEMPLOYMENT

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GEORGE ALLAN HARRISON

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ABSTRACT

The impact of unemployment insurance liberalization on measured unemployment was investigated, using two different regression models and quarterly data.

An initial model covering the period 1959 to 1976 showed strong evidence of a structural shift at about the first quarter of 1970. When the model was re-estimated for the period from 1970 to 1976, no statistical evidence for an impact on unemployment due to the 1971 liberalization of unemployment insurance benefits was found. When a second model was fitted to the same period, it also failed to show any impact of unemployment insurance liberalization of unemployment rates.

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INTRODUCTION

An effect of unemployment insurance benefits (UIB) which has received considerable attention in recent years is the increase in measured unemployment which might be attributable to them. It is frequently argued that the presence of UIB reduces the incentive to seek new employment for those who are currently unemployed and receiving UIB and increases the incentive to become unemployed for those persons who are employed in occupations which are covered by unemployment insurance.

A number of techniques for estimating this assumed effect have been developed, which typically depend on measuring the effect of variations in the degree of liberality of UIB on either unemployment levels, unemployment rates, or the duration of spells of unemployment.

In Canada attention has centered on the 1971 revisions of the Unemployment Insurance Act, since these revisions sharply increased the levels of UIB over a fairly short period of time, providing an opportunity to observe the effect of this change on the level of unemployment, or the duration of spells of unemployment.

In the U.S. the majority of studies rely on inter-state and inter-personal comparisons of the levels of UIB and attempt to measure the effect of variations in these levels on duration of spells and levels of unemployment. In a few studies changes in state programs which affect the liberality of UIB provide the source of variation.

The purpose of this study is to present some of the problems encountered in trying to single out the effect of UIB on unemployment, and to re-estimate the effect using a different method to that commonly used, in the process some criticisms of a number of Canadian studies are developed.

The Nature of the 1971 Changes in the Unemployment Insurance Act

Based on a House of Commons resolution of 1968, the Unemployment Insurance Commission conducted an extensive revision of the Unemployment Insurance Act. The effect of the revision was to greatly increase the level of benefits, to lengthen the waiting period for claimants, and to increase the number of persons covered by Unemployment Insurance to include virtually all wage and salary earners. A comparison of the pre and post 1971 Unemployment Insurance acts is given in Appendix A.

While benefits were liberalized under the new regime, they were also made taxable; in addition work related income was deducted from Unemployment Insurance benefits if it exceeded 25% of the benefit rate, as opposed to 50% under the pre-1971 regime. Maternity benefits were made available to women with 20 or more weeks of insured employment in the year prior to their confinement.

The combined effect of these revisions was probably to make Unemployment Insurance look more attractive to workers, although some changes made claiming Unemployment Insurance benefits less desireable.

CHAPTER ONE

THE CHOICE BETWEEN LABOUR AND LEISURE

The simplified standard model of the choice between an extra unit of work and an extra unit of leisure for an individual worker can be represented by a set of indifference curves, and a price line, given by the prevailing wage. Assuming 'rational' behaviour on the part of the workers, an increase in non-labour earnings, or a decline in the relative price of labour will lead to a reduction in hours worked. This result can be illustrated graphically as in Figure 1.

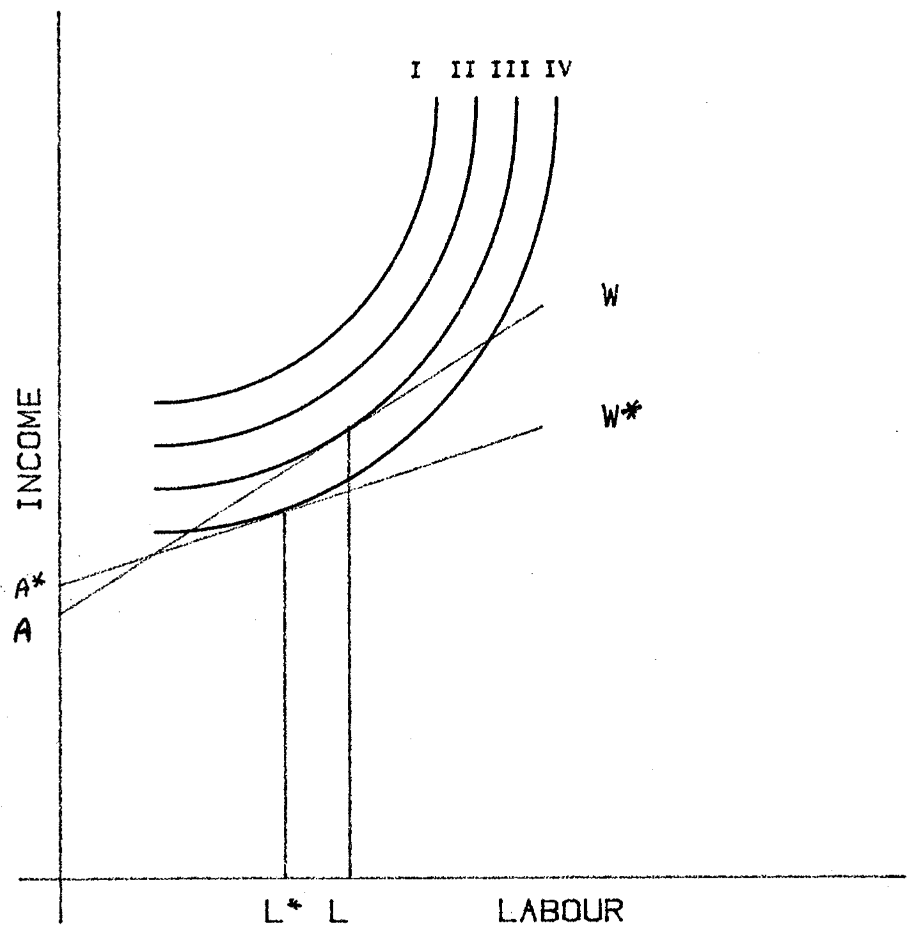
In this figure, the worker's preferences between income and time worked are given by a set of indifference curves. Since the worker is presumed to feel that labour is not intrinsically desirable, the indifference curves slope upward to the right of the figure. The prevailing wage in the absence of unemployment insurance benefits is given by the straight line 'W', the effective wage for a worker eligible for unemployment insurance benefits is given by the line 'W^*', while the non-wage income of the worker is given by the point at which the wage-line intercepts the income axis at 'A'.

As the figures shows, lowering the wageline moves the worker to a lower indifference curve in which he/she chooses less income and less work than would be the case in the absence of unemployment insurance benefits (UIB).

If leisure is a normal good, in other words, a decline in its relative price should lead to a rise in its consumption and <u>vice versa</u>. This is the assumption underlying the neo-classical theory of labour supply.

If one assumes that increased UIB result in the lowering of the effective wage facing the worker, it follows that the result will be for the worker to choose more leisure, or less labour. This is the theoretical basis for most work relating to the relationship between unemployment and the level of UIB.

Figure 1
The Choice Between Income and Leisure



The analysis so far, however, is static in that it does not take into account the worker's consideration of the future consequences of his/her actions and the existence of uncertain outcomes of actions which plague real world decisions. To relate the level of unemployment to the level of UIB, we must analyze the appearance of UIB to a worker in a real world situation.

Consider a worker who is working in a job covered by UIB. The effective wage facing the worker is not the money wage 'W' but the difference between the money wage and product of the earnings replacement ratio and the current wage which is given by 'xW'. The earnings replacement ratio is the ratio of the weekly benefits a worker is eligible to receive to the weekly wage which the worker is earning in his/her present job. In addition, the duration of benefits is normally a positive function of the duration of qualifying employment so that by working an extra week, the worker can add part of a week to his or her future claim. (We denote the number of weeks added to a claim by working one week by 'B' and assume that 'B' is less than one). Taking 'W' to be the weekly wage and 'W^*' to b the effective wage in the presence of UIB, we can express this relation as follows:

1. $$W^* = W - xW + xWB = W(1 - x(1 - B).$$

W^* will be less than W unless x goes to zero or B goes to one.

It should be noted that the duration of a claim does not increase indefinitely with the duration of qualifying employment. Thus, after some period of time, B becomes zero and the effective wage becomes:

2. $$W^{**} = W(1 - x).$$

A worker faced with a lowered effective wage may be induced to quit work. If this happens, the worker will normally face a substantial waiting period. Suppose the worker can claim N weeks of income at xW. By quitting, these benefits are postponed D weeks, where D is the standard waiting period for workers who quit their jobs without cause. Average weekly income during the period of unemployment then becomes:

3. $W = N(xW)/(N + D) = xW(N/\ N + D\ ).$

If D is appreciably large with respect to N then average weekly income is reduced by quitting.

So far we have not considered the fact that most people tend to value future benefits and costs less than current ones. In this case, waiting periods occur immediately upon quitting, while benefits are received some time in the future and additional weeks of benefits are received even farther into the future. If workers discount the future, the approximate formula for the value of a stream of income becomes:

4. $$NPV(W) = \sum_{n^1}^{n^2} xW/(1 + r)^i,$$

where;

$NPV(W)$ = Present value of the future income stream,
xW = weekly UIB,
r = the worker's private rate of discounting the future,
n^1 = time period when benefits begin,
n^2 = time period when benefits end.

If r is postive then $NPV(W) < xWN$, so that apparent benefits will be less than indicated actual benefits paid.

The presence of regular financial obligations can make 'r' quite high. The worker may incur various forms of unpleasantness in the present if rent is not paid, loan payments are not kept up, mortgage payments fall behind, or the food budget is exhausted. Such constraints tend to place a high premium on present income as opposed to future income.

Finally, every worker who quits must face the prospect of eventual job search. This will normally be perceived as an onerous task, both because of the effort involved, and the risk of not finding a suitable job before benefits are terminated. Thus the worker is actually faced with a choice between the following alternatives: to continue working and receive income 'W' and leisure '1 - L' each week, where 'L' is the normal work week, or to quit and receive unlimited leisure, and a net benefit stream:

5. $$\emptyset = \sum_{n^1}^{n^2} xW/(1 + r)^i - \sum_{n_1}^{n_2} C^i/(1 + r)^i - \sum_{n_{3+1}}^{n_3} C^i P(C^i)/(1 + r)^i$$

Where $\emptyset$ = the net present value of future costs and benefits,

xW = the value of UIB each week of covered unemployment,

C^i = the costs of job search in week i,

$P(C^i)$ = the probability of incurring a job search in week i.

The last term is included to distinguish between costs incurred in routine search during the period when a worker is supposedly enjoying leisure, and the possibility of incurring a search cost during a period of genuine unemployment after benefits have run out or the worker has decided to return to work.

A further consideration is that a person who quits without just cause, or is fired may be disqualified entirely for an infringement of the rules governing UI. A worker comtemplating quitting must assign a probability to the receipt of benefits. If the probability of a claim being approved is P, then the expected stream of benefits becomes:

6. $$\emptyset' = \emptyset^* P(\emptyset).$$

0' will necessarily be less than or equal to 0.

While the foregoing does not change the qualitative impact of UIB or of liberalization of UIB; and while improved benefits and extended benefit periods will still make the prospect of quitting more tempting, it does imply that it might be rash to assign <u>a priori</u> a large impact on job quits to UI liberalization. This is especially true when the longer waiting period associated with the liberalized benefit regime is taken into account.

In fact, some people may be induced to work more by the liberalization of UIB. If a person would not normally work at a given wage, the prospect of a benefit period after working for a qualifying period may increase the effective wage enough to cause such a person to take a job. The effective wage becomes:

$$7. \quad W^{**} = \sum_{n_0}^{n_1-1} W/(1 + r)^i + \emptyset'.$$

Where $\emptyset$ is the right hand side of equation 5.

The second term which represents the present value of UIB after quitting will be positive, so W^{**} will be larger than W. This will induce some workers to seek employment.

For an unemployed worker currently receiving UIB, the effect of liberalization would be ambiguous, since the reduced return to work in the present will be at least partly offset by the future benefits of new eligibility which result from obtaining employment.

Thus we have not the relatively simple case of a worker facing a reduced effective wage, but a much more complex choice between two alternative streams of benefits. It is not nearly so clear that a worker would be induced to switch jobs due to a fairly temporary 'improvement' in the income/leisure cost composition of his or her benefit stream as is implied in the simple model.

There are also demand side factors. It might be argued that the presence of UIB makes employers more willing to lay workers off during slack periods, leading to higher cyclical levels of unemployment. This argument might be considered in the light of a common argument that firms in the primary labour market engage in labour hoarding to preserve their workforce.

A counter argument might be formulated claiming that the availability of seasonal or cyclical workers who will accept UIB for extended periods as an alternative to waged income makes it possible for some firms to stay in business, thereby decreasing long-run unemployment, although perhaps aggravating seasonal or cyclical variations in unemployment.

A third factor is the potential effect of UIB on aggregate demand. The unemployed are likely to be low savers, hence an expansion of UI payments will likely have an upward effect on aggregate demand, leading to an increase in employment.

In summary, we can say that a liberalization of UIB will tend to induce higher measured rates of unemployment, but this effect will be mediated by several factors:

- the effective earnings ratio,
- the period of disqualification for quitting,
- the entitlement period for payment of benefits,
- the rate at which workers discount future benefits and future costs (which may not be the same),
- the degree of risk aversion in workers,
- the probability of having a claim rejected,
- the ratio of non-wage income to current liabilities,
- cultural and social attitudes,
- enhanced job search due to the prospect of future benefits,
- effects of UIB on employer hiring practices, and
- increased employment due to expanded aggregate demand.

It is impossible <u>a priori</u> to assign a direction to the impact of liberalization, since at least two factors act to reduce unemployment and one is ambiguous in its effect.

Having combined all these impacts, it remains to be noted that the declaration of oneself as being in the labour force is not the same as re-entering the labour force. Discouraged workers who declare themselves as looking for work due to the inducement of UIB liberalization are not changing their actual status. They are unemployed in both cases; they are simply reporting it differently. In this case, an increase in job search activity might lead to an increase in the measured unemployment rate, surely a perverse result. It follows that any positive impact of UIB on measured unemployment rates will include an element of more accurate reporting of actual status, <u>ie</u> it will overstate the welfare relevant magnitude of the change.

Finally, if should be observed that much of what is euphemistically referred to as leisure is in fact unwage labour. If a person who refuses a waged job does so in order to maintain his/her household, the welfare loss to society is not measured by the lost output from that job, but by the difference between that output and the output of the worker in household maintenance. Again, we have an overstatement of the impact of UIB if we look at any detectable increase in unemployment which might result.

CHAPTER TWO

LITERATURE REVIEW

The mainstream of research on the impact of UIB liberalization has largely ignored the complicating factors discussed in the previous chapter and have typically assumed a fairly simple model in which UI benefits reduce the wage facing the worker. This is true of Canadian studies as well as most American studies. This chapter reviews Canadian work in the field and then looks at some American studies.

Canadian Studies

Grubel et al. (1975) use a four equation simultaneous equation model to estimate the effects of UIB related variables on the measured rate of unemployment. The equations fitted were as follows:

$$1.\quad \ln U = -15.15 + 2.35\frac{UCB}{AWW} - 0.03PCGNP - 0.05PCGNP^{-1}0.12FLFPR + 0.17MLFPR - 0.02INEL$$
$$(2.71)\ (2.76)\quad (4.32)\ (7.07)\quad (3.85)\quad (2.76)\quad (4.26)$$

$$2.\quad \frac{UCB}{AWW} = 0.39 - 0.001U - 0.002\ O/MH + 0.003\ MAXBEN$$
$$(16.16)\ (0.4)\quad (7.92)\quad (8.54)$$

$$3.\quad FLFPR = 25.22 - 0.41U + 0.22AWW^{-1} + 2.87\ F40\text{-}49$$
$$(5.29)\ (4.15)\quad (36.36)\quad (8.54)$$

$$4.\quad MLFPR = 21.93 - 0.40\ U + 0.03\ AWW^{-1} + 1.45\ M25\text{-}44$$
$$(3.20)\quad (4.22)\quad (2.20)\quad (9.42)$$

A single equation model was also fitted:

$$5.\quad \ln U = 8.09 + 2.54\frac{UCB}{AWW} - 0.04PCGNP - 0.05PCGNP^{-1} + 0.08FLPR + 0.9MLFPR - 0.02INEL$$
$$(1.88)\ (3.89)\ (5.01)\quad (7.40)\quad (3.37)\quad (1.96)\quad (4.63)$$

The included variables were:

U = the unemployment rate, the endogenous variable of interest,

$\frac{UCB}{AWW}$ = the ratio of unemployment compensation benefits to the average weekly wage,

PCGNP = the percentage change in GNP from the previous year,

FLFPR = the female labour force participation rate,

MLFPR = the male labour force participation rate,

INEL = the ratio of disqualifications of claims to the total number of claims filed,

O/MH = output per worker hour,

MAXBEN = the maximum level of benefits allowed,

F40-49 = the proportion of women aged 40-49 in the labour force,

M25-44 = the proportion of men aged 25-44 in the labour force.

The logic of the model is that demand variables (PCGNP), supply variables (FLFPR and MLFPR), and UI related variables (UCB/AWW and INEL) all affect the rate of unemployment (UR). Since the rate of unemployment can be thought to affect the level of unemployment benefits, Equation 2 was introduced to deal with this possible relationship. Since participation rates also depend on employment conditions, equations 3 and 4 were introduced to capture this relationship.

To determine the effect of UI liberalization on the unemployment rate, values of UCB/AWW and INEL appropriate to the pre-liberalization period were inserted into the equation for the post-liberalization period. The resulting reduction in the estimated rate of unemployment is taken to be the effect of UIB liberalization on the unemployment rate. Grubel et al. estimate that UIB liberalization contributed 0.8 percentage points to the rate of unemployment in 1972.

Green and Cousineau (1976) develop three different methods of estimating the effect of UI liberalization on the level and rate of unemployment.

The first method is based on the work of Gujurati (1972) and Foster (1973) on Great Britain. A hyperbolic relation between the unemployment rate and the job vacancy rate is posited of the form:

$$1nU = b^0 + b^1 1nv + b^3 +$$

where: U = the unemployment rate,

V = the job vacancy rate, suitably defined,

t = a linear time trend.

Or, following Foster:

$$\ln U = a^0 + a^1 \ln V + a^3 \ln U^{-1}$$

where: U^{-1} = the lagged unemployment rate.

Unemployment rates are then forecast for the post-1971 period based on data from the pre-1971 period and the difference between forecast and actual values is attributed to the effect of UI liberalization.

Green and Cousineau find that their data shows an increase in the unemployment rate of 0.75 percentage points for 1971, 1.79 for 1972 and 1.81 for 1973, based on the Gujurati model. The effects are 0.38 for 1971, 0.87 for 1972, and 0.67 for 1973 if the Foster model is used.

Cousineau and Green then fit the following model to unemployment data over the period 1959(1) to 1973(4).

$$U = f(G,L,LFP)$$

where: U = quarterly unemployment,

G = deviations of GNP from its long term trend, suitably lagged,

L = deviations of the labour force from its long-term trend,

LFP = trend value of the labour force.

A number of specifications of the lag distribution were tried out and the authors note that the model underpredicts for nine out of ten quarters from 1971(3) to 1973(4).

The model was then augmented with two UI related variables; B/W, the ratio of average benefits to the average weekly wage, and X/P, the ratio of the disqualification rate to the proportion of the labour force covered by UI.

Again several specifications of the lag structure were tried out. From the estimated values of the coefficients of B/W and X/P, the authors estimate the impact of UI liberalization by applying these coefficients to the average change in the level of the two variables, for each year. The estimated impacts are 0.03 percentage points in 1971, 0.63 or 56,788 in 1972, and 0.35 or 32,880 in 1973 (based on their equation 1 in their table 3) (Cousineau and Green, 1976).

The effect of liberalization on the labour force is estimated by the following equation:

$$PR = \underset{(29.9)}{55.0} + \underset{(4.3)}{8.6W^{-2}} + \underset{(2.1)}{6.77\ G^{-2}} - \underset{(-9.2)}{0.44\ BR} + \underset{(2.1)}{3.59\ (T-1)/T}$$

$$R^2 = .881 \qquad D.W. = 1.48 \qquad S.E.E. = .282$$

In this equation:

W^{-2} = the logarithm of the deviation of the average weekly wage from its long-term trend, lagged two periods,

G^{-2} = deviation of Real GNP from its long-term trend lagged two periods (in logarithms),

BR = birth rate,

T = a time trend.

The deviations of the predicted from actual PR was taken as a measure of the impact of UI liberalization on labour force participation. The increase in the labour force due to UI liberalization was derived and the regression of the short-term changes in the labour force from the previous equation was used to estimated the effect of this increase on levels of unemployment. This "indirect" effect was added to the "direct" effect given above to produce a total increase of 0.67 percentage points or 60,038 in 1972, and 0.7 or 65,130 in 1973.

Cousineau and Green's third estimate is based on the following identity:

$$R = E - a(1 - r)$$

which is derived from the following set of relations;

$$E = U/LF$$

$$R = U^{*}/LF$$

$$U^{**} = XB = XN$$

$$LF = U + N$$

$$U = U^{*} + U^{**}$$

where:

U = average weekly number of unemployed persons,

U^{*} = average weekly number of persons unemployed for non-UI related reasons,

U^{**} = average weekly number of unemployed due to UI related reasons,

X = number of persons who some time during the year choose to become unemployed due to UI inducements,

B = number of weeks of UI induced unemployment experienced by those becoming unemployed due to UI inducements.

Estimates of X are derived from disqualification data and estimates of B are derived from studies of labour supply effects of income maintenance. The net effect of UI liberalization on unemployment is estimated to be 1.1 percentage points during 1972-73.

Warren (1977) challenges the relevance of the U-V models developed by Gujurati and Foster, claiming they rely on a static equilibrium model while observed vacancy and unemployment rates deviate from equilibrium values. When the appropriate model is substituted for these models, in the case of Great Britain, no statistically significant equilibrium relationship is found.

Maki (1977) estimates the following model:

$$AVDUR = F(UCB/AWW,\ DSQL,\ 12MPCEI,\ PFEM),$$

using seasonal dummies and polynomial lags, for ten provinces and for Canada. The variables included are:

AVDUR = average duration of unemployment,
UCB/AWW = ratio of unemployment benefits to the average weekly wage,
DSQL = disqualifications as a percentage of month end claimants,
12MPCEI = percentage change in employment index over 12 months, and
PFEM = proportion of claimants who were women.

The variables UCB/AWW and DSQL were intended to capture the effect of variations in the liberality of the UI program. 12MPCEI is a proxy for labour market conditions and PFEM was included to correct for the longer duration of unemployment experienced by women.

The fitted equation for Canada, for the period December 1962 to October 1974 was:

$$\underset{}{AVDUR} = \underset{(0.66)}{-0.89} + \underset{(5.96)}{28.25^{*}}\ UCB/AWW - \underset{(2.40)}{0.24^{*}}\ DSQL - \underset{(0.97)}{0.11^{*}}\ 12MPCEI + \underset{(5.31)}{0/140}\ PFM$$

$$R^2 = 0.83 \qquad D.W. = 1.62^{**} \qquad RHO = 0.T4^{**}$$

* sum of 2nd degree polynomial lags
** Hildreth-Lu routine used for estimation

The estimated coefficient for UCB/AWW was multiplied by the change in UCB/AWW between pre and post-liberalization periods. The change was estimated to increase average duration by two weeks. No estimate of the effect on unemployment rates or levels is provided.

Lazar (1978b) first estimates duration and incidence of unemployment spells for Canada over the period February 1966 to December 1975 using a modification of a technique developed by Perry (1972). The estimated values are then used to fit a linear equation of the the form:

$$Y = a^0 + a^1CU + A^2 [(CU - CU^{-3})/CU^{-3}] + A^3 [(CU^{-3} - CU^{-6})/CU^{-6}] + \sum B^i S^i$$

where:

Y = the continuation probability, P, or the turnover rate, T,

CU = the capacity utilization rate, interpolated between quarters,

S^i = seasonal dummies.

Then, following Reid (1977) and Wilton (1975) a transitional dummy variable was introduced into the model, specified as follows:

$$a^0 = C^0 + C^1D(t)\ C^2D(t)^2 + \ldots\ldots\ldots + C^nD(t)^n$$

where:

D(t) = 0 for February 1966 to June 1971,

D(t) = 1/6, 2/6, 3/6, 4/6, 5/6 for July 1971 to November 1971, and

D(t) = 1 for December 1971 to December 1975.

The dummy was intended to capture the effects of the UI liberalization on the transition probability and the turnover rate. The average spell duration can be derived from

$$ADS = 1/P$$

and the estimated unemployment rate (UR) is derived from ADS and T.

Lazar estimated the effect of UI liberalization on ADS, T, UR and the number of unemployed for men and women by age group for 1972 to 1975. These results were aggregated to give the increment in unemployment and the unemployment rate over the period. The increases were 119,000 or 1.3

percentage points in 1972, 94,400 or 1.0 in 1973, 103,200 or 1.1 in 1974, and 150,000 or 1.5 in 1975.

Siedule et al. (1976) extend the Candide 1.1 labour force block to include an equation designed to capture the effect of UIB on the labour force participation rate. They then run simultations for the 1971-79 period with the UI related variable at their pre-1971 levels. The difference between the simulated value with UI variables at actual values and that with UI variables at pre-1971 levels is used to estimate the effect of UI liberalization on the level of unemployment.

For 1971, the liberalization was found to reduce the unemployment rate by .1 percentage points and unemployment by 7500. For 1972, the effect was to cause an increase of 0.74 percentage points or 74,100 persons.

Rea (1977) decomposes the 1971 revisions into wage and income effects and estimates the impact on labour supply, using a labour supply equation estimated from the Unemployment Insurance Commission longtitudinal client sample. The estimated impact on the unemployment rate is 0.61 percentage points.

Table 1 presents a summary of the findings of Canadian studies.

TABLE 1

SUMMARY OF ESTIMATED EFFECTS OF UNEMPLOYMENT INSURANCE ACT LIBERALIZATION

AUTHORS	EFFECT			
	Levels of Unemployment Absolute Increase	Unemployment Rate (% Age Point Increase)	Average Duration (% Age Increase)	Turnover Rate (% Age Increase)
Grubel *et al.*	--	0.8 in 1973	--	--
Green & Cousineau	60,038 in 1972 65,132 in 1973	0.67 in 1972 0.70 in 1973 0.38-0.75 in 1971 0.87-1.79 in 1972 0.67-1.81 in 1973 1.1 in 1972-73	-- --	-- --
Lazar	119,200 in 1972 94,400 in 1973 103,200 in 1974 150,800 in 1975	1.3 in 1972 1.0 in 1973 1.1 in 1974 1.5 in 1975	18.0-40.7 18.3-33.8 17.8-34.7 17.6-34.7	14.0-33.8 14.2-40.6 14.2-34.5 11.1-25.5
Rea	--	0.7 in 1972	--	--
Maki	--	--	20.0	--

American Studies

American studies have also uncovered a positive correlation between the liberality of UI programs and the level of unemployment.

Marston (1975) compared the expected duration of employment for insured vs. uninsured workers. He concluded that the unemployment rate in 1969 in the U.S. would have been between 0.19 and 0.34 percentage points lower if unemploy- ment insurance were unavailable. He suggested that the impact of UI on unemployment would be lower in a slack labour market than in a tight market.

Ehrenberg and Oaxaca (1976) used inter-state and inter-personal variations in the 'replacement ratio', the ratio of an individuals UI benefits to his or her pre-unemployment wage. They estimate effects on duration of employment for four groups; older men, older women, younger men and younger women. They found that an increase in the replacement ratio from 0.4 to 0.5 led to an increase in duration of a spell of unemployment of 1.5 weeks for older men, 0.3 weeks for older women, 0.2 weeks for younger men and 0.5 weeks for younger women.

Holen (1977) used interstate variations in the replacement ratio to estimate the effect of benefit levels on duration of spells of unemployment. She found that a $10.00 increase in weekly benefits would lead to a 0.8 week increase in the duration of a spell of unemployment.

Classen (1977) compared the duration of claims for persons in the same state whose replacement ratio differed because of changes in UI benefits. In both Pennsylvania and Arizona, where some recipients were receiving benefits under the old and new UI regimes, it was found that duration of claims was longer for those receiving the newer, higher benefits. Classen found that in Pennsylvania an increase of $10.00 in benefits would lead to a 1.1 week increase in duration of claims. In Arizona the increase was one week. An increase of this magnitude would lead to a 0.4 percentage point rise in the unemployment rate.

Mortensson (1977) criticizes studies such as the foregoing for ignoring the incentive effects of UI on unemployed workers not eligible for UI benefits. An improvement in the benefit system will encourage such workers to obtain employment since the returns to employment will be increased. Workers obtaining insured jobs would become eligible for future benefits and would accept jobs sooner, reducing their average duration of unemployment. For this reason, studies which focus only on the behavior of insured workers will over-estimate the impact of UI liberalization on duration of spells of unemployment and on the unemployment rate.

Barron and Gilley (1979) investigate this possible effect. They find no significant relationship between search intensity of individuals not receiving UI benefits and the expected value of UI benefits should they obtain a job and subsequently quit.

Horowitz (1977) investigates the effect of more stringent eligibility controls on the rate of unemployment, and finds a significant relationship. He estimates that a five percentage point reduction in the fraction of claimants passing a work test would lead to a reduction of the unemployment rate in the U.S. of about 0.6 percentage points.

Grubel and Maki (1978) use a model very similar to the one used by Grubel et al. (1975) for Canada and find a positive effect of UI liberality on unemployment rates. They estimate that the elasticity of the unemployment rate with respect to the ratio of benefits to wages is about 6.0 for time series analysis and 0.9 for a cross-section analysis relying on interstate variations.

The Impact of Unemployment Benefits on Re-Employment Earnings

Some American studies have attempted to discover whether improved benefits to unemployed workers lead to more effective job search. If workers face lower search costs, they may search longer to find better jobs leading to higher post employment earnings and reduced worker turnover.

Burgess and Kingston (1977) find that increased UI benefits lead to increased earnings in post-employment jobs. Based on a sample of 1719 UI claimants in three U.S. cities, they found that an increase in the maximum weekly benefit of $1.00 was associated with a $25.00 increase in annual earnings. A one week increase in potential compensated duration was associated with an increase in annual earnings of $69.00. They conclude that improved UI benefits enable workers to resist the pressure to accept low-paying or unstable jobs.

Holen (1977) finds a very high return to additional job search. An increase of $10.00 in weekly benefits leads to an increase of $360 per year in subsequent earnings. An increase of one week in potential insured duration of claim leads to an increase of $10.00 in annual earnings.

Classen (1977) in the study referred to above found that weekly benefits had no influence on post-unemployment earnings. Classen's sample included layoffs, which might have reduced the effect of weekly benefits, but even when layoffs were excluded no relation between weekly benefits and post-unemployment earnings was detected.

Ehrenberg and Oaxaca (1977) found that for older men and women, and increase in the ratio of benefits to wages from 0.4 to 0.5 would increase post-unemployment earnings by 7% and 1.5% respectively, but could find no significant relationship for younger men and women.

CHAPTER THREE

SOME METHODOLOGICAL PROBLEMS WITH CURRENT RESEARCH

The Canadian studies discussed in the previous chapter, with the exception of Rae (1977) share a common methodological basis. A simple model of the determination of excess supply in the labour market is developed and a variable is introduced which is intended to capture the effect of variations in UI liberality. This model is then fitted to a times series on the specified variables and a one-tailed t-test is conducted on the coefficients of the UI related variable(s). In some of the more ambitious studies simultaneous equations are used to deal with the jointly endogenous nature of unemployment and unemployment insurance benefits.

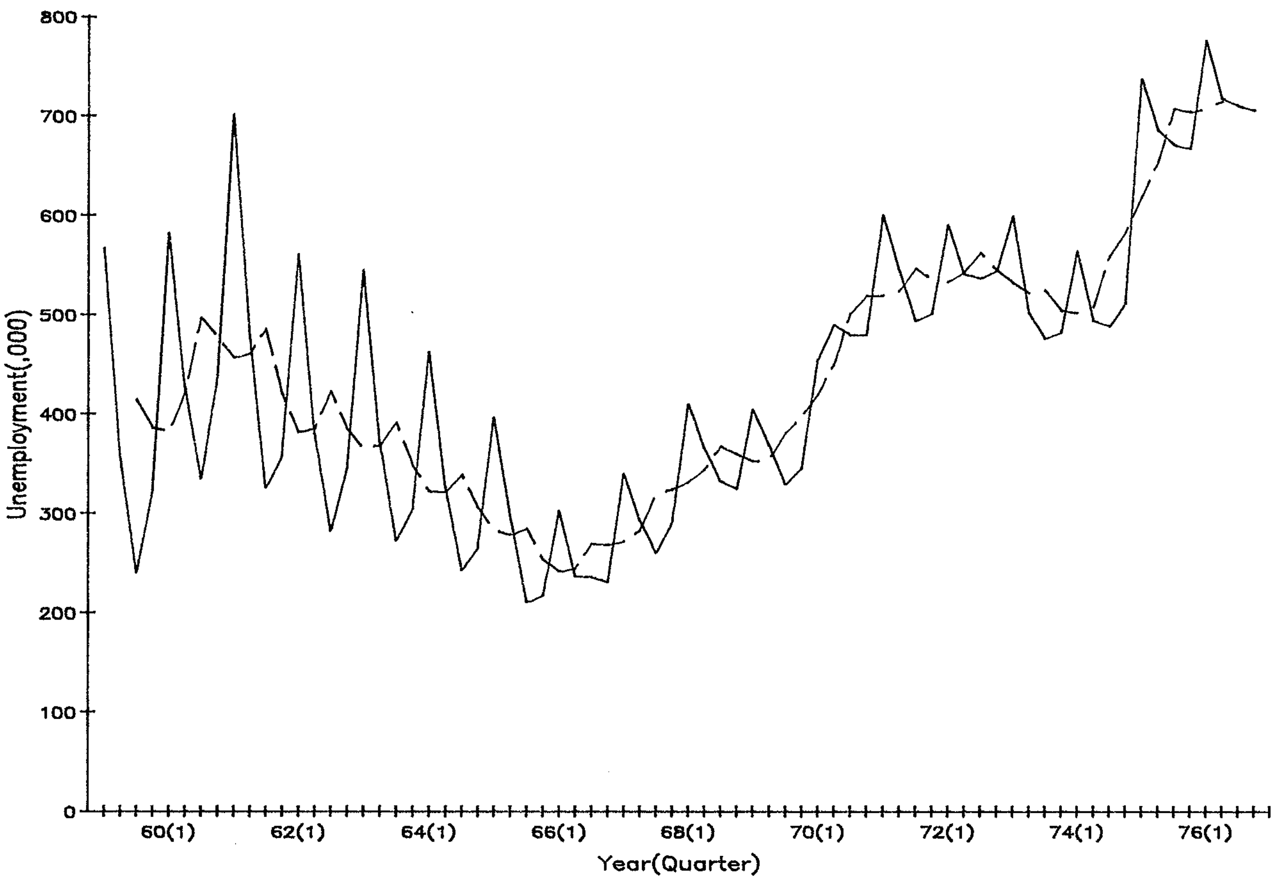
These models embody an implicit assumption about the nature of the causal link between unemployment insurance and unemployment. Concentrating on the simple choice between labour and leisure as developed early in Chapter Two of this paper, a hypothesis is developed that an increase in UI liberality will lead to an increase in measured rates of unemployment (or levels of unemployment).

What is not typically done is to deal explicitly with the possibility of causation acting in the other direction. That is to say, no one seems to take seriously the possibility that UI benefits might have been improved because unemployment had become high. At best the possibility is conceded, and then tucked away in some dusty corner of the econometric edifice being built; perhaps on the assumption that simultaneous equations methods will deal with the issue anyway.

Yet the possibility that such might be the case is strongly suggested by a casual inspection of the data as presented in Figure 7. The graph of actual unemployment in Canada shows a sharp upturn around 1966, which rose to historically high levels around 1970. This is of course the period in which UI benefits were increased, coverage was greatly extended, and waiting periods much reduced. Figure 7 illustrates this phenomenon by including a centred, five-quarter moving average of unemployment levels, to partially eliminate seasonal influences and make the underlying trends more apparent.

Figure 1.1

Actual and Adjusted Unemployment for Canada
1959(1) to 1976(4)



Source:Statistics Canada

A second problem revolves around the use of a variable such as the benefit/wage ratio to measure the effects of UI liberalization. This variable is typically highly correlated with unemployment on a seasonal, and cyclical basis, with the direction of causation running from the level of economic activity through unemployment to the benefit/wage ratio. As unemployment rises, wage increases tend to slow down and more highly paid workers tend to be laid off. This in turn increases benefit levels, which are determined by the applicants wages, the result is an increase in the benefit/wage ratio coincident with an increase in unemployment.

Particularily when single equation models are used to estimate the effect of UIB on unemployment, this effect, which is quite marked, gets confounded with any effect which might result from an actual shift in the level of UIB, inflating the estimate of the effect of UI liberality on unemployment (Hammermesh, 1978).

Finally the use of long time periods to estimate the coefficients of the models specified introduces the risk of combining structurally different periods into one model, thereby biasing the estimates of the coefficients.

This problem is potentially very acute in cases such as this where a dummy variable, or a "quasi-dummy" variable such as the benefit/wage ratio is used as a proxy for the liberality of UIB. This term is used to describe the behaviour of variables such as the benefit/wage ratio which, when short-term seasonal and cyclical fluctuations are removed, consists of a variable with essentially two levels and a short period of transition between levels. As figure 1.1 shows, when, seasonal fluctuations are removed by a five period moving average process, the benefit wage ratio begins to look very much like a two level variable. The slight secular decline from 1959 to 1968 is due primarily to average wages rising more rapidly than benefit levels, a phenomenon. The behaviour of the variable is dominated by the change in levels from 1971(2) to 1971(4).

The problem is acute because the introduction of a dummy variable into a model combining two or more structurally different periods is almost certain

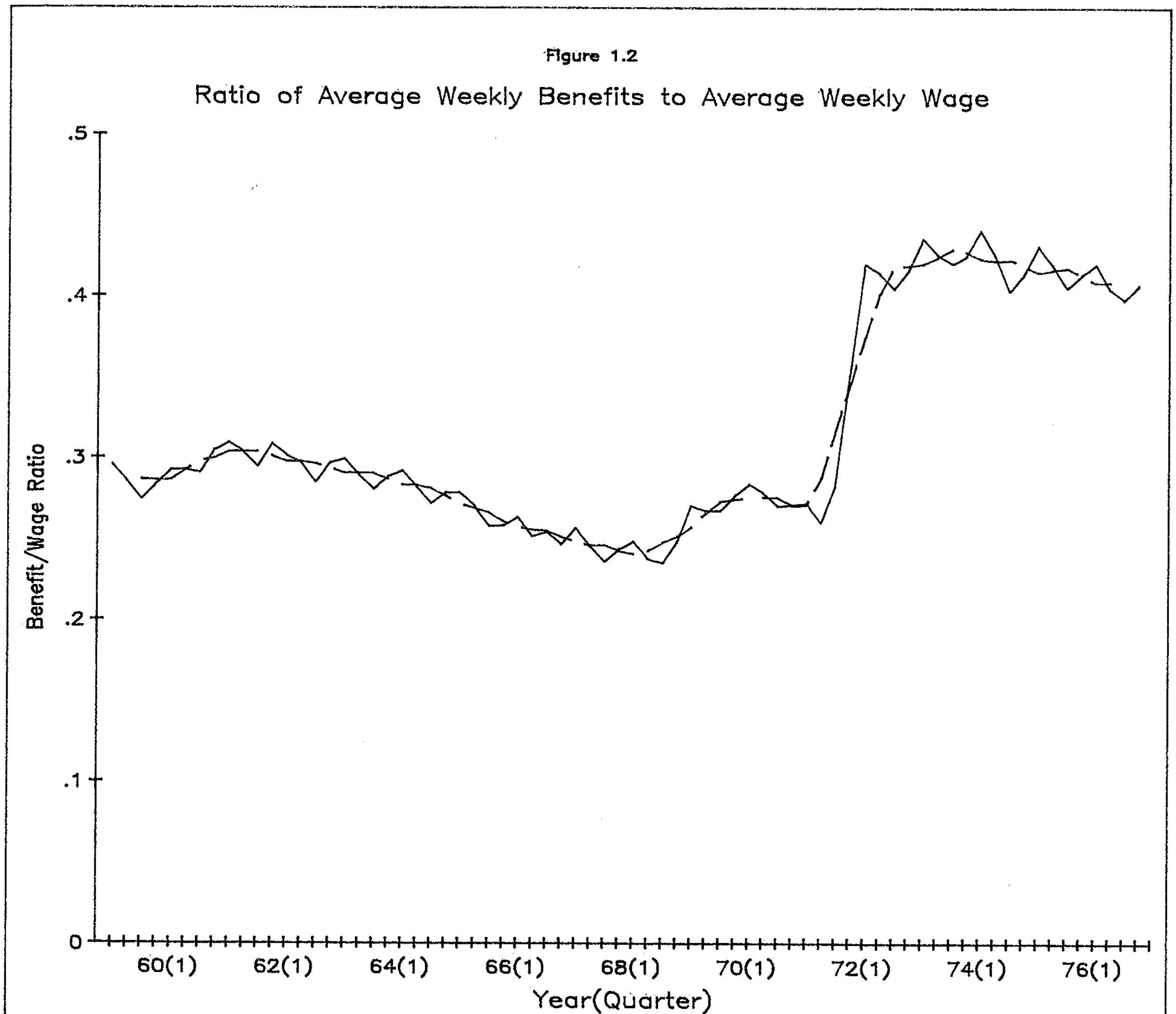


Figure 1.2
Ratio of Average Weekly Benefits to Average Weekly Wage
Benefit/Wage Ratio
.5
.4
.3
.2
.1
0
60(1)
62(1)
64(1)
66(1)
68(1)
70(1)
72(1)
74(1)
76(1)
Year(Quarter)

to improve the quality of the estimated model's fit to the data, and therefore appear with a statistically significant coefficient. This is so even if the dummy changes value several periods earlier or later than the actual time of shift.

To illustrate this effect, an artifical model was developed involving a single independent variable and two structurally different periods. The model was specified as follows:

1. VAR(I) = 1 + DUMMY + 0.3T + e(1) for I = 1 to 3 and,
2. VAR(I) = 1 + 2*DUMMY + 0.3T + e(2) for I = 4 to 6.

In this model e(1) was specified as a 0.5*N(0,1) random variable and e(2) was specified as a N(0,1) random variable, and DUMMY was a variable whose value was zero for periods 1 through 52 and whose value was 1 for periods 53 through 72. In all, six regressions were run for each of 40 models which were specified as follows:

VAR(I) = A + B*D(J) + C*T.

In each case D(J) was a dummy variable such that:

D(J) = 0 T J AND D(J) = 1 T = J for J=32 to 72.

For example: D(39) would take on a value of zero for periods 1 through 38 and a value of one for periods 39 through 72. D(40) in turn would take on a value of zero for periods 1 through 39 and a value of one for periods 40 through 72.

By regressing the generated observations on the variable T, taken with each dummy variable D(32) through D(72) in turn, a series of models can be fitted to the data. In each set of regressions, however, only the fitted models which includes D(53) is correctly specified.

When the coefficient of the dummy variable included in the mis-specified models is examined, however, it turns out in many cases to be statistically significant. In fact, as Table 2 shows, a dummy variable which changes from zero to one as much as seven periods before the actual shift, or up to twelve periods after the actual shift, can appear with a statistically significant coefficient.

The implications for the methods used in many studies of UI liberalization is that if a structural change actually occurred anywhere during the latter part of the 1960's or early 1970's, it might show up as a significant effect of UIB liberalization on measured unemployment rates.

A possible solution then would be to first try to determine when any structural changes may be taken place and to confine the use of econometric analysis to those periods which are likely to be internally homogenous, with respect to the other explanatory variables in the model, or to excluded variables.

Fortunately, Quandt (1958) has developed a method which allows this to be done, at least to some degree. This method will be discussed in more detail in Chapter Four, when the methods and results of the empirical part of this study are discussed.

TABLE 2

VALUES OF THE T-STATISTIC FOR THE DUMMY VARIABLE WITHIN TWELVE PERIODS OF THE ACTUAL SHIFT OF INTERCEPT

PERIOD OF SHIFT OF DUMMY	EQUATION					
	1	2	3	4	5	6
41	0.58	-0.21	-0.01	-0.27	0.06	-0.57
42	1.36	0.39	0.22	-0.57	0.42	0.58
43	1.21	-0.13	0.91	0.27	0.46	-0.07
44	1.30	-0.19	0.91	0.27	-0.11	0.46
45	1.47	-0.50	1.17	0.39	0.19	0.74
46	2.02**	-0.19	1.90***	0.80	0.51	0.43
47	2.81	-0.10	2.23**	0.87	1.55	0.19
48	3.01*	0.48	2.98*	1.60	2.01**	0.21
49	3.58*	0.80	2.98*	1.60	2.01**	0.21
50	3.17*	0.92	3.15*	2.25**	2.28**	0.82
51	3.74*	1.77***	3.96*	2.42**	2.45**	1.79***
52	4.25*	2.24**	4.67*	3.42*	2.95*	3.04*
53	3.96*	3.01*	4.54*	3.47*	3.57*	3.85*
54	4.33*	3.35*	5.80*	3.31*	4.28*	4.66*
55	3.23*	2.26**	5.37*	2.74*	3.33*	4.01*
56	2.86*	2.07**	5.33*	1.53	3.33*	3.53*
57	2.42**	2.33**	5.15*	1.43	2.89*	2.67*
58	2.12**	2.37**	4.07*	1.29	2.57*	2.46*
59	1.76***	1.70***	3.74*	2.20**	1.53	1.90**
60	2.19**	1.43	3.13*	2.05**	1.92***	1.88***
61	1.93***	1.91***	2.55**	1.88***	1.48	2.35**
62	2.72*	1.87**	2.78*	2.16**	1.11	2.32**
63	2.70*	1.72	2.50**	1.27	0.36	2.40**
64	2.32**	2.01**	2.32**	1.29	0.13	1.81***
65	2.49**	2.07**	2.39**	1.26	-0.36	1.22

NOTE:

*significant at the 0.01 level
**significant at the 0.05 level
***significant at the 0.10 level for a one-tailed test against the hypothesis that the coefficient of the dummy is zero

CHAPTER FOUR

METHODS AND RESULTS

The approach adopted to investigate the relationship between UIB and unemployment was to follow as closely as possible the method of one study, namely Cousineau and Green (1978), up to the point of model specification. After choosing a general descriptive model a number of variations in assessing the actual impact of UIB were included which lead to quite different conclusions from those of Cousineau and Green.

Following Cousineau and Green, it is hypothesized that unemployment is a function of long-term labour force growth, the level of economic activity and short run variation in the labour force.

The model was fitted to data covering the period 1959 to 1976. This period is longer than that used for any of the Canadian studies used here. The decision to extend the data was made to provide an adequate number of data points beyond the period of the 1971(3) revisions to ensure statistical reliability for short period estimates. On the other hand, extending the time series much beyond 1976 would increase the possibility of inadvertently including more structurally distinct subsets of data in the regressions. The choice of 1976 seemed to be a plausible time to cut off the series in light of these conflicting constraints.

Long-term labour force growth is measured by the trend growth of the labour force is derived from the equation:

1. $LF^{t} = A^{o}e^{at}$

or

2. $1nLF^{t} = a + at.$

Equation 2 was fitted to data derived from CANSIM to obtain the following fitted equation:

3. $$1n(LF) = \underset{(1449.6)}{8.689} + \underset{(53.1)}{0.0076^{t}}$$

$R^2 = 0.98$ DW = 0.30

MSE = 0.00063   DFE = 70

From equation 3, predicted values for the labour force, LFP, were derived, as were deviations of the labour force from its predicted value, L. Actual and predicted values of the labour force and the calculated residuals are given in Figure 2.

The level of economic activity was measured by the deviation of quarterly GNE from its trend line. The fitted log-linear equation was:

$$4.\quad \ln(\mathrm{GNE}) = \underset{(714.7)}{9.412} + \underset{(41.6)}{0.0313t}$$

$R^2 = 0.96$ DW = 1.85
MSE = 0.0031 DFE = 70

Again, the deviation of GNE from its trend line, G, was calculated. The results are shown in Figure 3.

A linear model relating quarterly unemployment, $UQ^{\dagger}$, to $G^{\dagger}$, $LFP^{\dagger}$ and $L^{\dagger}$ including lagged values of $G^{\dagger}$ and seasonal dummies was then fitted.

The choice of the appropriate lag structure for G was made using a stepwise regression procedure--(SAS's PROC STEPWISE with the MAXR option). This method chooses the best possible regression for models of all sizes ranging from a user specified number of included variables specified up to the maximum number of variables entered in the model for consideration, leaving the actual choice of model to the researcher.

After applying this method to the data, the following equation was chosen as the one most appropriate for the study:

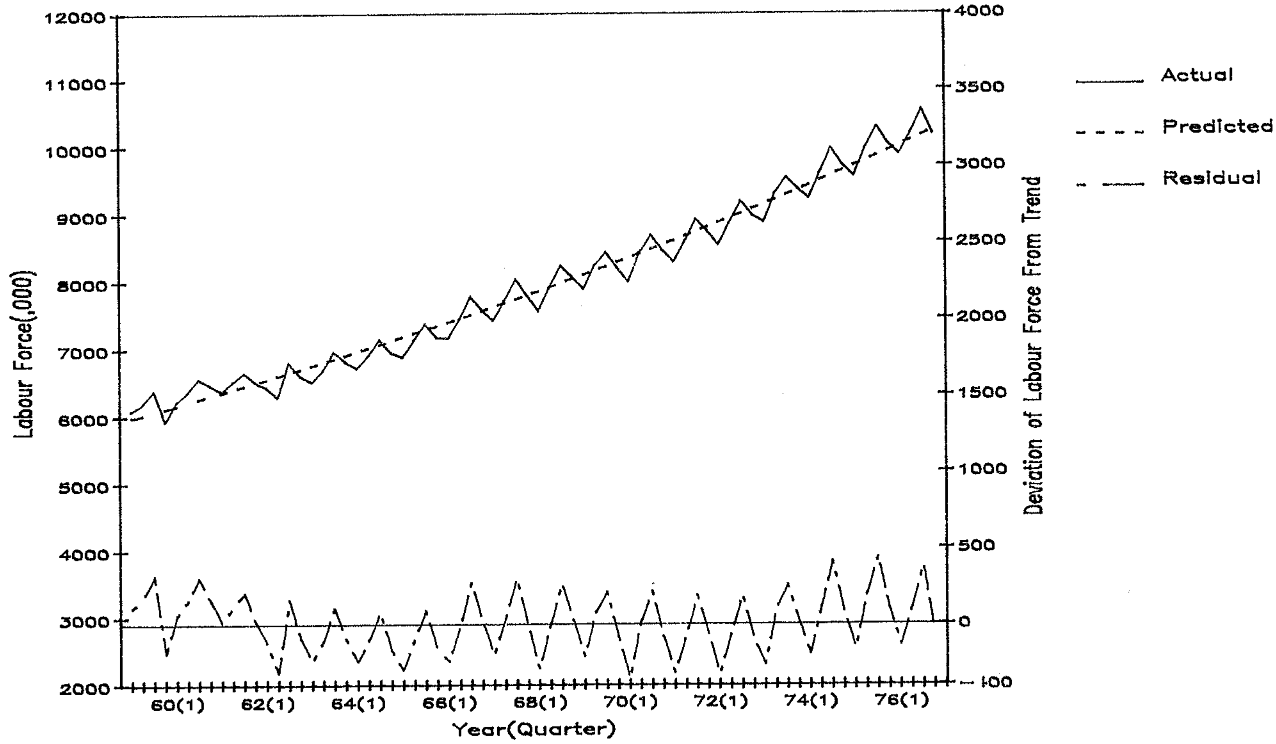
$$5.\quad UQ = \underset{(0.14)}{-28.5} - \underset{(2.73)}{0.066G^{-1}} - \underset{(2.73)}{0.041G^{-2}} + \underset{(9.33)}{0.066LFP} + \underset{(5.07)}{0.41L} + \underset{(3.63)}{2.56.7S1}$$

$$\underset{(2.73)}{-127.4S2} \; \underset{(7.40)}{-327.5S3}$$

$R^2 = .80$ DW = 0.53
MSE = 4967.7 DFE = 62

Figure 2

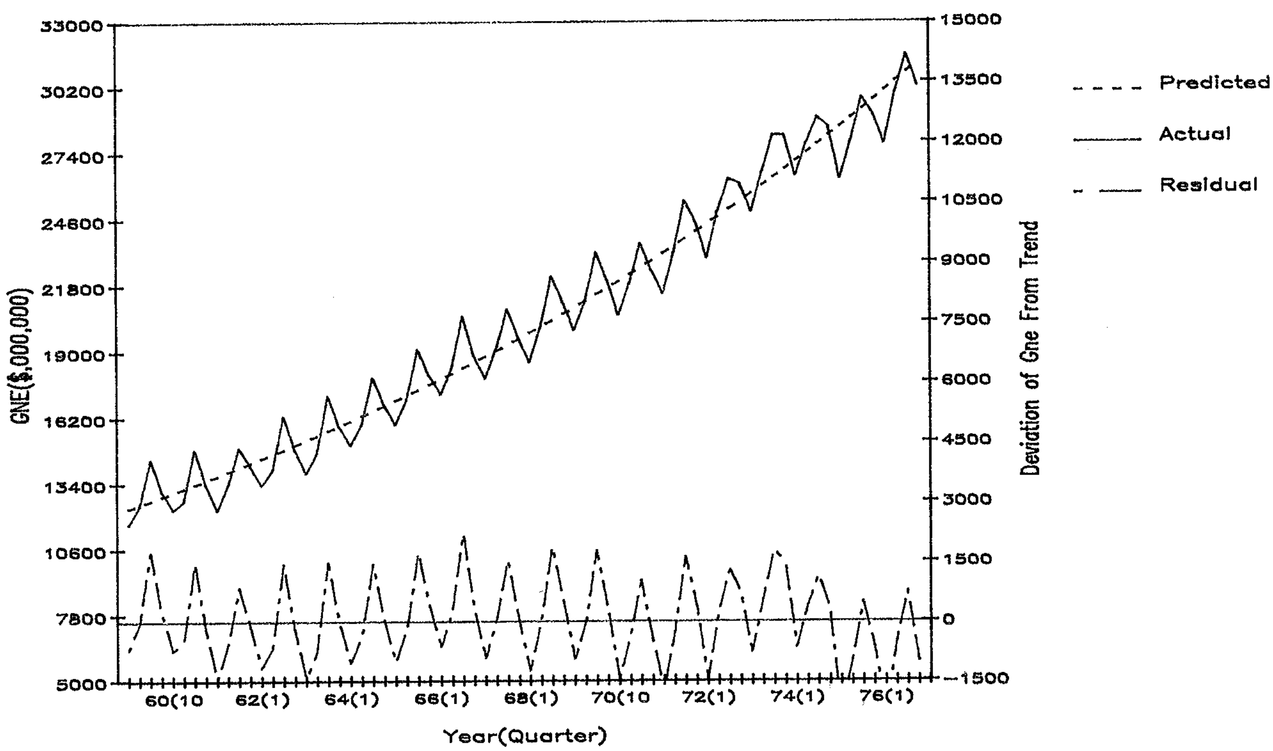
Actual and Predicted Labour Force for Canada
1959(1) to 1976(4)



Source:Statistics Canada

Figure 3

Actual and Predicted GNE for Canada 1959—1976
1971 Dollars



Source:Cansim

Figure 4 shows the actual and predicted values for UQ from this equation and the residuals.

The next step was to compare the residuals from the fitted equation with those of the Cousineau and Green model. The major difference is that the model chosen here underpredicts unemployment, from 1970(1) through 1974(2) and not from 1971(2) as in the Cousineau and Green model (Cousineau & Green, 1978, p. 76).

At this point the analysis presented here diverges from that of Cousineau and Green. Instead of selecting appropriate variables to measure the impact of UI liberality on unemployment, an investigation was carried out to determine if there was a possibility of an underlying structural change not related to UI which might be masked by the specified model.

ESTIMATION OF THE PROBABLE TIME OF A SHIFT OF REGRESSION PARAMETERS

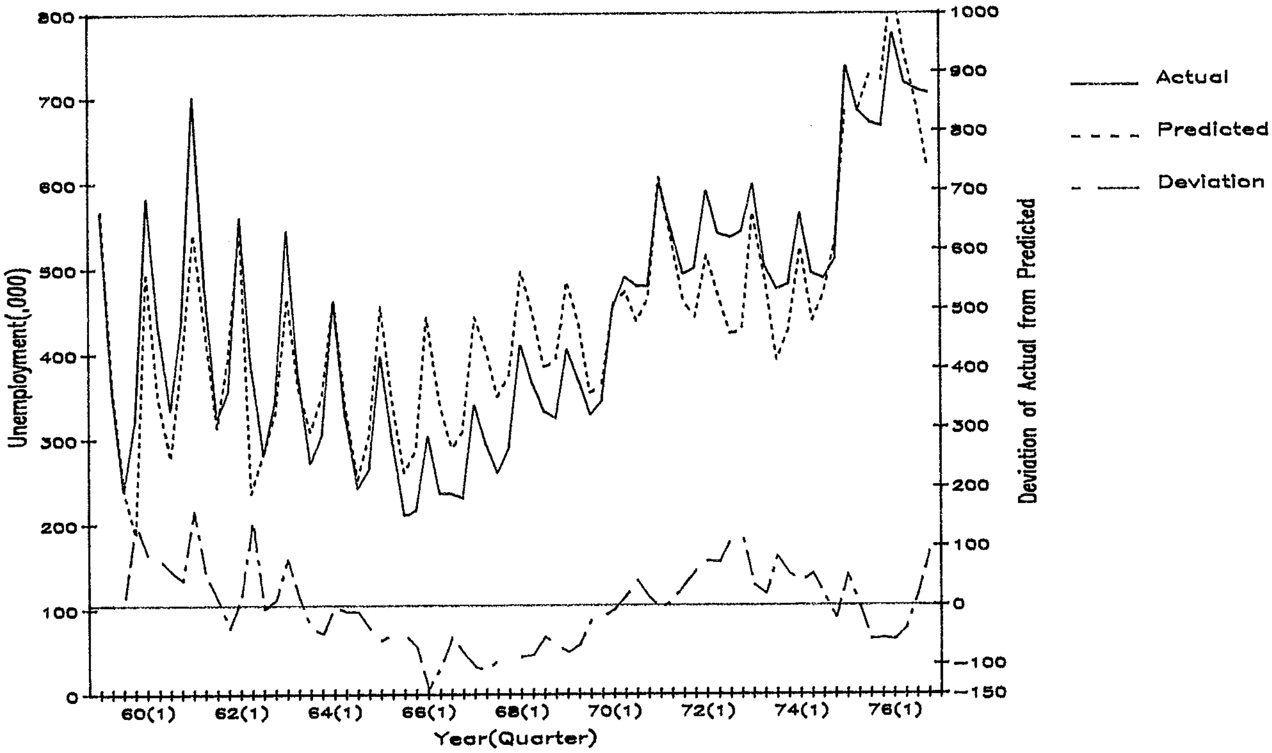
If one were to take the point of time of the change from over-prediction to under-prediction as an estimate of the time of a structural change in the labour market, one might conclude that there was a factor other than UIB liberalization involved since the change occurs at 1970(1), well before the introduction of UI Act changes. Such an inspection, however, leaves much to be desired as a means of estimating a structural shift. Two other methods were therefore used to develop a more sophisticated estimate of the point of shift.

Maximum Likelihood Estimates of the Time of Shift of Parameters

Two possible cases were considered. The entire relation might be shifted upward at some point in time--a shift of intercept. Alternately, one or more slope parameters may change at some point in time--two separate regression regimes may be involved. Each case was to be considered in turn and maximum likelihood estimates (MLE) of the time of change in parameter values were derived.

Figure 4

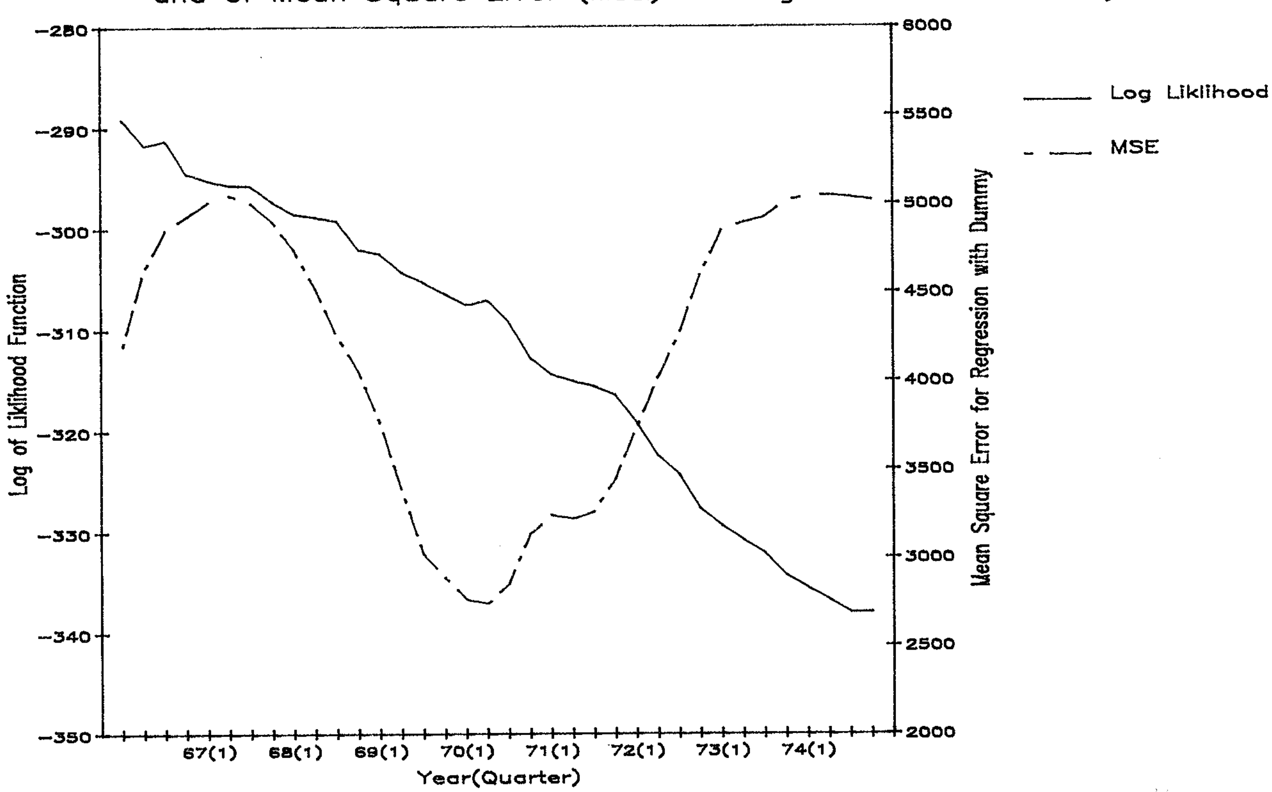
Actual and Predicted Unemployment for Canada 1959(1) to 1976(4)



Source:Statistics Canada

Figure 5

Plot of Log of Liklihood Function for Split Regressions and of Mean Square Error (MSE) for Regressions with Dummy



MLE of the Upward Shift of the Regression

This situation would apply if some factor uncaptured by the model being used underwent a sharp change which resulted in higher unemployment than previously for a given set of values of included variables.

Quandt (1963) gives a method of deriving an MLE of the point in time at which a change occurs in the slope parameter of a single variable regression model. This method generalizes in a straight forward way to both a multiple regression case and to a shift of intercept.

Here the model is extended to the case of a shift of intercept with unchanged coefficients for the explanatory variables. If the variance of the residuals is assumed to be equal in both periods (before and after the shift), the estimating procedure reduces to a very simple process.

A dummy variable is introduced into the model which was formerly fitted to the entire period. The model now is specified as:

6. $UQ = f \quad g^{-1}, G^{-2}, L, LFP, D(T)$, seasonal dummies

where:

7. $D(T) = (0{:}t \quad T, 1{:}T \quad =T)$

for some value of T between 1959(1) and 1976(4). All other variables remain as specified earlier. The MLE of the time at which a shift of intercept occurs is that value of T which minimizes the standard error of estimate for the regression, when D(T) is included in the model. To illustrate this approach, the MSE for the models discussed in Chapter 3 are given in Table 3. In each case, the empirical results lend heuristic support to the method proposed. In five of the six cases, the MSE obligingly falls to a global minimum when the dummy corresponding to the actual time of shift of intercept is used. In the exception (Equation 4), the estimate is cut by only one period.

TABLE 3

MEAN SQUARE ERROR OF FITTED EQUATION FOR DUMMY VARIABLES WITHIN TWELVE PERIODS OF ACTUAL SHIFT OF INTERCEPT

PERIOD OF SHIFT OF DUMMY	EQUATION					
	1	2	3	4	5	6
41	.315	.375	.354	1.312	1.285	1.313
42	.317	.375	.350	1.314	1.288	1.313
43	.316	.375	.342	1.315	1.291	1.310
44	.314	.374	.348	1.312	1.291	1.303
45	.306	.375	.337	1.303	1.287	1.310
46	.290	.375	.331	1.301	1.248	1.313
47	.286	.374	.314	1.268	1.220	1.312
48	.273	.372	.301	1.227	1.213	1.294
49	.283	.371	.310	1.225	1.201	1.301
50	.269	.359	.289	1.213	1.188	1.255
51	.256	.350	.269	1.124	1.147	1.158
52	.263	.332	.273	1.119*	1.090	1.081
53	.254*	.323*	.238*	1.135	1.020*	0.999*
54	.281	.342	.250	1.186	1.113	1.065
55	.289	.354	.251	1.272	1.112	1.112
56	.298	.348	.256	1.277	1.152	1.190
57	.304	.347	.286	1.284	1.179	1.207
58	.310	.360	.295	1.229	1.249	1.247
59	.303	.365	.310	1.239	1.226	1.249
60	.307	.357	.324	1.251	1.251	1.216
61	.293	.357	.319	1.232	1.269	1.219
62	.293	.367	.325	1.285	1.289	1.212
63	.300	.355	.329	1.284	1.291	1.254
64	.297	.354	.327	1.286	1.289	1.285
65	.296	.367	.340	1.246	1.291	1.238

NOTE:

*denotes global minimum mean square error
**denotes period of shift of intercept in underlying model

Although data in the model covers the period 1959(1) to 1976(4), the existence of structural changes much before the time of UI liberalization are of minor concern to this analysis. Such changes are unlikely to have seriously affected the estimate of the impact of UI liberalization on unemployment. Therefore to determine the MLE of a shift of intercept during the period of interest here, values of T were allowed to range from 1966(1) to 1976(4). From Figure 5, it can be seen that there is a minimum in the SEE at T =1970(2), which gives the MLE of the time of shift of intercept.

MLE of the Time of Change of Regression Regime

This case corresponds to a change in the behaviour of workers or employers which results in a changed response to economic and labour market conditions. Following Quandt (1963), the observations for the entire period were partitioned into two periods 1959(1) T-1 and T to 1976(4), and separate regressions were fitted to each subset. The choice of T was varied from 1966(1) to 1974(4), the last period for which a regression could be fitted to the period T to 1976(4). The likelihood function was calculated for each partition. The results are shown in Figure 5. A global maximum occurs at 1966(1) and a local maximum occurs at 1970(2). Again, 1970(2) appears as a MLE of the time of change in regression parameters within the period of interest. The maximum at 1966(1) may be an artifact resulting from the merging of time series spanning the periods 1959-1965 and 1966-1976.

Three estimates of the most likely time of a structural change have now been made. The first, derived from an inspection of the residuals, gives 1970(1) as a time of shift. The second, based on a MLE of the time of shift of intercept gives 1970(2). The third, based on MLE of the time of change of regression regime, also gives 1970(2). Of the three, the third least inspires confidence due to its local nature and the odd behaviour of the log of the likelihood function.

What has not been established is that such a shift has actually taken place. To establish this shift, it is not sufficient to choose the estimated time for the structural shift and conduct a Chow test of equality of regression on the two periods, before and after. Such a test is biased toward rejecting the hypothesis of no difference between the regression regimes in the two periods involved. A better approach is to drop a few observations on each side of the estimated time of shift, fit regressions to the two resulting subsets of observations and test the equality of regressions (Quandt, 1963). The appropriate test statistic is:

$$F = \frac{\left[(SSE)restricted - SSE(unrestricted)\right]/k+1}{SSE(unrestricted)/\ n-2(k+1)}$$

where:

SSE/restricted = the sum of squares for error assuming one set of parameters for both periods,

SSE/unrestricted = the sum of squares for error allowing the parameters to differ in each subset of observations,

k = the number of explanatory variables in the model,

n = the number of included observations.

This statistic has an F-distribution with k and n-2(k+1) degrees of freedom if there is no difference between regressions.

In this case, eight observations for the years 1969 and 1970 were discarded. When the test was conducted, the resulting F-statistic was 8.51 which was larger than the critical value of F(0.05) = 2.18. This leads to rejection of the hypothesis of no change between periods.

On the basis of this test, it seems safe to conclude that a break in the regression occurred in 1970(1) or possibly 1970(2). It was decided to use the period 1970(1) to 1976(4) because this preserves degrees of freedom, while being unlikely to contaminate the latter period significantly. Such a choice is supported by the fact that 1969 is generally held to mark the beginning of a major recession in Canada. It is possible to think of 1970 as marking the beginning of a recessionary structure.

ESTIMATION OF THE IMPACT OF UIB OVER THE SHORT PERIOD

Having identified the period 1970(1) to 1976(4) as structurally different from the period 1959(1) through 1969(4), the next step was to refit the model for the latter period, using a proxy for the Benefit/Wage Ratio to estimate the impact of the 1971(3) unemployment insurance changes on the level of reported unemployment. To ensure that the effects of previous structural change were excluded, while including reasonable number of observations from the period prior to the UI liberalization, the model was fitted for the period 1970(1) to 1976(4).

When the model developed for the long period was fitted to the short period, a problem appeared in that one of the independent variables--L, the deviation of the labour force from its trend value--ceased to be significant, in fact L appears in this equation with the wrong sign. To maintain consistency with the Cousineau and Green methodology, however, L was left in the equation. The fitted equation is given as equation 8. Note that some auto-correlation exists still.

$$8.\quad UQ = \underset{(3.76)}{-592.8} - \underset{(2.29)}{0.0361G^1} - \underset{(3.56)}{0.0571G^2} + \underset{(7.49)}{0.125LFP} - \underset{(1.32)}{0.126L} + \underset{(2.36)}{109.5S1} - \underset{(0.66)}{32.6S2}$$

$$\underset{(1.67)}{-81.8S3}$$

R^2 = .94 MSE = 720.47

D.W. = 1.58 DFE = 20

Rather than correct for auto-correlation, however, a new variable was introduced which was intended to capture the effect of UIB liberalization on measured unemployment. This was the variable BWRDUM, which was specified as follows:

BWRDUM = (0.0:1970(1) to 1972(2))
(0.6:1971(3))
(1.0:1971(4) to 1976(4)).

BWRDUM was so specified to capture the pattern of change in the BWR, following Lazar (1978).

The extended model was fitted to the same data and the resulting regression equation is:

9. $UQ = 358.4 - 0.0457G^1 - 0.0588G^2 + 0.098LFP - 0.101L + 109.6S1 - 57.8\ S2$
$\qquad (1.79) \quad (2.73) \quad (3.75) \quad (4.46) \quad (1.06) \quad (2.57) \quad (1.14)$

$\qquad -104.3S3 + 33.19BWRDUM$
$\qquad (2.57) \quad (1.45)$

$R^2 = .95$ DW = 1.76

MSE = 683.3 DFE = 19

In this equation the introduction of BWRDUM results in a better fit, with slightly improved R^2 and overall reduced auto-correlation. However, BWRDUM is not itself significant at the 5% level and seems to be somewhat collinear with L and LFP. Although the multiple correlation among the explanatory variables is not extremely high--R^2BWRDUM.G^1, $G.^2$, LFP, L, S1, S2, S3 is only 0.72--the exclusion of the last four observations results in a fairly large change in the coefficients of LFP and L as well as BWRDUM. The equation for the shorter period is:

10. $UQ = -222.2 - 0.050G^1 - 0.0628G^2 + 0.833LFP - 0.0424L + 131.2S1$ VARIABLES.
$\qquad (0.56) \quad (2.78) \quad (3.61) \quad (1.88) \quad (0.16) \quad (2.21)$

$\qquad -64.0S2 - 133.2S3 + 45.76BWRDUM$
$\qquad (1.13) \quad (2.21) \quad (1.50)$

$R^2 = 0.92$ DW = 1.81

MSE = 732.4 DFE = 15

Both L and LFP show a marked change from the previous equation, while the coefficient of BWRDUM shows a definite though smaller change. These results are sufficient to suggest that the model is poorly specified for the short period.

Sas's PROC REG provides an option for testing collinearity among the explanatory variables following Belsley, et al. (1980). Inspection of the condition indices shows that one has a large value; the associated variance proportions indicate that the variables, BWRDUM, LFPRED, LFRESID, the intercept, and one of the seasonal dummies, S3, are subject to collinearity problems. As a result, both the actual values of the co-efficients of these variables, and their associated statistical errors are subject to a high probability of error.

AN ALTERNATIVE MODEL

An alternative model specification was developed by converting the annual model presented by Barber & MacCallum (1980) into a quarterly model. The respecified model is based on the assumption that unemployment is a function of: the level of activity in the American economy, the relative unit cost of labour in Canada compared to American unit labour costs, the terms of trade, and the degree of government stimulus of the economy as measured by the full employment deficit/surplus for all levels of government, and by the generosity of UI benefits. These assumptions can be expressed as:

UR = F(USUR,COMP,TOT,FISC,BWRDUM.SEASONAL DUMMIES)

where:

UR = the quarterly unemployment rate in Canada,

USUR = the quarterly unemployment rate in the U.S.,

COMP = comparative labour costs, measured by the ratio of the U.S. Bureau of Labour Statistics' Unit Labour Cost estimate to the Bank of Canada Index of Labour Cost per Unit of Output,

TOT = the ratio of the implicit price deflator for imports as published by Statistics Canada to that for exports,

FISC = the full employment surplus or deficit, estimated by interpolating the annual ratio of the surplus to GNE published by the Treasury Board of Canada between years to produce quarterly estimates and applying these estimates to the quarterly GNE figures provided by Statistics Canada,

BWRDUM = a transitional dummy defined as in the previous section to capture the effect of UIB liberalization.

The model was estimated using data unadjusted for seasonal variation, so three seasonal dummies were included to capture these variations. In addition various lag structures were tried to FISC, COMP, and TOT to capture any lagged influences these variables might have on the unemployment rate.

The model was first fitted to the 1970(1) to 1976(4) data without including BWRDUM. Again the SAS procedure STEPWISE with the MAXR option was used to choose the fitted model. The final fitted equation chosen was:

$$11.\ UR = \underset{(0.99)}{1.95} + \underset{(5.46)}{).41USUR} - \underset{(2.94)}{0.12INFLATN} + \underset{(4.25)}{6.98COMP} - \underset{(2.42)}{4.2TOT^{-3}} + \underset{(9.32)}{1.15S1}$$

$$+ \underset{(2.63)}{0.31S2} - \underset{(1.34)}{0.18S3}$$

$R^2 = 0.94$ D.W. = 2.09

MSE = 0.051 DFE = 20

The equation shows a high degree of fit with the data, low auto-correlation and generally significant coefficients, except for the seasonal variables. However, it has not been expanded to include a UI related variable. This deficiency was corrected by refitting the model including the variable BWRDUM. The result is given in equation 12. Note that BWRDUM is not only non-significant in this equation, it appears with the wrong sign. The fitted equation has a high fit with the data, and low auto-correlation. Moreover, changing the sample period to remove the four quarters for 1976(1) to 1976(4) does not have drastic effect on the coefficients of BWRDUM. When the model was fitted for the period 1970(1) - 1975(4), the coefficient of BWRDUM increased slightly to -0.006 but remained statistically insignificant. The actual results for the longer period are:

$$12.\ UR = \underset{(2.03)}{1.95} + \underset{(4.97)}{0.41USUR} - \underset{(2.90)}{0.11INFLATN} + \underset{(4.16)}{7.11COMP} - \underset{(2.40)}{4.31TOT^{-3}} + \underset{(9.12)}{1.15S1}$$

$$+ 0.31S2 - 0.16S3 - 0.058BWRDUM$$
$$(2.50) \quad (1.32) \quad (0.40)$$

$R^2 = 0.94$ D.W. = 2.06

MSE = 0.054 DFE = 19

The fitted equation for the shorter period, 1970(1) - 1975(4), was:

13. $$UR = 5.01 + 0.47USUR - 0.09INFLATN + 3.93COMP - 5.1TOT^{-3} + 1.17S1$$
$$(3.46) \quad (4.82) \quad (1.59) \quad (1.16) \quad (2.45) \quad (8.04)$$

$$+ 0.31S2 - 0.15S3 - 0.0096BWRDUM$$
$$(2.12) \quad (1.02) \quad (0.06)$$

$R^2 = .93$ D.W. = 2.11

MSE = 0.061 DFE = 15

There is however, enough variation in the coefficients of the variables in this equatio to raise the possibility that multi-colinearity may still be a problem.

When the collinearity diagnostics were examined, three condition indices indicated problems, although none involved BWRDUM. Collinearity seemed to exist primarily between USUR and TOT^{-3}. In other cases, the eigenvalues with high condition indices did not contribute a large variance proportion to more than one variable.

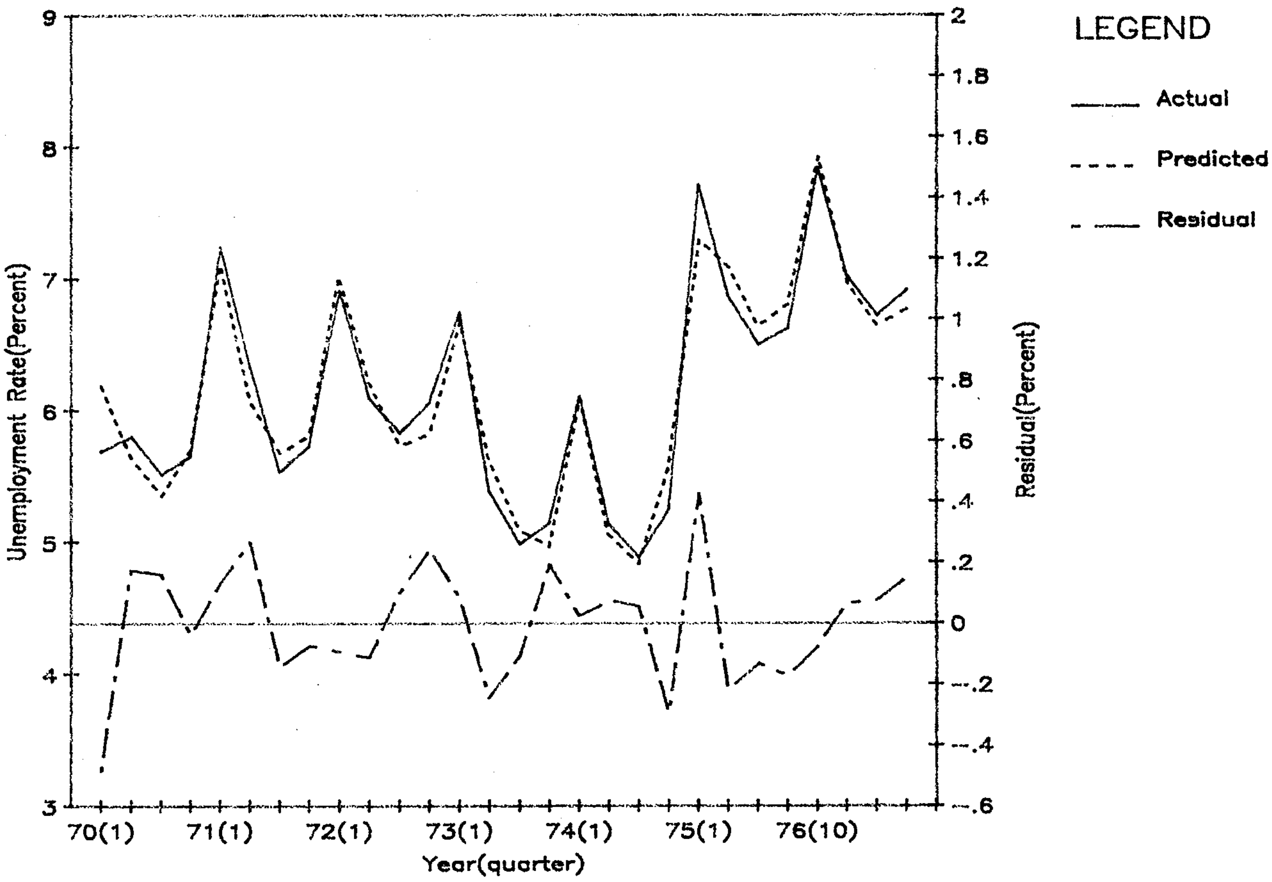
The actual and predicted values of UR for the 1970-1976 period are shown in Figure 6.

The results lead to the conclusion that the relationship between UIB and the unemployment rate is not statistically significant when tested for a structurally uniform period as in this case.

The results presented here lead to sharply different conclusions from most Canadian studies. Indeed they suggest that any studies which include the period prior to 1970 should be regarded with caution, as they may be confusing

Figure 6

Actual and Predicted Unemployment Rates for Canada
1970(1) to 1976(4)



Source:Statistics Canada

a structural change unrelated to UIB with the effect of UIB liberalization in 1971. In particular, caution should be exercised in dealing with studies which try to capture the effect of UIB liberalization through the use of a dummy variable or "quasi-dummy" variable such as the benefit/wage ratio. At the very least, these results suggest that the significance of UI related variables is strongly dependent on the choice of time period and of model.

CHAPTER FIVE

SUMMARY AND CONCLUSIONS

This study was undertaken to investigate the effects of UI liberalization on unemployment in Canada. Quarterly time series data over the period 1959(1) to 1976(4) was used to fit a single equation model of the relation between unemployment and economic and labour market indicators. Two maximum likelihood tests for the location of a structural change in the relationship were carried out. Both tests placed the structural shift well before the UI liberalization of 1971(3). A test performed on the same model for the period after the estimated time of transition led to rejection of the hypothesis that a further shift occurred at the time of the 1971(3) revisions of the UI regulations. A dummy variable introduced to capture the effect of changes in the benefit/wage ratio at this time proved to be not statistically significant.

The fitted model, although not supporting the hypothesis that UI benefits affect measured rates of unemployment, did give the correct direction of effect. The model, moreover, showed signs of being poorly specified for the period 1970(1) to 1976(4). Accordingly, an attempt was made to develop a model, based on the work of Barber and McCallum (1980), which could be fitted to the same period. This model led to a strong rejection of the hypothesis. In addition, the new model appeared to be fairly well specified for the short period. Dropping the last four observations led to only small changes in the coefficients of the included variables and did not improve the performance of the UI related variables.

These results are consistent with the theoretical discussion developed in an earlier chapter which suggested that many other factors might be involved in work/leisure choices which could mask any effect improved benefit ratios might have on work seeking behavior of insured individuals.

The results also point up a flaw in the methods typically used in Canadian and American Studies to assess the impact of UI on unemployment. Most rely heavily on relatively simple models which cannot establish the source and direction of causation among such factors as the level of economic activity, labour supply changes, and the influences of unemployment insurance. Theory cannot give a clear indication of the direction and magnitude of such effects, so regression coefficients are simply disguised correlation coefficients, without causal significance. When these are used to measure the effect of unemployment insurance benefits on unemployment, their magnitude is suspect. When dummy variables are used, their interpretation depends crucially on their location in the time series. This is equally true when disguised dummy variables such as the benefit/wage ratio are used instead of actual dummies or transitional dummies.

These results suggest that it might be fruitful to follow other avenues of analysis of the relationship between unemployment and UI liberalization. One such line of advance might be to consider the relationship between worker militancy during the period 1970 to 1974 and changes in the UI Act. In such a model, the changes might be hypothesized to be in response to a heightening of labour's demands leading to temporary concessions by the state. Another line might be to explore the possibilities that an important function of UI liberalization was to prop up aggregate demand at a time when economic expansion had come to a halt. In both such models, the role of the state in liberalizing UIB is a reactive one, a response to social and economic conditions. This reverses the usual econometric treatment of the activities of the state, wherein, flying in the face of common sense and daily experience, the actions of the state are treated as "exogenous", i.e., not to be explained within the model being considered, but happening much like the weather.

Since 1969 is the beginning of the first major recession in North America since the Great Depression, and marks the beginning of a long period of economic stagnation and high inflation, it is inviting to link this to the apparent unexplained increase in unemployment which is normally attributed to unemployment insurance liberalization. To do this, however, is beyond the scope and intent of this paper.

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APPENDICES

APPENDIX A:

SUMMARY OF 1971 UNEMPLOYMENT INSURANCE ACT REVISIONS

SOME COMPARAISONS BETWEEN OLD AND NEW ACTS

Old Program

Coverage

Certain employments such as teachers, civil servants, members of armed forces, salaried workers over $7,800 per year, are not covered as well as a variety of inconsiderable employments.

In order to protect their insured status, persons may elect to contribute both employer and employee share when their employment becomes excepted by virtue of a salary increase beyond $7,800 per year.

Age is not a factor.

Contributions

Contributions are according to a table with employers and employees paying equally and the government contributing 20% of the combined amount and paying administration costs as well.

A fund is created and contribution rates are adjusted periodically.

Benefits

Waiting period — 1 week.

Regular — Claimant can draw one week for each two weeks of contributions if (a) he has had 30 contributions in the last 104 weeks of which 8 were in the last 52; and (b) he meets on a continuous basis and through bi-weekly reports the conditions of availability, capability and seeking work.

New Program

Coverage

Coverage is universal for all regular members of the labour force for whom there exists an employer-employee relationship. There is only one measure of inconsiderable employment, i.e. less than $30 per week or 20 times the provincial hourly minimum wage, whichever is the lesser.

There is no need for any election as all persons engaged in insurable employment are insured for the first $150 per week.

Universality becomes effective January 2, 1972.

Coverage, contributions and benefit entitlement cease at age 70.

Contributions

Employers and employees absorb the benefit cost for the initial and re-established benefit periods as well as the administration cost, with the employer rate being 1.4 times the employee rate. The government share is confined to the benefit cost for the extended benefit periods as well as the excess cost of benefits for the initial and re-established benefits that are due to a national unemployment rate greater than 4%.

There is no fund and employer and employee contributions are adjusted yearly.

National Revenue/Taxation commence collection of the contributions effective January 2, 1972.

Persons formerly not contributing either because of their occupation or by virtue of being over the salary ceiling will pay a preferred rate for the first 3 years. For those who had been occupationally excluded, the preferred rate is portable. However, in the case of persons formerly excluded because of the salary ceiling, the preferred rate continues only so long as the employee remains with the January 2, 1972, employer.

An experience rating formula for employers may be introduced in 1974 to reflect the additional benefit expense generated by large employers who have above-average layoff experience.

Benefits

Waiting period — 2 weeks.

Regular — The duration of benefits under the new program is not determined solely by the length of attachment to the labour market. A claimant can draw to a maximum of 51 weeks depending on his employment history and prevailing economic conditions, providing (a) he has at least 8 weeks of contributions in the last 52 and (b) he meets the conditions of availability, capability and searching for work.

Persons with 20 or more weeks of insured earnings (major labour force attachment) are eligible for a wider range of benefit that includes a pre-payment of 3 weeks of regular benefit for work-shortage lay-offs; benefit payments when the interruption of earnings was caused by illness or pregnancy; and 3 weeks retirement benefit for older workers.

APPENDIX B:

RESULTS OF DUMMY VARIABLE REGRESSION

Value of t-statistic of Coefficient of Dummy Variable and MSE of Equation for Forty Dummy Variables each fitted to Six Underlying Models Containing a Shift of Intercept at Observation Number Fifty-Three.

Period of Shift of Included Dummy Variable	T-Statistic / Mean Square Error: Equation 1	2	3	4	5	6
29	-1.99	-0.87	-2.20	-1.40	-1.26	-1.89
	.31	.37	.33	1.28	2.26	1.25
30	-1.27	-1.07	-1.74	-1.63	-1.67	-2.10
	.32	.37	.34	1.27	1.24	1.23
31	-0.90	-1.27	-1.912	-1.78	-2.13	-1.81
	0.32	0.37	0.34	1.26	1.21	1.25
32	-0.86	-0.53	-1.57	-1.61	-1.25	-1.35
	0.32	0.38	0.34	1.27	1.26	1.28
33	-0.72	-0.54	-1.24	-1.21	-1.14	-0.57
	0.32	0.37	0.35	1.29	1.27	1.31
34	-0.16	-0.94	-0.79	-1.39	-1.24	-0.56
	0.32	0.37	0.35	1.28	1.26	1.31
35	.12	-1.33	-0.65	-1.63	-0.91	-0.37
	0.32	0.37	0.35	1.27	1.28	1.31
36	0.16	-1.38	-1.04	-1.94	-1.47	-0.17
	0.32	0.37	0.35	1.25	1.27	1.31
37	-0.42	-1.42	-1.31	-0.97	-0.06	-0.74
	0.32	0.36	0.36	1.30	1.28	1.31
38	-0.38	-0.52	-0.60	-0.49	-0.30	-0.65
	0.32	0.37	0.35	1.30	1.29	1.31
39	-0.36	-0.44	-0.26	-0.65	-0.39	-0.69
	0.32	0.37	0.35	1.31	1.29	1.30
40	0.58	-0.22	-0.01	-0.27	-0.06	-0.57
	0.32	0.38	0.35	1.31	1.29	1.31

Period of Shift of Included Dummy Variable	T-Statistic / Mean Square Error — Equation 1	2	3	4	5	6
41	1.36 0.32	0.39 0.37	0.22 0.35	.42 1.31	.58 1.28	-0.12 1.31
42	1.21 0.32	-0.13 0.38	0.91 0.35	.27 1.31	.46 1.29	-0.07 1.31
43	1.30 0.32	-0.19 0.37	1.58 0.34	.17 1.31	-0.11 1.29	0.46 1.31
44	1.47 0.31	90.50 0.37	1.17 0.35	.39 1.31	.19 1.29	.74 1.30
45	2.02 0.31	-0.19 0.38	1.90 0.34	.80 1.30	.51 1.29	.43 1.31
46	2.81 0.29	-0.10 0.38	2.24 0.33	.87 1.30	1.55 1.25	.19 1.31
47	3.01 .29	0.48 .37	2.98 .31	1.60 1.27	2.01 1.22	.21 1.31
48	3.58 .27	0.80 .37	3.49 .30	2.23 1.23	2.11 1.21	1.03 1.29
49	3.16 .28	.92 .37	3.14 .31	2.25 1.23	2.28 1.20	.82 1.30
50	3.74 .27	1.77 .36	3.96 .29	2.41 1.21	2.45 1.19	1.79 1.25
51	4.24 .26	2.24 .35	4.67 .27	3.42 1.12	2.95 1.15	3.04 1.16
52	3.96 .26	3.02 .33	4.54 .27	3.47 1.12	3.57 1.09	3.85 1.08
53	4.33 .25	3.35 .32	5.80 .24	3.31 1.14	4.28 1.02	4.66 1.00
54	3.23 .28	2.62 .34	5.38 .25	2.74 1.19	3.33 1.11	4.01 1.06
55	2.86 .29	2.07 .35	5.33 .25	1.53 1.27	3.34 1.11	3.53 1.11

Period of Shift of Included Dummy Variable	T-Statistic / Mean Square Error — Equation 1	2	3	4	5	6
56	2.42	2.33	5.15	1.43	2.90	2.67
	.30	.35	.26	1.28	1.15	1.19
57	2.12	2.37	4.04	1.29	2.57	2.46
	.30	.35	.29	1.28	1.18	1.21
58	1.76	1.70	3.74	2.20	1.53	1.91
	.31	.36	.29	1.23	1.25	1.24
59	2.19	1.43	3.13	2.05	1.92	1.88
	.30	.36	.31	1.23	1.25	1.25
60	1.93	1.91	2.55	1.88	1.48	2.35
	.31	.36	.32	1.25	1.25	1.22
61	2.71	1.87	2.78	2.16	1.11	2.32
	.29	.36	.32	1.23	1.27	1.22
62	2.70	1.27	2.49	1.27	.36	2.40
	.29	.37	.33	1.29	1.29	1.21
63	2.32	2.02	2.32	1.29	.13	1.81
	.30	.35	.33	1.28	1.29	1.25
64	2.49	2.07	2.39	1.26	-0.36	1.22
	.30	.35	.33	1.29	1.29	1.29
65	2.51	1.27	1.74	1.95	0.03	1.46
	.30	.37	.34	1.25	1.29	1.27
66	2.13	1.24	1.78	1.94	-0.07	.57
	.30	.37	.34	1.25	1.29	1.31
67	2.31	1.11	2.26	1.58	-0.33	.78
	.30	.37	.33	1.27	1.29	1.30
68	1.76	0.76	1.89	1.20	.28	.90
	.31	.37	.34	1.29	1.29	1.30

APPENDIX C:

SOURCE DATA

LISTING OF VARIABLES INCLUDED IN ANALYSIS

VARIABLES FOR EQUATIONS 1 THROUGH 10

PERIOD	GNE	GNEP	GO	LF	LFPRED	LFRESID	UQ	BWRDUM
1959(1)	11685	12377.4	-692.4	6092.7	5987.8	104.82	566.670	0.0
1959(2)	12511	12540.0	-29.0	6208.7	6033.4	175.27	356.000	0.0
1959(3)	14445	12704.8	1740.2	6397.0	6079.3	317.72	240.000	0.0
1959(4)	13096	12871.7	224.3	5935.0	6125.5	-190.52	323.670	0.0
1960(1)	12311	13040.8	-729.8	6229.0	6172.1	56.88	583.330	0.0
1960(2)	12706	13212.1	-506.1	6376.7	6219.1	157.60	426.670	0.0
1960(3)	14862	13385.7	1476.3	6570.3	6266.4	303.96	334.330	0.0
1960(4)	13352	13561.5	-209.5	6468.0	6314.0	153.97	439.000	0.0
1961(1)	12278	13739.7	-1461.7	6375.0	6362.1	12.94	702.670	0.0
1961(2)	13374	13920.2	-546.2	6527.3	6410.5	116.88	480.000	0.0
1961(3)	14948	14103.0	845.0	6665.7	6459.2	206.46	325.330	0.0
1961(4)	14141	14288.3	-147.3	6516.7	6508.3	8.32	357.670	0.0
1962(1)	13323	14476.0	-1153.0	6434.3	6557.8	-123.51	561.330	0.0
1962(2)	14017	14666.2	-649.2	6284.7	6607.7	-323.06	379.670	0.0
1962(3)	16296	14858.9	1437.1	6812.3	6658.0	154.35	281.670	0.0
1962(4)	14839	15054.1	-215.1	6608.0	6708.6	-100.63	346.000	0.0
1963(1)	13823	15251.9	-1428.9	6512.0	6759.7	-247.66	545.670	0.0
1963(2)	14712	15452.2	-740.2	6697.0	6811.1	-114.08	371.670	0.0
1963(3)	17134	15655.2	1478.8	6967.3	6862.9	104.45	272.000	0.0
1963(4)	15818	15860.9	-42.9	6816.7	6915.1	-98.42	305.000	0.0
1964(1)	15023	16069.3	-1046.3	6718.0	6967.7	-249.68	463.000	0.0
1964(2)	15930	16280.4	-350.4	6915.0	7020.7	-105.68	326.000	0.0
1964(3)	17923	16494.3	1428.8	7149.0	7074.1	74.92	242.670	0.0
1964(4)	16734	16710.9	23.1	6950.7	7127.9	-177.22	266.000	0.0
1965(1)	15909	16930.5	-1021.5	6877.7	7182.1	-304.44	397.000	0.0
1965(2)	16971	17152.9	-181.9	7132.3	7236.7	-104.40	297.670	0.0
1965(3)	19100	17378.2	1721.8	7381.3	7291.8	89.55	210.330	0.0
1965(4)	18001	17606.5	394.5	7174.3	7347.2	-172.91	217.670	0.0
1966(1)	17195	17837.8	-642.8	7154.7	7403.1	-248.46	303.333	0.0
1966(2)	18319	18072.2	246.8	7447.0	7459.4	-12.44	236.667	0.0
1966(3)	20513	18309.6	2203.4	7787.3	7516.2	271.15	235.667	0.0
1966(4)	18817	18550.1	266.9	7585.0	7573.4	11.65	230.000	0.0
1967(1)	17860	18793.8	-933.8	7424.7	7631.0	-206.29	340.333	0.0
1967(2)	19116	19040.8	75.3	7724.0	7689.0	35.00	293.667	0.0
1967(3)	20812	19290.9	1521.1	8044.0	7747.5	296.52	260.000	0.0
1967(4)	19556	19544.3	11.7	7794.7	7806.4	-11.75	290.000	0.0
1968(1)	18543	19801.1	-1258.1	7567.7	7865.8	-298.13	410.667	0.0
1968(2)	20043	20061.2	-18.2	7922.7	7925.6	-2.95	365.333	0.0
1968(3)	22176	20324.8	1851.2	8246.7	7985.9	260.76	332.000	0.0
1968(4)	21102	20591.8	510.2	8067.7	8046.6	21.02	324.000	0.0
1969(1)	19896	20862.3	-966.3	7890.0	8107.9	-217.86	405.000	0.0
1969(2)	21133	21136.4	-3.4	8243.7	8169.5	74.14	368.333	0.0
1969(3)	23186	21414.0	1772.0	8439.0	8231.7	207.33	329.000	0.0
1969(4)	22010	21695.4	314.6	8202.7	8294.3	-91.61	346.000	0.0
1970(1)	20526	21980.4	-1454.4	7993.7	8357.4	-363.70	455.333	0.0
1970(2)	21875	22269.1	-394.2	8432.3	8420.9	11.40	489.67	0.0

1970(3)	23582	22561.7	1020.3	8692.7	8485.0	207.68	479.000	0.0
1970(4)	22407	22858.1	-451.1	8461.3	8549.5	-88.20	479.333	0.0
1971(1)	21428	23158.4	-1730.4	8288.3	8614.6	-326.23	600.667	0.0
1971(2)	23178	23462.6	-284.6	8599.3	8680.1	-80.75	544.667	0.0
1971(3)	25376	23770.9	1605.1	8925.0	8746.1	178.89	493.667	0.6
1971(4)	24468	24083.2	384.8	8742.0	8812.6	-70.63	501.667	1.0
1972(1)	22946	24399.5	-1453.5	8542.0	8879.7	-337.66	590.667	1.0
1972(2)	24907	24720.1	186.9	8874.0	8947.2	-73.21	540.333	1.0
1972(3)	26323	25044.8	1278.1	9196.7	9015.3	181.41	536.000	1.0
1972(4)	26072	25373.9	698.1	8976.7	9083.8	-107.17	544.667	1.0
1973(1)	24917	25707.2	-790.2	8877.0	9152.9	-275.93	599.333	1.0
1973(2)	26581	26044.9	536.1	9313.3	9222.5	90.78	501.333	1.0
1973(3)	28170	26387.1	1782.9	9545.7	9292.7	252.97	475.667	1.0
1973(4)	28144	26733.8	1410.2	9369.0	9363.4	5.62	482.667	1.0
1974(1)	26430	27085.0	-655.0	9231.3	9434.6	-203.27	564.333	1.0
1974(2)	27831	27440.8	390.2	9614.0	9506.4	107.63	493.667	1.0
1974(3)	28926	27801.3	1124.7	9984.0	9578.7	405.33	487.667	1.0
1974(4)	28491	28166.5	324.5	9728.3	9651.5	76.80	511.667	1.0
1975(1)	26298	28536.6	-2238.6	9558.7	9724.9	-166.28	737.667	1.0
1975(2)	27913	28911.4	-998.4	9984.0	9798.9	185.08	685.000	1.0
1975(3)	29760	29291.3	468.7	10307.3	9873.4	433.88	670.000	1.0
1975(4)	29034	29676.1	-642.1	10045.3	9948.6	96.78	666.000	1.0
1976(1)	27817	30065.9	-2248.9	9883.7	10024.2	-140.56	776.000	1.0
1976(2)	29962	30460.9	-498.9	10203.7	10100.5	103.20	717.000	1.0
1976(3)	31593	30861.1	731.9	10551.3	10177.3	374.04	709.000	1.0
1976(4)	30240	31266.5	-1026.5	10186.0	10254.7	-68.71	704.667	1.0

LISTING OF VARIABLES INCLUDED IN ANALYSIS

VARIABLES USED IN CONSTRUCTING THE VARIABLE FISC

PERIOD	FISC	SURPLUS
1968(1)	50.99	0.275
1968(2)	80.17	0.400
1968(3)	177.41	0.800
1968(4)	253.22	1.200
1969(1)	318.34	1.600
1969(2)	422.66	2.000
1969(3)	428.94	1.850
1969(4)	374.17	1.700
1970(1)	318.15	1.550
1970(2)	306.25	1.400
1970(3)	259.40	1.100
1970(4)	179.26	0.800
1971(1)	107.14	0.500
1971(2)	46.36	0.200
1971(3)	57.10	0.225
1971(4)	61.17	0.250
1972(1)	63.10	0.275
1972(2)	74.72	0.300
1972(3)	52.65	0.200
1972(4)	26.07	0.100
1973(1)	0.00	0.000
1973(2)	-26.58	-0.100
1973(3)	-28.17	-0.100
1973(4)	-28.14	-0.100
1974(1)	-26.43	-0.100
1974(2)	-27.83	-0.100
1974(3)	43.39	0.150
1974(4)	113.96	0.400
1975(1)	170.94	0.650
1975(2)	251.22	0.900
1975(3)	-52.08	-0.175
1975(4)	-159.69	-0.550
1976(1)	382.48	1.375
1976(2)	-599.24	-2.000
1976(3)	-600.27	-1.900
1976(4)	-544.32	-1.800

LISTING OF VARIABLES INCLUDED IN ANALYSIS

VARIABLES FOR EQUATIONS 11 THROUGH 13

PERIOD	UR	USUR	COMP	TOT	FISC	BWRDUM	D1	D2	D3
1968(1)	5.42660	4.0	0.877275	1.02381	50.99	0.0	1	0	0
1968(2)	4.61124	3.5	0.879415	1.01615	80.17	0.0	0	1	0
1968(3)	4.02587	3.6	0.869474	1.01398	177.41	0.0	0	0	1
1968(4)	4.01603	3.5	0.869095	1.00320	253.22	0.0	0	0	0
1969(1)	5.13308	3.4	0.867929	1.01481	318.34	0.0	1	0	0
1969(2)	4.46808	3.5	0.856208	1.01466	422.66	0.0	0	1	0
1969(3)	3.89857	3.6	0.863886	1.00312	428.94	0.0	0	0	1
1969(4)	4.21814	3.6	0.857534	1.00000	374.17	0.0	0	0	0
1970(1)	5.69618	4.2	0.838789	1.00811	318.15	0.0	1	0	0
1970(2)	5.80701	4.7	0.823312	1.01515	306.25	0.0	0	1	0
1970(3)	5.51039	5.2	0.828394	1.02058	259.40	0.0	0	0	1
1970(4)	5.66499	5.8	0.823707	1.02045	179.26	0.0	0	0	0
1971(1)	7.24713	6.0	0.832947	1.00102	107.14	0.0	1	0	0
1971(2)	6.33382	5.9	0.838386	0.99799	46.36	0.0	0	1	0
1971(3)	5.53128	6.0	0.831747	1.00000	57.10	0.6	0	0	1
1971(4)	5.73858	5.9	0.839396	1.00298	61.17	1.0	0	0	0
1972(1)	6.91485	5.8	0.843717	1.00000	63.10	1.0	1	0	0
1972(2)	6.08895	5.7	0.848405	1.00487	74.72	1.0	0	1	0
1972(3)	5.82820	5.6	0.867699	1.00872	52.65	1.0	0	0	1
1972(4)	6.06758	5.3	0.880985	1.01729	26.07	1.0	0	0	0
1973(1)	6.75153	5.0	0.868647	1.02728	0.00	1.0	1	0	0
1973(2)	5.38296	4.9	0.854766	1.03822	-26.58	1.0	0	1	0
1973(3)	4.98306	4.8	0.856185	1.05737	-28.17	1.0	0	0	1
1973(4)	5.15174	4.7	0.850508	1.09983	-28.14	1.0	0	0	0
1974(1)	6.11324	5.0	0.844476	1.14391	-26.43	1.0	1	0	0
1974(2)	5.13487	5.1	0.838166	1.15615	-27.83	1.0	0	1	0
1974(3)	4.88448	5.6	0.855469	1.13256	43.39	1.0	0	0	1
1974(4)	5.25955	6.6	0.868904	1.11525	113.96	1.0	0	0	0
1975(1)	7.71725	8.1	0.874691	1.08632	170.94	1.0	1	0	0
1975(2)	6.86098	8.8	0.908017	1.08469	251.22	1.0	0	1	0
1975(3)	6.50023	8.6	0.942500	1.09737	-52.08	1.0	0	0	1
1975(4)	6.62994	8.4	0.925746	1.12062	-159.69	1.0	0	0	0
1976(1)	7.85134	7.7	0.940964	1.11248	382.48	1.0	1	0	0
1976(2)	7.02689	7.6	0.967761	1.11693	-599.24	1.0	0	1	0
1976(3)	6.71953	7.7	0.962963	1.12516	-600.27	1.0	0	0	1
1976(4)	6.91799	7.7	0.975287	1.12172	-544.32	1.0	0	0	0

LISTING OF VARIABLES INCLUDED IN ANALYSIS

VARIABLES USED IN CONSTRUCTING THE VARIABLES TOT AND COMP

PERIOD	TOT	CANIMPDX	CANIMDI	COMP	USUNIT	CANLCI
1968(1)	1.02381	94.6	92.4	0.877275	102.6	90.008
1968(2)	1.01615	94.4	92.9	0.879415	103.1	90.668
1968(3)	1.01398	94.3	93.0	0.869474	104.7	91.034
1968(4)	1.00320	94.1	93.8	0.869095	106.6	92.645
1969(1)	1.01481	95.9	94.5	0.867929	108.6	94.257
1969(2)	1.01466	96.9	95.5	0.856208	110.6	94.697
1969(3)	1.00312	96.3	96.0	0.863886	112.5	97.187
1969(4)	1.00000	96.5	96.5	0.857534	114.7	98.359
1970(1)	1.00811	99.5	98.7	0.838789	117.7	98.725
1970(2)	1.01515	100.5	99.0	0.823312	118.4	97.480
1970(3)	1.02058	99.2	97.2	0.828394	119.0	98.579
1970(4)	1.02045	99.8	97.8	0.823707	121.1	99.751
1971(1)	1.00102	98.4	98.3	0.832947	120.9	100.703
1971(2)	0.99799	99.5	99.7	0.838386	122.3	102.535
1971(3)	1.00000	101.0	101.0	0.831747	123.1	102.388
1971(4)	1.00298	101.0	100.7	0.839396	123.2	103.414
1972(1)	1.00000	102.4	102.4	0.843717	125.0	105.465
1972(2)	1.00487	103.1	102.6	0.848405	125.0	106.051
1972(3)	1.00872	104.1	103.2	0.867699	124.5	108.028
1972(4)	1.01729	105.9	104.1	0.880985	125.2	110.299
1973(1)	1.02728	109.2	106.3	0.868647	127.4	110.666
1973(2)	1.03822	114.1	109.9	0.854766	132.2	113.000
1973(3)	1.05737	119.8	113.3	0.856185	134.2	114.900
1973(4)	1.09983	126.7	115.2	0.850508	137.8	117.200
1974(1)	1.14391	139.9	122.3	0.844476	142.1	120.000
1974(2)	1.15615	150.3	130.0	0.838166	148.3	124.300
1974(3)	1.13256	157.2	138.8	0.855469	153.6	131.400
1974(4)	1.11525	161.6	144.9	0.868904	157.9	137.200
1975(1)	1.08632	163.6	150.6	0.874691	162.0	141.700
1975(2)	1.08469	166.5	153.5	0.908017	160.9	146.100
1975(3)	1.09737	171.3	156.1	0.942500	160.0	150.800
1975(4)	1.12062	172.8	154.2	0.925746	164.3	152.100
1976(1)	1.11248	172.1	154.7	0.940964	166.0	156.200
1976(2)	1.11693	172.9	154.8	0.967761	167.5	162.100
1976(3)	1.12516	175.3	155.8	0.962963	170.1	163.800
1976(4)	1.12172	175.1	156.1	0.975287	174.0	169.700