1983

Age-structure And Unemployment: Some Consequences Of The Post-war Baby Boom

David Kerr Forrest

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LA THÈSE A ÉTÉ MICROFILMÉE TELLE QUE NOUS L'AVONS RECEUE
AGE-STRUCTURE AND UNEMPLOYMENT:
SOME CONSEQUENCES OF THE POST-WAR BABY BOOM

by

David Kerr Forrest

Department of Economics

Submitted in partial fulfillment
of the requirements for the degree of
Doctor of Philosophy

Faculty of Graduate Studies
The University of Western Ontario

London, Ontario

March, 1983

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Abstract

The paper examines the impact of demographic change on the rate of unemployment in Canada. The central question asked is whether the increase in the number of young people in the population during the 1970s put upward pressure on the overall unemployment-rate. It was expected to have done so because (i) a greater fraction of the labour-force would now fall within an unemployment-prone age-category, and (ii) in the absence of sufficient wage flexibility, a shift in the supply curve of a particular type of labour (defined by age) would generate unemployment in that age-group (this is the so-called cohort-crowding hypothesis). Through these two channels (termed, respectively, the weights and rates effects), and also by influencing participation-rates, changes in the age-composition of the population were predicted to alter the overall unemployment-rate. The theoretical rationale for this prediction is extensively reviewed.

The empirical analysis uses Canadian labour-force data covering 1953-78. The weights effect was estimated by statistical decomposition of changes in the unemployment-rate. To estimate the rates effect, multiple regression techniques were employed. Eight equations were estimated, each 'explaining' the unemployment-rate of a specific demographic group, defined by age and sex. Amongst the independent variables was the proportion of the adult population which fell within the particular demographic category. The performance of this variable was used to test the cohort-crowding hypothesis and to make estimates of the impact of a group's relative size on its group-specific (and thereby on the overall) rate of unemployment.
The results of the empirical analysis indicate that the increase in the number of young people in Canada had only a small effect on the overall unemployment-rate but a major effect on the structure of unemployment. Youth unemployment increased as a result of the entry into the labour-force of the baby-boom cohorts; older male workers experienced a fall in unemployment.

A less detailed empirical analysis is presented for the United States. The results are broadly similar to those for Canada.
ACKNOWLEDGEMENTS

First, I must acknowledge the valuable advice of the members of my thesis Committee: Professors Chris Robinson, David Laidler and Edward Saraydar. Next, because the work was written during periods spent on the faculties of McMaster, Liverpool and Manchester Universities, I am greatly indebted to colleagues at those institutions for their generosity in reading and commenting on earlier drafts of the several chapters. In this connection I must thank all of the following: Alan Harrison, Robert Hart, Leslie Bobb, Michael Vealí, Simon Blackman, Barry Naisbitt, Simon Damberger, Ian Walker and Richard Blundell. I am particularly grateful to Christos Ioannidis for fruitful discussions on econometric techniques, and to Michele McWilliams and Vivienne Oakes for fast and efficient typing. Lastly, my greatest debt must be to my parents without whose encouragement my university studies could not have begun. The work is therefore dedicated to the memory of my father.
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CHAPTER I

UPWARD TRENDS IN THE RATE OF UNEMPLOYMENT

1. An Overview

In most Western countries, unemployment-rates were higher, on average, during the nineteen-seventies than they were during the nineteen-fifties and nineteen-sixties. Possible reasons for the increase in unemployment constitute the subject matter of this paper. This first chapter begins by presenting evidence on the extent of the rise in unemployment, stressing particularly the international nature of the phenomenon. Next it reviews (theoretically and empirically) the currently popular hypothesis that changes in unemployment - insurance arrangements have been responsible for much of the rise in unemployment. This hypothesis, however, will be shown to be incapable of explaining all of the increase in unemployment. In the second half of the chapter, we therefore introduce another hypothesis designed to explain rises in unemployment-rates: this is the hypothesis that demographic change played a role in generating increases in unemployment. This latter hypothesis will in fact be the subject of the rest of the paper (beyond Chapter I). It will be examined theoretically in Chapters II and III and empirically in Chapters IV and V. Finally, Chapter VI will summarise our findings and discuss their implications.

We begin then by considering the extent of the rise in unemployment in Canada (the country which will be the main empirical
focus of the paper). Chart One shows the annual unemployment-rate for each year from 1953 to 1978. Figures shown are based on series of labour force statistics for this period, supplied by Professor Frank Denton and his co-workers at McMaster University.¹ It is evident from the chart that Canada has generally experienced a secular rise in unemployment-rates – except for unusually high unemployment-rates in 1958-61, the rate of unemployment has shown a consistent upward trend over the whole of the 1953-78 period.

Canada has not been alone in experiencing a trend rise in the rate of unemployment. Table 1-1 and Chart Two summarise the experiences of eight Western industrialised countries for which unemployment figures are published by the International Labour Organisation. For most of the eight countries, the mean rate of unemployment increased consistently from one semi-decade to the next over 1960-78. In every case, save that of Sweden, the rate of unemployment was sharply higher in the late nineteen-seventies than it had been in the earlier years.

The data cited above point to a swelling in the ranks of the unemployed during the last few years. In most countries, this has thrust unemployment back into the centre of the stage of political, social and economic debate. This study, which considers the trend-increase in unemployment, should therefore be of interest to an audience wider than that of professional economists alone.

---

¹ The methodology for the generation of these statistics is described in Appendix I. Appendix II tabulates the values for the principal series employed in this paper.
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1962-65 only  
Sources: Canada: See Appendix 1  
Other Countries: Year Book of Labour Statistics (I.L.O., various issues)
Chart 1: Percentage Unemployment Rate, Canada, 1953-78 (Annual Averages)
Chart 2: Unemployment Rates in Eight Industrial Countries, 1960-78
2. Unemployment as an International Problem

Given that unemployment rates have shown an upward trend in nearly all countries since the mid- or late-nineteen-sixties, any explanation of increased unemployment will lack plausibility unless it can account for the international nature of the problem.

An international explanation could take either of two forms:

(a) it could point to factors in the international economy (trends in trade, currency upheavals, etc.);

(b) it could point to trends within domestic economies, trends that are apparent in several countries simultaneously.

A currently popular theory takes approach (b). It points to the increasingly generous unemployment-insurance benefits introduced in most Western economies over the past decade.

3. Unemployment Insurance and the Rate of Unemployment

The effect of unemployment benefit on the rate of unemployment has been a matter of increasing interest in recent years. Such interest may be explained by the historically large increase in benefits that has been made available in several countries and by the fact that these increases in benefit have coincided with increases in unemployment-rates. Attempts to draw a link between benefit-rates and unemployment-rates have used as their theoretical basis either the traditional analysis of the allocation of time as between labour and leisure or the job-search approach.
popularised by the Phelps volume (1970). We first consider the impact of unemployment-insurance schemes on the labour-leisure trade-off.

**Unemployment-Insurance and the Labour-Leisure Trade-Off**

In this section, the choice to be analysed is that of an individual worker choosing how many weeks per year to spend in employment and how many weeks per year to allocate to other activities (these other activities we subsume under the title 'leisure'). The worker is assumed to be able to choose freely for himself the number of work-weeks and to be able to do so without facing any costs from moving in and out of employment. In the absence of an unemployment-insurance scheme, and with non-labour income equal to zero, the income-leisure frontier is AB in Figure 1-1: the economic agent whose choice is depicted here will receive no income if he takes 52 weeks of leisure (point A); if he works every week of the year, he will be at point B, receiving an income of 52w (where w is the relevant weekly wage); alternatively, he may choose any point along the frontier AB, the slope of which is equal to the (negative of) the weekly wage.

We now examine the impact of an unemployment-insurance scheme on this income-leisure opportunity locus for our "representative" worker. The form that the impact will take will depend upon the rules of the particular scheme. Most schemes only allow benefit to be paid to those who work some minimum number of weeks per year: for example, for most of the period of operation of the 1971 Unemployment Insurance Act (of Canada), this qualifying period has been set at eight weeks. Our examination of the effect of a benefit
Figure 1-1: The Effect of Unemployment-Insurance on the Income-Leisure Opportunity Locus of an Individual Worker
scheme is framed in the context of such an eight week qualifying period. Subject to this qualification, it is assumed that benefit is payable for any week in which the worker is not at work. It is also assumed that workers do not bear any of the costs of the scheme (for example, through lower wage-rates).

Consider again, then, the income-leisure choice depicted in Figure 1-1. Where the number of work-weeks is less than eight, no benefit is payable and the income corresponding to any given number of leisure-weeks is the same as when there was no unemployment-insurance system: AC, therefore, remains as part of the income-leisure frontier. However, for any number of weeks of leisure in the range 0-44 weeks, there is, corresponding to any given amount of leisure, a greater amount of income than before. For example, a person working eight weeks per year will previously have had an annual income of 8w but will now receive 8w + 44b, where b is weekly unemployment-insurance benefit. Across the range 0-44 weeks of leisure, there is, therefore, a shift in the income-leisure frontier. Of course, the intercept on the vertical axis will remain at 52w because a person taking no leisure will receive no unemployment-benefit. The new income-leisure frontier is, therefore, ACDB. The slope of the frontier is (the negative of) the opportunity cost of a week of leisure. This opportunity cost is w in the range AC but only (w-b) in the range DB. For workers employed for at least eight weeks per year, the change in the frontier thus reflects a fall in the opportunity cost of leisure.
How will a worker's behaviour change in response to these changes in his leisure-income trade-off possibilities? For a worker already employed for at least eight weeks per year (i.e., for one who maximises utility by locating at a point on the line BC), there will be an increase in the number of leisure-weeks. The result assumes a diminishing marginal rate of substitution between income and leisure; also, it is assumed that leisure is a normal good. The substitution effect raises the number of leisure-weeks because the opportunity cost of leisure has fallen. The income effect is in the same direction because the worker is able to attain a higher indifference curve by moving to a point along BD and, if leisure is a normal good, more of it will be consumed at the higher real income level. Thus, the number of weeks of leisure is unambiguously increased. Amongst a labour-force composed entirely of such workers, the unemployment-rate will rise so long as the extra leisure-weeks are counted as 'unemployment-weeks' (as they are likely to be, since a worker can only claim benefit if he describes his leisure as unemployment).

The overall unemployment-rate will also be affected by the behaviour of those who worked less than eight weeks per year before the introduction of the benefit scheme. These workers previously located at point A (zero work) or along the line AC. Some of these workers will be induced to reduce their number of leisure-weeks by moving to a higher indifference curve at a point where it is tangent to the section BD of the new income-leisure opportunity locus. If all such workers were previously counted as part of the labour-force for the whole year, the result will be to reduce the recorded rate-of-unemployment. However, for those who worked not at all prior to the scheme, it is likely
that they will be entering the recorded labour-force for the first time; and even for those working up to eight weeks in the initial situation, it is possible that weeks previously counted as weeks "not in the labour-force" will now be counted as "weeks of unemployment" (they must be declared as such to qualify for benefit). To the extent that these changes are recorded in the way suggested, there will be an increase in the labour-force participation-rate. This increase will raise the overall unemployment-rate if the new entrants have higher-than-average unemployment experience - as seems likely given that the workers in question are those who place the highest value on leisure.

The number of people who move from zero employment to an annual employment of eight weeks will, of course, depend both on the generosity of the scheme and on the strength of preference for leisure amongst the population in general. In the context of the Canadian system as it stood in August, 1977, Grubel and Walker (1978, p. 11) show that for a worker able to obtain the average weekly Canadian wage, the unemployment-insurance scheme raised the income attached to eight weeks work per year from $2,080 to $8,548. This is a decidedly non-trivial change in income possibilities; the number of persons entering the labour-force in response to the unemployment-insurance scheme, therefore, seem likely to be large.

Whether one considers only those who previously worked more than eight weeks per year or whether one also considers those previously working less than eight weeks per year, the analysis yields a prediction that the introduction of an unemployment-insurance system will raise the recorded rate-of-unemployment. Of course, not all workers will behave
in the way predicted: the scale of job-switching costs (assumed zero in this analysis) will, in fact, inhibit some of those who would otherwise have benefited by taking more leisure. Nevertheless, traditional labour-leisure trade-off analysis appears to provide a theoretical foundation for the proposition that unemployment-insurance schemes raise unemployment-rates.

Not all workers will take more unemployment in response to unemployment-insurance schemes. In some cases, the job-switching costs (e.g., the job-search costs at the end of the leisure period) will be sufficiently high to deter the worker from quitting. This brings us to the job-search framework for analysing unemployment. Most explanations of the link between unemployment-benefits and unemployment-rates have, in fact, been presented within this latter framework.

**Unemployment-Insurance and the Duration of Search Unemployment**

The job-search concept was popularised by the Phelps volume (1970). Contributions to this book focused on voluntary unemployment that was taken by those experiencing it in order to 'specialise' full-time in job-search activity. The key assumption was that full-time job-search was a productive activity: it yielded information about different employments, and prolonging search always involved the possibility of locating a job that gave better overall returns (in income or in other respects) than those jobs about which information had previously been gathered.
A simple model of job-search considers a worker who happens to become unemployed. How long will he remain unemployed if his aim is to maximise the expected present value of life-time income and if a job could be obtained at any time so long as the worker had acquired information about that job? His decision on this matter will depend on both the benefits and the costs associated with job-search. Ignoring search costs, the marginal benefit of another week of job-search is the expected increase in the present value of life-time income from spending a week in gathering information about jobs. The marginal benefit would be expected to decline with the duration of job-search: as time passes, information is gathered on a greater and greater proportion of the pool of possible jobs and the probability of finding a job superior to those already considered becomes less and less.¹

Now consider the marginal cost of the nth week of job-search. This will consist of the income foregone from spending the nth week in job-search rather than employed in the most satisfactory job located in the first n-1 weeks of search. Marginal cost is expected to rise with time since, if better jobs are located by search-activity, the best weekly wage available will increase with length of search. Our income-maximising worker will continue search until the marginal cost of search is equal to its marginal benefit: this will give the equilibrium.

¹ In the limit, when information about all jobs has been obtained, there are no further benefits from search.
unemployment-duration for our worker. His decision will, of course, be changed by anything that shifts either the marginal cost or the marginal benefit function. The introduction of an unemployment-insurance system reduces the marginal cost of a week of search because income foregone is now equal to the weekly wage that could have been earned minus the unemployment benefit paid for that week. This shift in the marginal-cost-of-search function will raise the equilibrium search duration chosen by the optimising worker.

The analysis, then, predicts that the availability of unemployment-benefits will raise the expected duration of a given spell of unemployment. The unambiguity of the prediction is, however, only achieved by ignoring the qualifying-rules attached to all unemployment-insurance programmes. A worker must spend a specified number of weeks in employment in order to qualify for benefit. Mortensen (1977) pointed out that for workers not yet qualified for benefit, or for those whose benefit-rights have been almost exhausted, the presence of an insurance system may shorten the duration of unemployment. The reason is that the benefit from taking a job at any time will now include not only the wage-income from the job but also the value placed on the rights (to be acquired on-the-job) to draw unemployment-benefit in the future. This effect is most likely to be important for new labour-force entrants, none of whom will yet have acquired benefit-rights.

1. In a graph, the equilibrium would occur where a rising marginal cost curve cut a falling marginal benefit curve.
The overall effect of an unemployment-insurance system on expected unemployment-duration for a group of workers who would have become unemployed in any case, is ambiguous on the basis of the foregoing analysis. However, in a system such as that in force in Canada, the likely effect on mean duration is to prolong it: because of a short qualifying period, the proportion of workers entering unemployment without benefit entitlement is small and it is only in this group that behaviour will be such as to shorten job-search. The presumption made here is, therefore, that unemployment-benefit provision lengthens the expected duration of a given spell of unemployment.

What are the consequences of increased duration for the overall unemployment-rate? The overall unemployment-rate for a year may be expressed as the product of the number of spells of unemployment and the average duration (in weeks) of a spell of unemployment divided by 52 times the labour-force. From this identity, it is clear that if the average duration of a spell of unemployment is increased by the introduction of (or rise in) unemployment-benefit, and if there is no effect either on the number of spells of unemployment or on the size of the labour-force, then a rise in unemployment-benefit necessarily raises the unemployment-rate.

Of course, the *ogetris paribus* assumptions just specified are unlikely to be fulfilled. A discussion of how unemployment-benefit affects the unemployment-rate, therefore, requires a consideration of its impact on the number of spells of unemployment and on the size of the labour-force.
The Effect of Unemployment-Insurance on the Number of Spells of Unemployment

The number of spells of unemployment will be altered if unemployment-insurance provision changes either the number of voluntary quits or the number of lay-offs. The effect on the number of quits is ambiguous for the same reasons as the effect on mean unemployment-duration was ambiguous. A worker who maximises expected life-time income will quit if the expected income gain from locating a higher paying job by full-time search is greater than the cost of search in terms of income foregone. Unemployment-benefit lowers the opportunity-cost of full-time search and therefore raises the proportion of workers who can benefit from quitting. The unambiguity of this result depends, however, on ignoring benefit-qualification rules: most systems disallow or delay the start of payments to workers who have left jobs voluntarily; to this extent, quits are made less likely because some workers may choose to wait for a possible lay-off rather than quit. The overall effect on quits is therefore, strictly ambiguous. However, a presumption will be made here that the effect will be to increase the true quit-rate in any country with rules similar to those in force in Canada. In Canada, workers are entitled to benefits either if they were laid off "with cause" (e.g. because their work was below acceptable standards) or if they can persuade the government official handling their claim that conditions in the former job were "intolerable". With respect to the former provision, Grubel and Walker (1977, p. 19) point out that, since contributions by employers to the Canadian unemployment-insurance system are unrelated to their record of laying-off workers, they have no incentive to deny that workers left their jobs "with cause" - to say otherwise would penalise
employees who quit and thereby make it more difficult to recruit new employees at the current wage. We, therefore, assume here that the effect of benefit provision is definitely to increase the true number of quits (though these may be classified as lay-offs in the official figures). The assumption would be more questionable in a country where rules were more harsh in the treatment of those making voluntary job-separations.

Turning to the employer's side, it seems likely that there will be a greater incentive to lay-off workers the more generous is a country's unemployment-benefit scheme. A simple model generating this prediction assumes an industry facing a demand pattern that is unstable but predictable. A firm has to decide whether to lay-off workers who are not needed to maintain output during slack-periods. A particular group of workers will experience lay-off if the benefit to the firm is greater than the cost of lay-off. The benefit is, of course, the saving in current wage costs. The cost arises because income-maximising workers will join firms that regularly lay-off workers only if the wage-rate during employment periods is sufficient to compensate for income losses during lay-off periods. Lay-offs, therefore, carry the cost to the firm of raising payments to workers during those periods when the firm's labour-force is fully employed. The decision of a profit-maximising firm on how many workers to lay-off will depend on its assessment of these benefits and costs. If unemployment-insurance payments are increased, the costs of laying-off workers will be lowered: corresponding to any given pattern of lay-offs, the supply price of labour will be reduced. This is because income-maximising workers will need less wage-rate compensation for anticipated periods of lay-off. The overall
result is to raise the proportion of workers whom firms will find it worthwhile to lay-off.

The preceding argument is based on a partial equilibrium model. A general equilibrium model in the same spirit divides the economy into stable- and unstable-demand industries. Ceteris paribus, unstable industries generate more lay-offs relative to their labour-force because their output is subject to seasonal or cyclical influences. The supply price of labour to these industries reflects the expectation of workers that their incomes will be reduced from time to time by lay-off. Unemployment-benefit reduces the degree of income loss from this source and therefore reduces the supply-price of labour to the unstable sector which, because it now faces lower factor costs, will expand its share of output and employment. A change in industry-mix in favour of the unstable sector raises the proportion of workers subject to cyclical or seasonal lay-off - an outcome reinforcing the partial equilibrium result that (so long as firms do not fully bear the costs of paying benefits), the introduction or increase of benefit payments

1. A formal model of this sort is provided by Feldstein (1976).
2. See Topel and Welch (1980).
3. In Canada and Britain, firms' contributions do not depend on whether they have laid-off workers in the past; in the U.S.A., firms partially bear the costs of lay-off through a system that varies contributions according to an 'experience rating'.
raises the number of lay-offs.\textsuperscript{1}

From our analysis of the effects of benefit on both quits and lay-offs, the prediction is made that an increase in benefit-provision will raise the number of job-separations. It has already been concluded that benefits will lengthen the mean duration of given spells of unemployment. To conclude our analysis of the effect on overall unemployment, it remains only to consider the effect of benefits on the labour-force participation-rate.

The Effect of Unemployment-Insurance on the Labour-Force Participation Rate

Unless the incidence of the costs of the benefit scheme happen to fall entirely on workers, the institution of an unemployment-insurance plan (or an increase in its scale of benefits) will create a greater inducement for those who are currently year-round non-participants to join the labour-force. Prior to the introduction of benefits, this part of the population rejected participation - the expected weekly return to participation was insufficient to compensate for loss of time to paid employment and job-search. After the scheme has been introduced, the expected weekly return to being in the labour-force is raised from $\frac{eW}{52}$ to $\frac{eW + UB}{52}$ (where $W$ is the weekly wage, $B$ the weekly benefit-payment, $e$ the expected number of weeks of employment, and $u$ the expected number of weeks of non-employment). Some year-round non-participants will now

\textsuperscript{1} The argument focussed on temporary lay-offs and ignored permanent lay-offs. Feldstein (1978, p. 834) justifies this by pointing out that 75% of employees laid-off in U.S. manufacturing return to their old employer.
switch to being counted as year-round participants (even if they are not effective participants for some of the year, they will be counted as participants because benefits depend on declaring oneself available for work).

Mean participation-rates will also vary to the extent that there is a change in the behaviour of part-year participants in the labour-force. Some of these will increase the number of weeks of effective participation in order to secure benefit-qualification. Those who already work a sufficient number of weeks to qualify for benefit may lower their period of effective participation but now have an incentive to declare themselves participants in the non-work part of the year; they will, therefore, increase their period of registered participation. These changes in behaviour and changes in the way in which behaviour is described will both raise recorded participation-rates. The extent of the rise depends on the extent to which it is possible, at low search-cost, to take employment for short periods of the year (say the eight weeks required to qualify for benefits in Canada in the nineteen-seventies). For Canada, Swan (1975) estimated that the change in benefits made by the Unemployment Insurance Act (1971) raised the labour-force participation-rate by two percentage points.

The effect on the unemployment-rate of an increase in labour-force participation depends on whether those joining the labour-force experience

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1. If 'leisure' is a normal good, the effect of an increase in benefit-rates will be to reduce the number of work-weeks.
more or less unemployment than the average worker. In fact, they may be expected to experience more unemployment than average. Those who were year-round non-participants before must have valued leisure relatively highly and can be expected to continue to spend more weeks than average in non-employment (in terms of Figure 1-1, many may move from a corner solution at A to a corner solution at B); those who were part-year participants before and who now-register themselves as in the labour-force for other periods (solely in order to be eligible for benefit) will add extra participant-weeks but all of these weeks will now be counted as unemployment-weeks. Another point is that, since the great majority of adult men are already participants, those joining the labour-force in response to benefit provision will be women and young males - groups with above-average unemployment-rates in all countries (including Canada).

An Assessment of the Overall Impact of Unemployment-Insurance on Unemployment-Rates

The preceding account supports the prediction that the introduction or expansion of an unemployment-benefit system will raise the overall unemployment-rate; this rise will occur because of the impact of benefits on:

(i) the mean duration of a given spell of unemployment;\(^1\);
(ii) the number of spells of unemployment; and
(iii) the labour-force participation-rate.

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1. Although we predict that this will increase, the mean duration of all spells of unemployment could decline depending on the mean duration of those unemployment spells induced by unemployment-insurance provision.
Our presumption in favour of the unemployment-insurance hypothesis of increased unemployment has been supported for several countries and periods; for a representative sample, see the following: for the United States, Bailey (1974), Grubel and Maki (1976) and Komisar (1968); for Great Britain, Maki and Spindler (1975); for a selection of readings on other countries, Grubel and Walker (1978). Canadian studies are reviewed briefly below. A critical view of all such studies — one that emphasises the inconclusive nature of the evidence — is available from Kaliski (1975 and 1976).

Studies of the Effect of Unemployment-Insurance in Canada

The principal empirical analysis in this paper will examine the causes of the increase in unemployment in Canada. It is, therefore, relevant to review briefly those unemployment-insurance studies that have taken the Canadian case as their context.

Grubel, Maki and Sax (1975) assessed the impact of unemployment-insurance changes by estimating (on the basis of annual data for the period 1953-72) an equation in which the dependent variable was the log of the rate of unemployment. There were six independent variables. Two represented unemployment-insurance conditions — these were first the replacement-ratio (the ratio of the weekly rate of unemployment-benefit to the average weekly wage) and, secondly, the proportion of benefit claims ruled ineligible. The other independent variables were

1. For a more detailed survey, see Bockin and Cournoyer (1978).
2. This variable was intended as a proxy for the severity with which the rules of the scheme were being implemented.
the male and female labour-force participation-rates and the nominal G.N.P. for both the current and preceding year. Estimation was by two-stage least-squares and the coefficients on all six variables proved significant at the 5% level. The sign of the estimated coefficient on the replacement-ratio was positive and that on the proportion of claims rejected was negative. The impact of the 1971 Unemployment Insurance Act was quantified by using the mean values of the two insurance variables over 1953-70; these means were fed into the estimated equation together with the true values of each of the other variables. The estimated equation predicts that the rate of unemployment would have been 5.5 per cent (instead of 6.3 per cent) had these insurance variables remained at their 1953-70 means (instead of having been raised by the 1971 legislation). Grubel, Maki and Sax, therefore, attributed to the 1971 Act a 0.8 percentage point rise in Canadian unemployment.

Maki (1975) offered an analysis very similar to that of Grubel, Maki and Sax (1975) except that he applied the model at the regional level. The dependent variable was the same as in his work with Grubel and Sax but the independent variables numbered one fewer because the overall labour-force participation-rate replaced the two sex-specific participation-rate series. The unemployment increases attributed to the 1971 Act varied substantially across the five standard Canadian regions—the increases estimated were as much as 1.6 percentage points in the Atlantic and British Columbia regions but only 0.2 percentage points in Ontario (where the coefficient on the replacement-ratio was not statistically significant).
Green and Cousineau (1976) presented to the Economic Council of Canada a study on the effect of unemployment-insurance. They examined the unemployment-vacancies relationship and found that an econometric estimation of it based on the period up to 1971 underpredicted the 1972 and 1973 relationships; their conclusion was that there had been a structural break in the relationship in 1971, the year of introduction of the new unemployment-insurance system. They also estimated an equation for the absolute number of unemployed persons; they used the two unemployment-insurance variables employed by Grubel, Maki and Sax (1975) and the coefficient on each proved statistically significant. Overall, they estimated that if the unemployment-insurance system in 1972 and 1973 had been the same as that which prevailed in the nineteen-sixties, the unemployment-rate would have been 0.5 to 0.7 percentage-points lower.

Siedule, Skoulas and Newton (1975) adapted the complex CANDIDE model of the Canadian economy to assess the impact of the 1971 Act. By using simulations, they estimated that it added 0.4 percentage-points to the unemployment-rate. However, Bodkin and Cournoyer (1978), surveying their work, suggest that this may be an underestimate.

A consensus drawn on the basis of the Canadian studies cited would point to there having been upward pressure on the unemployment-rate as a result of the introduction of the Unemployment Insurance Act; the impact of the Act was to raise the unemployment-rate by up to a full percentage point. However, none of the writers attributed all of the nineteen-seventies unemployment increase to the benefits explanation. Thus, it seems desirable to search for complementary explanations of the apparent trend rise in unemployment. It also seems desirable that any
search for such complementary explanations should take account of the 
international nature of the unemployment problem. The great strength of 
the benefits explanation is, after all, that an expansion of unemploy-
ment-insurance systems was occurring in several of those countries in which 
unemployment was rising; econometric evidence on the importance of 
unemployment-insurance variables is available for several countries. 
In this paper, we examine another set of variables which moved with some 
synchronisation across countries and ask whether this set of variables 
may also add something to our understanding of the causes of rising 
unemployment. The set of variables chosen for examination consists of 
measures of the age-composition of the population.

4. Changes in the Age-Structure of the Population and Labour Force

Demographic trends provide one set of variables that have shown a 
fair degree of uniformity across Western economies. Thus Chart Three 
graphs birth-rate experience for four countries, Canada, the United 
States, the United Kingdom and West Germany. Wartime data are not 
available for Germany and the pre-War experience for that country was 
different from that of the other three. In the other three countries, 
birth rates were at historically low levels during the nineteen-thirties 
but rose sharply from the early years of the War to reach a peak in 1947. 
During the nineteen-fifties, birth-rate suffered a trough in Britain and 
West Germany, only to rise again towards the end of the decade. In 
Canada and the United States, the baby boom showed no sign of abatement 
after 1947, the birth-rate showing only marginal changes between then 
and 1957. That year was the beginning of the end of the baby boom. In 
the two North American economies, birth-rate declined sharply after 1957
Chart 3: Birth Rate per 1000 persons
and the decline has remained without interruption almost up to the present. It was 1963 before Britain and Germany began a similar experience but in these countries also the decline in birth-rate, once started, was steady and sharp. In all four countries, birth-rates fell to their lowest recorded levels during the nineteen-seventies.

Table 1-2: Birth-Rates in Four Countries
Live births per 1,000 population: annual averages

<table>
<thead>
<tr>
<th>Year</th>
<th>Canada</th>
<th>U.S.A.</th>
<th>U.K.</th>
<th>West Germany</th>
</tr>
</thead>
<tbody>
<tr>
<td>1937-39</td>
<td>20.4</td>
<td>17.3</td>
<td>15.4</td>
<td>not available</td>
</tr>
<tr>
<td>1940-45</td>
<td>23.1</td>
<td>19.8</td>
<td>17.6</td>
<td>not available</td>
</tr>
<tr>
<td>1946-59</td>
<td>27.6</td>
<td>24.5</td>
<td>16.9</td>
<td>16.4*</td>
</tr>
<tr>
<td>1960-69</td>
<td>22.0</td>
<td>20.3</td>
<td>17.8</td>
<td>17.6</td>
</tr>
<tr>
<td>1970-76</td>
<td>16.1</td>
<td>15.8</td>
<td>14.2</td>
<td>11.1</td>
</tr>
</tbody>
</table>


*1948-59.

For our purposes, the number of births is of more interest than the birth-rate. Chart Four, therefore, traces the number of births in one country (Canada) for each year from 1925 into the nineteen-seventies. Long swings of substantial amplitude mark the Canadian experience in this period. Before the Second World War, the number of births was stagnant, being only as high in 1940 as in 1925. After 1940 and until 1959, there was a steady rise in the number of births so that twice as many Canadians were born in 1959 as in a typical year of the nineteen-thirties. In 1960, the number of births began a sharp decline that has continued almost to the present day.
Chart 4: Number of Live Births in Canada, 1925-1970
What has all this to do with unemployment? The point is that variations in the fertility of the population have a lagged effect on the age-structure of the labour-force. Below, it will be argued that the unemployment-rate is partially dependent upon the age-structure of the labour-force.

It is necessary to be more precise about the lagged effect of birth-rate patterns on labour-force composition. The argument here is that the number of labour-force entrants in any year will depend crucially on the number of births in the country approximately twenty years earlier. Of course, the population-cohort born twenty years ago will have been augmented in size by net inward migration and will have lost members through mortality. However, the number of persons reaching working-age from year to year will still vary directly with the number of births twenty years earlier, so long as the influence of mortality and immigration do not vary greatly from one cohort-group to the next. In the absence of swings in labour-force participation-rates, variations in the number of persons reaching working-age will then lead directly to variations in numbers of young people entering the labour-force.

Thus, one will have an 'echo' effect from numbers of births to numbers of young workers. This echo effect is very clear from Chart Five which shows the number of young workers (defined as 20-24 year olds in the labour-force) in Canada for each year from 1953 to 1978. The number was fairly static before 1960 but there then began a sharp and steady rise, numbers more than doubling between 1961 and 1975. These variations in the number of young workers echo faithfully earlier
variations in the number of births; evidently, the influence of migration, mortality and participation-rate changes on successive cohorts was insufficient to outweigh the effects of the substantial variations in birth numbers on subsequent labour-market entry figures.

The argument so far has been couched in terms of the lagged impact of the number of births on the number of young workers. A similar argument explains the impact of the birth-rate on the proportion of young workers in the labour-force. Easterlin (1980, p. 17) points out that the birth-rate in year t is the ratio of those born at that date to the total population in that year. The size of this ratio will be reflected twenty years later in the ratio of those 20 years old in year t+20 to those aged more than 20 in year t+20. The ratios will not be exactly the same but, subject to there being little inter-generational variation in the influence of immigration and mortality, the latter ratio will vary in size directly with the former ratio. Thus, an increase in the birth-rate in year t will yield an increase in the proportion of young people in the adult population two decades later. Again, so long as swings in the participation-rate are not too great, it will be reflected also in an increase in the proportion of the labour-force accounted for by young workers. This 'echo' is very evident in Chart Six and Seven.

(a) Chart Six shows the proportion of the labour-force aged 20-24. This

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1. A similar echo effect for the U.S.A. was pointed out by Johnson (1976)
fell in every year until 1961; echoing the post-1940 baby boom, it increased every year from then until 1977; the rate of increase was fastest from 1963 to 1969.

(b) If one expands the definition of 'young workers' to include all those aged 15-24, the pattern is similar except that the increase in the proportion of the labour-force accounted for by the group was reversed after 1974 (this reversal reflected the fall in birth-rate after 1957 (Chart Seven)).

One again, the influence of mortality, immigration, and participation-rate changes is insufficient to prevent a strong 'echo effect' appearing in the labour-force data.

5. Why Might Changes in Age-Structure Influence the Rate of Unemployment?

The idea to be explored in this paper is that part of the explanation for the upward trend in unemployment-rates may lie in the changes that have occurred in the age-structure of the working-age population. Specifically, it is argued that the increase in the proportion of the labour-force that consists of young persons may have put upward pressure on the rate of unemployment. Such an effect may be expected for the following reasons.¹

(a) The young age groups (15-19, 20-24) have shown consistently above-average unemployment rates. Why this should be so is discussed later in

¹. Only a preliminary outline of reasons is provided; more detailed justification is postponed until later in the paper.
the paper. Whatever the reason, an increase in the proportion of young workers represents an increase in the size of a group with a high incidence of unemployment and, ceteris paribus, this implies an increase in the (overall) rate of unemployment.

(b) Within the young age-groups, an increase in numbers may lead to an increase in the rate of unemployment unless the economy is experiencing boom conditions: the market may have difficulty in absorbing a high rate of influx of (inexperienced) young workers into the labour force (without considerable downward flexibility in wage rates).

(c) An increase in the number of young workers may induce unemployment in other age-groups also: if there is some decline in the relative youth wage, employers may substitute some young workers for members of the middle-aged group. As a result, the unemployment-rate for the latter group would rise unless there were sufficient flexibility in the relevant wage-rates.

(d) On the other hand, to the extent that the labour-force is increased in size by the entry of the baby-boom cohort, this may increase the demand for experienced labour to complement the greater employment of inexperienced labour; this may reduce the recorded unemployment rate for older workers.

(e) Finally, the labour supply decisions of older workers may also be affected. An increase in the number of young workers will raise the number of families where there are several potential wage-earners (this arises because many teenagers live in the parental home and become additional wage-earners on entering the labour force). In such
households, there will then be greater freedom for other household members to 'specialise' (from time to time) in non-market work, leisure or job-search (according to where their comparative advantage lies). This will be another factor generating upward pressure on the measured rate of unemployment.

For all these reasons, an increase in the proportion of young persons may be expected to influence the overall unemployment-rate. This prediction implies that the children of the baby-boom generation may have been expected to suffer from particularly unfavourable conditions when they reached the age for entering the labour-market. Easterlin (1980) makes this prediction and relates it to changes in wage-rates and in the incidence of crime, suicide, and marriage within the age-cohort in question.

Note that the reasons for this influence of demographic change on unemployment are very different from those advanced by the theorists of the nineteen-thirties (Hansen (1939)). Those theorists saw the size of the labour-force relative to the total population as being of key importance: the more people were outside the labour force, the better would be the outlook for a reduction in unemployment since these non-workers would still have a demand for goods but would not compete for jobs. Modern theory now rejects this emphasis on the demand-side linkage between demographics and the macro-economy: it is presumed that monetary and fiscal policy could offset any variations in aggregate demand that resulted. The emphasis is now rather on the
supply side and on those effects that result from the different characteristics of young versus old workers. 1

1. The emphasis on the supply-side in this paper reflects the spirit of most current writing on macro-economics. The shift in emphasis away from demand-side considerations is a relatively recent phenomenon and has presumably occurred as a reaction to the failure of the followers of Keynes to include a serious analysis of supply functions in the building of their models. The importance of this shift in professional thought is discussed by Klein (1978).
Chapter II

The Decomposition of Changes in the Overall Rate of Unemployment


This section sets a framework within which it will be possible to evaluate the hypothesis that changes in age-composition affected the overall rate-of-unemployment in Canada during the nineteen-seventies. As a starting-point, consider the truism that the national unemployment-rate is the weighted sum of the unemployment-rates of the several demographic groups that make up the labour-force; the weights are given by the proportion of the labour-force accounted for by each demographic group. In symbols

\[ U = \sum_{i=1}^{n} w_i u_i, \]

Here, \( U \) is the national unemployment-rate; \( w_i \) is the proportion of the labour-force taken up by members of demographic group \( i \); and \( u_i \) is the unemployment-rate for group \( i \). There are \( n \) demographic groups.

From this way of defining the unemployment-rate, it is evident that a change in the demographic composition of the labour-force may affect the aggregate unemployment-rate in either or both of two ways. The first possibility is that the demographic change will alter the unemployment-rate by changing the weights, \( w_i \). This will be termed here the weights effect of demographic change. The second possibility is that demographic change will affect group-specific unemployment-rates (the \( u_i \)) and thereby influence \( U \). This will be termed here the rates effect of demographic change. The empirical problem of course is
to isolate the weights effect and the rates effect.

For the sake of clarity in exposition, it is worth stressing again that the term rates effect is not defined here as referring to that part of the change in $U$ 'caused' by changes in the $u_1$. Rather, it includes the effects of only those changes in the $u_1$ that can be attributed to the influence of demographic change. Similarly, the term weights effect refers to the influence of demographic change operating via changes in the labour-force weights, $w_1$.

It should be noted that several authors have assessed the impact of demographic change by estimating what we call a weights effect, though they have not generally taken the analysis further to discuss rates effects. Thus, the procedure to be employed in this chapter (or variations of it) have been used by inter alia Perry (1970), Denton et al. (1975), Council of Economic Advisors (1974) and Clogg (1979). Cheshire (1973) followed an analogous procedure when he decomposed differences between regional unemployment-rates in Britain into a part attributable to differences in regional industrial composition and a part attributable to differences in regional industry-specific unemployment-rates.

2. Isolating The Weights-Effect

The change in the aggregate unemployment-rate between two periods, $0$ and $1$, may be divided arithmetically into three parts. The first part of the change is that unambiguously attributable to changes in the
labour-force weights, $w_i$. This may be written as

$$
\sum_{i=0}^{1} w_i (w_i - w_i) (u_i - u_i)
$$

where superscripts refer to the periods 0 and 1.

The second element in the change in $U$ between the two periods is that unambiguously attributable to changes in the group-specific unemployment-rates, $u_i$. This element may be written as

$$
\sum_{i=0}^{1} w_i (u_i - u_i)
$$

There now remains a third element in the change in $U$ between periods 0 and 1. This is given by the expression

$$
\sum_{i=0}^{1} (w_i - w_i) (u_i - u_i)
$$

which is known as the interaction term. This reflects the joint effects of the $w_i$ and $u_i$ varying together and is therefore known sometimes as the covariance term.

We have now decomposed the change in $U$ between two periods into three parts:

$$
U - U = \sum_{i=0}^{1} w_i (w_i - w_i) + \sum_{i=0}^{1} w_i (u_i - u_i) + \sum_{i=0}^{1} (w_i - w_i) (u_i - u_i)
$$

The interpretation of the covariance term (the last part of the identity) requires some care and, specifically, some view on the direction of causation between the variables in the economic model. A simple model that would highlight the issues involved might embody
the following set of relationships:

\[ U = \sum_{i} w_i \]  \hspace{1cm} (i)

\[ w_i = p_i \cdot \frac{PR_i}{PR} \]  \hspace{1cm} (ii)

\[ \frac{PR_i}{PR} = g(u_i, U) \]  \hspace{1cm} (iii)

\[ u_i = h(X, w_i) \]  \hspace{1cm} (iv)

Here, \( X \) stands for a set of 'other exogenous variables' (including, for example, the state of aggregate demand in the economy); \( p_i \) is the proportion of the working-age population accounted for by group \( i \); \( PR_i \) is that group's labour-force participation-rate; \( PR \) is the economy-wide participation-rate; and \( w_i \) and \( u_i \) have the same meaning as before.

In the model, (i) is simply the identity established above. Equation (ii) is an identity that states that the labour-force weights depend on both population-weights and participation-rates. Equation (iii) states that the participation-rate for a group is influenced by its unemployment rate (and that for the whole economy by the overall unemployment-rate): the link would be via the discouraged-worker effect. In equation (iv), the proposition is that a group's rate of unemployment is dependent on its weight in the labour-force and on certain exogenous variables.

For the moment, we will ignore 'other exogenous variables' (\( X \)). With these excluded from the model, population-proportions would then constitute the only exogenous variables. They are truly exogenous because they are determined by past birth, death and
migration activity. There is therefore no chain of causation running towards the population-weights. From the population-weights to the labour-force weights, there are, according to our model, two possible routes via which the former influence the latter (with a consequence for the overall unemployment-rate, U). First, with participation-rates constant, changes in the \( p_i \) would induce proportionate changes in the \( w_i \). Secondly, any such changes in the \( w_i \), according to our hypothesis of the rates-effect (equation (iv)), would have an influence on the \( u_i \); these effects on the \( u_i \) influence the \( P_i \) (equation (iii)) which in turn influence the labour-force weights, \( w_i \) (equation (ii)). Thus population-weights influence labour-force weights both directly and indirectly.

We defined the weights effect as the effect on overall unemployment of changes in population age-composition in so far as these effects occur by altering labour-force weights. On this basis, the model specified above (still ignoring 'other exogenous variables') would suggest that the interaction term should be added to the first term in our identity. This is because changes in \( u_i \) that cause changes in \( w_i \) have as their origin variations in population-proportions. If this procedure is adopted, the "full weights effect" is given by

\[
\begin{align*}
\sum_{i} \epsilon_i (w_i^1 - w_i^0) + \sum_{i} (w_i^1 - w_i^0)(u_i^1 - u_i^0).
\end{align*}
\]

This measure then includes both the direct and indirect weights effects of the variations in the population proportions.
The reasoning in support of this allocation of the interaction term to "weights" depends on our having ignored the "other exogenous variables" included in equation (iv). In fact, changes in these variables may be more important than changes in population-weights in explaining changes in $u_i$. To the extent that this is true, any changes in the $u_i$ that generate changes in the $PR_i$ and $w_i$ have as their principal origin non-demographic changes in the environment. In that case, the primary causation would be from rates to weights and one might then be content to exclude altogether the interaction term from any assessment of the weights effect. The resulting "pure weights effect" measure would be given by

$$\sum_{i} u_i^0 \left( w_i^2 - w_i^0 \right).$$

A third possible way of allocating the interaction term was suggested by Kitagawa (1955). Where one is unable to take other than an arbitrary view on the relative strengths of different directions of causation, Kitagawa recommended dividing the interaction term into two equal parts and adding one part to each of the first two terms in the identity. The "Kitagawa weights effect" would thus be given by

$$\sum_{i} \left( (u_i^0 + u_i^1) \right) \left( w_i^1 - w_i^0 \right).$$
We now have three statistical measures of the weights effect:

\[ (1) \sum_{i} u_i^0 (w_i^1 - w_i^0) \] "pure weights effect"

\[ (2) \sum_{i} \left( \frac{u_i^1 + u_i^0}{2} \right) (w_i^1 - w_i^0) \] "Kitagawa weights effect"

\[ (3) \sum_{i} u_i^0 (w_i^1 - w_i^0) + \sum_{i} (u_i^1 - u_i^0)(w_i^1 - w_i^0) \] "full weights effect"

The choice between these three statistical measures of the weights effect depends on what economic model of the labour-force is adopted; e.g., the strength of the link between unemployment-rates and participation-rates appears crucial. A priori, a choice between the three measures is difficult. However, it may be that our assessment of the weights effect will prove to be insensitive to our choice of measure. To check this possibility, various estimates for weights effects in Canada are computed in the next section.

3. Decomposition of Changes in the Canadian Unemployment-Rate

The procedure for examining the influence of demographic change in Canada involved dividing the labour-force into four groups according to age. The four age-groups were:

15 - 19 years
20 - 24 years
25 - 44 years
45 years and above.

The age-group divisions employed follow those in the Labour Force Survey and are suggested by the extent to which unemployment-rates are different across age-groups. Thus, there is a consistently higher
unemployment-rate amongst 15 - 19 year olds than amongst 20 - 24 year olds. Therefore, these groups were separately defined. However, there is no strong divergence in unemployment-rates between the 25 - 34 and 35 - 44 age-groups. Therefore these groups were not separately defined. A further subdivision of the 15 - 19 age group (into 15 - 16 and 17 - 19, e.g.) may have been obvious in the context of this reasoning but the data on these more finely-defined groups have been collected by Statistics Canada for only a limited number of years.

The period of analysis for Canada in this study is 1953-78; annual data are used. Presentation of results is facilitated by specifying four distinct trade-cycles within the period. These trade-cycles are defined from one peak in the Canadian unemployment-rate to the next peak in that rate; the cycles in question are 1954-58, 1961-68, 1968-72 and 1972-78. For each cycle, the mean unemployment-rates and mean labour-force weights were computed from annual data. The mean overall unemployment-rate for each cycle is given in Table 2-1.

Table 2-2 shows the change in mean unemployment-rate from cycle to cycle [column (1)] and in each case decomposes the change into three parts: the first that unambiguously attributable to between-cycle changes in labour-force weights (\( \Sigma u_i^0(w_i^1-w_i^0) \)); the second that unambiguously attributable to between-cycle changes in age-specific unemployment-rates (\( \Sigma w_i^0(u_i^1-u_i^0) \)); and the third the interaction
Table 2-1

Mean Unemployment-Rates Across Four Canadian Trade-Cycles

<table>
<thead>
<tr>
<th></th>
<th>percent</th>
</tr>
</thead>
<tbody>
<tr>
<td>1954-58</td>
<td>4.22</td>
</tr>
<tr>
<td>1961-68</td>
<td>4.54</td>
</tr>
<tr>
<td>1968-72</td>
<td>5.34</td>
</tr>
<tr>
<td>1972-78</td>
<td>6.77</td>
</tr>
</tbody>
</table>

Table 2-2

Decomposition of Changes in the Unemployment-Rate Across Four Canadian Trade-Cycles

(percentage points)

<table>
<thead>
<tr>
<th>trade-cycles</th>
<th>total change in</th>
<th>$\sum u_i^0(w_{i1}^1-w_{i1}^0)$</th>
<th>$\sum w_i^0(u_{i1}^1-u_{i1}^0)$</th>
<th>$\sum (u_{i1}^1-u_{i1}^0)(w_{i1}^1-w_{i1}^0)$</th>
<th>sum of columns (2),(3),(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
<td>(5)</td>
</tr>
<tr>
<td>1954-8 to</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1961-8</td>
<td>+0.32</td>
<td>+0.02</td>
<td>+0.22</td>
<td>+0.01</td>
<td>+0.25</td>
</tr>
<tr>
<td>1968-72</td>
<td>+1.12</td>
<td>+0.07</td>
<td>+0.95</td>
<td>+0.05</td>
<td>+1.07</td>
</tr>
<tr>
<td>1972-8</td>
<td>+2.55</td>
<td>+0.12</td>
<td>+2.19</td>
<td>+0.17</td>
<td>+2.48</td>
</tr>
<tr>
<td>1961-8 to</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1968-72</td>
<td>+0.80</td>
<td>+0.06</td>
<td>+0.73</td>
<td>+0.03</td>
<td>+0.82</td>
</tr>
<tr>
<td>1972-8</td>
<td>+2.23</td>
<td>+0.11</td>
<td>+1.97</td>
<td>+0.15</td>
<td>+2.23</td>
</tr>
<tr>
<td>1968-72 to</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1972-8</td>
<td>+1.43</td>
<td>+0.09</td>
<td>+1.28</td>
<td>+0.04</td>
<td>+1.41</td>
</tr>
</tbody>
</table>
or covariance term \( \Sigma (w_i^1 - w_i^0)(u_i^1 - u_i^0) \). Column (5) sums the three components of change; this sum is different from that shown in column (1) because of rounding errors in the calculation of the three components of change.

The principal feature of Table 2-2 is that the bulk of variations in the unemployment-rate occur because of variations in age-specific unemployment-rates rather than because of changes in labour-force weights. Our principal task in this paper will therefore be to assess the extent to which demographic change may have influenced age-specific unemployment-rates. In other words, if demographic change is to be accorded great significance it could only be via our rates effect. Most of Chapter IV will be devoted to evaluating this rates effect.

Consider now the weights effect. Table 2-3 gives the percentage of the change in the Canadian unemployment-rate between each pair of trade-cycles that can be attributed to the weights effect. In each case, three figures are given: these correspond to calculations of the pure weights effect, the Kitagawa weights effect and the full weights effect respectively. Whichever measure is used, the quantitative significance of the weights effect appears to be limited. For example, we are most interested in the increase in the Canadian unemployment-rate during the nineteen-seventies. Between the 1968-72 and 1972-8 trade-cycles, the mean unemployment-rate increased by 43 percentage points. Depending on the method of measurement chosen, between 6.2 and 9.3 percent of the change is attributable to the weights effect. Not only is this a small (though not trivial) part of the overall increase in
<table>
<thead>
<tr>
<th>trade cycle</th>
<th>pure weights effect</th>
<th>Kitagawa weights effect</th>
<th>full weights effect</th>
</tr>
</thead>
<tbody>
<tr>
<td>from 1954-8 to</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1961-8</td>
<td>8.8</td>
<td>10.8</td>
<td>12.7</td>
</tr>
<tr>
<td>1968-72</td>
<td>6.7</td>
<td>9.1</td>
<td>11.4</td>
</tr>
<tr>
<td>1972-8</td>
<td>4.8</td>
<td>8.2</td>
<td>11.7</td>
</tr>
</tbody>
</table>

| from 1961-8 to |                     |                        |                   |
| 1968-72 | 6.7 | 8.7 | 10.6 |
| 1972-8 | 4.9 | 8.4 | 11.9 |

| from 1968-72 to |                     |                        |                   |
| 1972-8 | 6.2 | 7.8 | 9.3 |
unemployment, it is also the case that the assessment of the weights effect seems insensitive to the choice of measure. The paper will therefore proceed with the assumption that the pure weights effect is appropriate. Justification for this choice will in fact be given in Chapter III; but as has been demonstrated here, the results would not differ greatly if one of the alternative measures were substituted. In Chapter IV, regression analysis will be employed in the estimation of a rates effect. Our weights effect will then be added to gauge the overall importance of changes in the age-structure of the population of Canada.

1. Additional discussion of the choice is included in Chapter 3.5
Chapter III

The Influence of Demographic Change on The
Rate of Unemployment: Some Theoretical
Considerations

1. Aims of the Chapter

This chapter seeks to review various strands in the literature that all point to the likelihood that changes in the age-structure of the population will induce changes in the overall rate of unemployment. The organisation of the chapter is based on the existence of separate channels through which the changes in age-composition may influence unemployment. In particular, three sets of considerations are discussed, each pointing to a different avenue of influence for demographic change.

First, suppose we take labour-force participation-rates as given for each demographic group. Population-composition changes would then generate proportionate changes in labour-force composition. Even if there were no consequential changes in age-specific unemployment-rates, the overall unemployment-rate would nevertheless be affected so long as different demographic groups have different unemployment-rates. We need to ask here why unemployment-rates should be expected to differ across age-groups, independently of the relative sizes of different age-groups. This, of course, is equivalent to asking why there should be a weights effect associated with demographic change.

Secondly, continuing to take participation-rates as given, we need to ask via what theoretical mechanism could changes in age-structure affect age-specific unemployment-rates. Why should there be what
Wachter and Kim (1975) call for 'short-overcrowding effects'. This, of course, is equivalent to asking why there should be a rates effect associated with demographic change.

Thirdly, we need to review the implications of the fact that it may be inappropriate to regard age-specific participation-rates as given. They may in fact be altered by changes in the age-composition of the population, either directly or via those changes in unemployment-rates that follow from demographic change. These possibilities will be explained and discussed.

2. Age and Unemployment

This section addresses the question of why it may be expected that unemployment-rates differ across age-groups, independently of the relative numbers in each of those age-groups. Two major strands in the literature cast light on the question. First, the theory of job-search is relevant to the extent that rates of return to job-search may differ across age-groups, with different periods of search resulting, according to the age of the economic agent concerned; this body of theory is capable of explaining why the incidence of voluntary unemployment would be expected to vary across age-groups. Secondly, the body of theory on the effects of minimum-wage laws suggests that the adverse consequences of such laws impinge with particular severity on young people; this theory is therefore capable of explaining why the incidence of involuntary unemployment will be expect to vary across age-groups.

Implications of Search Theory

Job-search theory, popularised by the Phelps volume (1970) and pioneered earlier by Stigler (1962), takes as its central assumption the

1. discussed in Chapter I above (pp.12-14)
idea that time spent in full-time job-search activity is time productively spent: It is productive time in the sense that the process of search is likely to yield information about the labour market that will enable an agent's lifetime (money and psychic) income to be increased. Searching for a better job-offer than those previously located therefore has a positive expected pay-off. However, it also has a cost. The cost is an opportunity cost: the cost of a week's full-time search consists of the earnings that could have been generated by spending the week at work instead. In considering whether or not to prolong a given period of search, the maximising agent will have to compare the cost of further search with the expected benefit. Similarly, an agent's decision whether or not to quit his current job to engage in full-time search for a better one will involve a balancing of the expected costs and benefits.

Within the framework set by job-search theory, any explanation of differences in unemployment experience between age-groups would depend on there being different expected costs and benefits from full-time job-search, depending on the age of the agent. If the young have a systematically greater incentive to engage in full-time search, then members of this age-group will, on average, spend more weeks in the year in search than will members of older age-groups. Ceteris paribus, higher unemployment-rates will then be recorded for the youth population although this 'search unemployment' is voluntary, it will nevertheless be counted as unemployment in the official figures because the official definition of unemployment, in Canada and elsewhere, embraces all those
who don't have a job but who are actively looking for one.

It seems in fact that conditions indeed favour a higher investment in job search on the part of younger workers as compared with those in the mid-career stage of their labour-force participation. On the one hand, wage rates are lower for young workers (this is to be expected because, in general, productivity will be less for those with little job experience): less income is therefore foregone by remaining without a job. On the other hand, potential benefits of prolonged search are likely to be greater for the young: it is more important to make an optimal choice if it is going to affect one's career progress for forty years rather than for twenty years. Again, just because of their inexperience, the young typically have more options to examine - they will have little first-hand knowledge of on-the-job conditions and, of course, will not have been tied into particular job-types by past training.

These last two factors are presented in greater detail in a review article by Lippman and McCall (1976). They propose that the optimal solution for the job-search has the form

\[
\text{continue searching if } X \leq \epsilon_i \\
\text{accept employment if } X > \epsilon_i, \quad i = 1, 2, \ldots, k.
\]

where \(X\) is the job-offer and \(\epsilon_i\) the reservation wage for the \(i\)th career (of the \(k\) for which entry may be under consideration). Now, for the young, inexperienced worker, knowledge of conditions in the several employments is likely to be highly imperfect and this is likely to be reflected in a bunching of the \(\epsilon_i\)s. Once a particular job-type has been tried, more information will have been acquired and the \(\epsilon_i\)s may therefore be expected to be revised in value, resulting possibly in a reversal of
the decision to enter the particular field for which experimentation has been made. "Experimentation and sorting would be anticipated early in the career-choice process when specific investments are small" [Lippman and McCall, (1976, p. 352)]. The result would be a high job-turnover rate amongst the young with the possibility of a high unemployment-rate coming from an above-average frequency rather than an above-average duration of unemployment.

Empirical support for this possibility is available for both the United States and Canada, and also for Britain. In Britain, Nickell (1980), in a study based on the 1972 General Household Survey, found that the under-25s were more than twice as likely to enter unemployment in any given week as were members of any over-25 age-group; at the same time, their mean duration of unemployment was the lowest of any of the seven age-groups studied. Similarly, in the U.S., Clark and Summers (1978) reported that 70% of teenage unemployment-spells ended in one month or less. For Canada, David Gower (1975) conducted a study of 'Job-Search Patterns'. In Gower's sample, the proportion of under 25's looking for a job was twice as high as the proportion of job-seekers in the labour-force generally. Nevertheless, the average duration of search was relatively short: young job-searchers of both sexes reported above-average success rates in both the receiving of job-offers and the procurement of full-time permanent jobs. Thus, in all three countries cited, high youth unemployment was the result of a high frequency of (relatively short-duration) unemployment. The implication drawn by Gower is that "the main problem for young workers is not to find a job but rather to find the 'right' job and to find jobs that offer
stable employment" [Gower (1975), p. 31]. In other words, young members of the labour-force experience high unemployment-rates partly because many of them are likely to engage in an extensive job-search, one that involves sampling a variety of different jobs over a relatively short period of time. Given their inexperience, and consequential ignorance about the characteristics of different jobs, this can be viewed as rational information-gathering behaviour.

Returning now to the point about the effects of training on the propensity to quit, the prediction may be made that the propensity to withdraw from a particular job (or job-type) will decline with length of experience on that job. This prediction depends on an assumption that the current wage received will include a positive element that will be greater the longer the individual worker has been engaged in that job: specific human capital, in other words, is assumed to be accumulated over time. If this is so, then while the opportunities available in other fields may be the same for all new entrants at any particular time, it will nevertheless take greater prospects elsewhere to cause an experienced worker to switch careers than it will to cause an inexperienced worker to switch careers: the latter will not yet have acquired the pecuniary rewards of having skills learned on-the-job and will therefore have less to lose by a move.

One doubt may be raised at this point. An alternative explanation for the above-average frequency of unemployment amongst the young would be that youth is simply fickle: the high separation-rate of this group.
might be claimed to be based not on any 'rational' behaviour pattern but rather on instability of personality or lifestyle. This possibility is examined by Leighton and Mincer (1974) in a paper that was able to draw upon large-scale samples of the United States labour-force. They established, first, that high youth unemployment in that country (like that in Canada) is generated by a high frequency of job-changing that is in turn associated with relatively recent entry into the labour-force: the age-profile of unemployment reflects the age-profile of job-changing. They then examined a sub-sample of workers of all ages whose distinguishing characteristic was recent migration to the United States; they also looked at a sub-sample of workers who had migrated within the United States. Both these types of migrant worker, workers who might reasonably be expected to be highly-oriented towards improving their income prospects, were also found to have abnormally high incidences of unemployment. Indeed, for the earliest period after migration, the incidence of unemployment amongst immigrants was greater than that for native youths; for internal migrants, incidence was comparable with that for young non-migrants. There seemed then to be some evidence that new entrants to a particular labour-market engaged in relatively frequent job-switching regardless of age: the young just happened to be composed of workers who had short tenures at work and this lead to a high separation rate as experiments in job-matching were undertaken by these workers (and by employers) 1.

1. Oi (1979) draws attention to empirical evidence that the young have a greater-than-average willingness to engage in risk. This may be an additional factor in explaining the willingness of young people to quit-to-search; presumably, those willing to migrate also have a degree of risk-preference above the average for society as a whole.
Marital Status and Job-Search Behaviour

Another factor, one that is related to age, that will influence the amount of search undertaken is marital status. It is of course the case that that part of the labour force aged below 25 years includes an exceptional proportion of single persons; and it is known from the Labour Force Survey that, independent of age group, single persons have a higher unemployment-rate than married persons. It follows that, even with all other things equal, the 'young' labour force will experience a higher-than-average incidence of unemployment.

Why should unemployment be higher for single persons? To some extent, the explanation may lie in discriminatory labour practices on the part of employers. However, it is not necessary to assume the existence of such practices: one can formulate a theory of supply-side behaviour that is adequate to explain the phenomenon in question. This theory runs in terms of a single person being more mobile than one with more family responsibilities: such a person may be expected to invest in a more extensive job-search. The details of the analysis are suggested by Jacob Mincer's theoretical and empirical paper on 'Family Migration Decisions' [Mincer (1978)].

Consider the case in which the taking-up of a new job requires a geographical move on the part of the worker. If the worker is a one-person household, with no need to consider any effect on other parties, the decision process is simple. He (or she) will move if and only if

$$G_1 = R_1 - C_1 > 0.$$
Here $G_1$ is the net real income gain from migration; $R_i$ are the gross returns; and $C_i$ are the costs of the move. The worker is assumed to maximise the present value of real income; the calculation of $R_i$ and $C_i$ will of course include a discounting procedure applied with an appropriate rate of discount.

Within this framework, then, the worker will move if $G_1 > 0$. However, for those in family households, this will not necessarily be the case. If decisions in a two-person household are based on the maximisation of joint returns (a plausible situation within a marriage partnership), then the move will take place if and only if

$$G_f = R_f - C_f > 0.$$  

Here, the subscripts indicate that the reference is now to family gains and costs: $R_f = R_1 + R_2$ and $C_f = C_1 + C_2$, where 1 and 2 are the individuals involved.

Now, if the costs of a move are the same for a single person as for a family, then the possibility of the existence of another family member having an effect on the decision is dependent upon $R_1$ and $R_2$ having opposite signs. If $G_1$ is greater than zero, a single person would move but if $(G_1 + G_2) < 0$, then a married person would not move even if $G_1 > 0$. This situation could arise within marriage if the spouse were working and if his or her earnings were likely to suffer from a change of residential location. Of all moves for which $G_1 > 0$, a proportion will be ruled out because of the negative effect on other family members.
Persons who would move if they were single but who do not because of the effect on other family members are termed 'tied stayers'. The phenomenon of the tied stayer will reduce geographical mobility amongst married workers compared with single workers. Of course, there is also an opposite phenomenon, that of the 'tied mover'. A tied mover would not migrate if he or she were single but in fact moves because, although \( G_1 < 0, \) \( G_f > 0 \). The existence of the tied mover will tend to raise geographical mobility amongst married workers compared with single workers.

The degree of mobility of members of family households compared with members of single households will depend upon the relative numbers of tied stayers and tied movers. Mincer presents illustrative figures which indicate that tied stayers are likely to be more numerous. A sufficient set of conditions to guarantee this result is as follows:

(i) \( G_1 \) and \( G_2 \) are drawn from distributions of \( G_1 \) that are normal and identical to each other;

(ii) \( G_1 \) and \( G_2 \) are paired randomly, and

(iii) the probability that \( G_1 > 0 \) is less than 0.5.

If all three conditions hold, marriage will have an inhibiting effect on mobility. This inhibiting effect will be reduced to the extent that either

(a) \( G_1 \) and \( G_2 \) have different means across their separate populations; or

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1. Since far less than 50% of single persons migrate in any year, condition (iii) may be accepted as reasonable.
(b) $G_1$ and $G_2$ are positively correlated with each other.

In extreme cases of (a) or (b), marriage would have no effect at all on the propensity to migrate. Thus, to take case (a) if only one family-member works, $G_1$ will always equal $G_2 (G_2 = 0)$ and the mobility of person 1 will be unaffected by the presence of person 2. \(^1\) Alternatively to take case (b), if $G_1$ and $G_2$ are perfectly and positively correlated, there will never be any conflict between the individual interests of different family members and again the mobility of each person will be unchanged by marriage. If (a) and (b) hold in less than those extreme forms, then the tendency of marriage to inhibit migration will be reduced; some inhibiting effect will nevertheless continue to exist.

For the reasons given, one would expect to find that job changes necessitating geographical mobility would occur with less frequency for members of family households than for members of single-person households. Note however that not all unmarried persons will be members of 'single person households': some will be 'tied' by the presence of parents or others rather than by the presence of a spouse.

Decisions to engage in search-unemployment depend on an evaluation of the probability of finding a job-offer that will be judged better than that currently available. For members of family households, the probability of a given job-offer from another area being judged 'better' may be expected to be less (as explained above). One would therefore expect married persons to engage in a geographically less extensive job-search than that undertaken, on average, by unattached single persons. The result would be a lower incidence of search-unemployment.

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1. Mobility overall may still be lowered if person 2 worked before marriage because, whereas person 2 used to have a positive probability of moving for work reasons, there is now no such incentive for her (or him) to move.
amongst married persons.

The argument may be extended to cover cases where the marriage partner is not a labour-force participant. For there to be a difference in the decision whether or not to take a job in another area, there must now be a difference between $C_i$ and $C_f$. Such a difference may arise if there are psychic costs to the spouse's move (as there may be, for example, if the original location is close to that of parents). Costs are more likely to diverge if children are present since the disruption to their schooling is known to be a powerful disincentive to geographical mobility [Long (1972) and Long (1975)].

**Implications of Minimum Wage Theory**

It is a well-known proposition of elementary economics that, in a competitive labour market, a minimum wage set above the market-clearing level will cause unemployment. This unemployment will not be evenly spread amongst different age-groups if the demand-price for labour is lower for some groups than for others. In other words, if the employer's demand-price is, on average, relatively low for youth labour, proportionately more young workers than older workers will suffer involuntary unemployment as a consequence of the existence of minimum wages.¹

Because many work-skills are learned on-the-job, workers who have been in the labour-force for only a short time are likely to be generally less productive than other workers. To the extent that this is true, an employer, faced with a choice between a middle-aged worker and a

¹. Minimum wage rates are set by legislation in Canada and the U.S.A. but may also be enforced informally by unions with monopoly power.
new entrant to the labour-force, is likely to hire the former unless
the new entrant's price is sufficiently low to offset his productivity
handicap.

A second reason for expecting employers to be willing to pay less
for young workers is that to hire this class of labour is to incur a
risk greater than that associated with taking on other classes of
worker. An employer can never be certain as to how a potential
employee will perform on-the-job; his information on the employee
will be limited. However, it will be more limited in the case of a young
worker because the worker will have little job experience on the
basis of which his prospects may be evaluated. To hire a young
worker, therefore, an employer must run a greater risk of making
a decision that will be regretted ex post (or else he must incur
higher hiring costs - for administration of aptitude and personality
tests, for example).

Finally on the demand side, one may mention the costs of on-the-
job training. To the extent that firm-specific on-the-job training
costs are borne by firms, an employer always runs the risk that his
'investment' will be wasted: he may 'pay' the training costs of a
worker only to find that the worker leaves his position before 'paying
back' the employer's costs. Since young workers are perceived as
being more mobile than those in the prime-age groups, the risk of a
wasted training-cost investment is greater in the case of young workers.
This is another reason for the young finding it relatively hard to
acquire satisfactory employment, at least in the absence of a wage-
structure that allows sufficiently low relative wages for the young.

The training-cost issue arises again in the context of the relative risks of lay-off faced by young and prime-age workers respectively. It is not just that young workers are worst affected by union-imposed 'Last In, First Out' rules. Even without such rules, an employer faced with a deterioration in market conditions, may be most likely to dispense with the services of young, inexperienced workers because they have little training 'invested' in them. The same line of reasoning may encourage him to retain his more experienced workers to avoid both the loss of investment and the incurring of recruitment costs when conditions improve and when such workers may be in short supply. Moreover, a worker with skills acquired on-the-job may be able to take over the simpler work of a new entrant whereas substitution in the reverse direction is likely to be impossible. All these considerations put young workers at most risk whenever lay-offs are contemplated. Oi's (1962) suggestion to this effect received strong empirical support from a study by Parsons (1972). Parsons also examined the extent to which experience on-the-job raises the cost of quitting and found that this too lowered the separation-rate for older workers.

This section has pointed out that a young worker is less likely to be hired than an older worker if their prices are the same. The point about price (i.e. wage) is that the handicaps experienced by young workers should not in themselves prevent their being employed. Provided that they are willing to work for less than established members of the labour-force, their wage-rates should be bid down until it becomes
worthwhile to absorb them all into the ranks of job-holders. However, in Canada, as in several other countries, there exist institutional impediments to the extent to which wages can be bid down; if law or convention stipulates a minimum wage that is 'too high', some members of the labour-force will not be absorbed into employed ranks: a worker whose marginal revenue product is less than the minimum wage will remain jobless. Since workers with very low productivity are concentrated amongst the young and inexperienced, the unemployment burden associated with a minimum wage will fall disproportionately on the young.

The Minimum Wage with Partial Coverage Across Industries

A formal analysis of the effects of minimum-wage laws which have only partial coverage was provided by Mincer (1976). This analysis considers the market for one type of labour. The employed part of the labour-force is divided between an uncovered sector and a covered sector. In the uncovered sector, wages are set competitively and there is thus no unemployment in this sector; in the covered sector, wages are set legislatively, at above the initial market-clearing level. In the simplest case, the model is one-period, workers are risk-neutral, and the work in each sector is identical.

In the context specified, the equilibrium condition for wages
in the covered sector is given by the (equivalent) formulae

\[ pW_m = W_n \quad \text{or} \quad p = \frac{W_n}{W_m} \quad \text{or} \quad \frac{1 - p}{p} = W \]

Where \( W_m \) and \( W_n \) are the post-legislation wage-rates in the covered and uncovered sectors respectively, \( p \) is the probability of gaining covered-sector employment; and \( \bar{W} = (W_m - W_n) / W_n \). The condition defines equilibrium because only with expected wage-rates equal in the two sectors will there be no incentive for further migration between the two sectors.

In the covered sector, let employment be denoted by \( E_m \). Abstracting from any growth in the demand for labour, the number of vacancies in the formal sector in each period will be equal to the product of the separation- (or turnover) rate, \( \delta \), and the covered-sector employment-level, \( E_m \). These \( \delta E_m \) vacancies are filled as soon as they appear, and a remaining pool of unemployed searchers is observed. The probability of employment in a period is given by

\[ p = \frac{\text{number of vacancies}}{\text{number of searchers}} = \frac{\delta E_m}{U + \delta E_m} \]
We now define the covered sector unemployment-rate as

$$u_m = \frac{U}{E_m + U}$$

Using the equilibrium condition $W = \frac{1 - P}{p}$, and the result, noted above, that $P = \frac{\delta E_m}{U + \delta E_m}$, the equilibrium rate of unemployment in the covered sector can be derived as

$$u_m^* = \frac{\delta W}{1 + \delta W}$$

Defining the coverage-ratio of minimum-wages as $k = \frac{E_m}{E_m + E_n}$, and recalling that the market for labour in the uncovered sector is assumed always to clear, the corresponding (equilibrium) aggregate unemployment rate is

$$u_A^* = \frac{k \delta W}{1 + k \delta W}$$

Mincer concluded from this model that the unemployment-rate induced by the imposition of minimum wages was proportional to the percentage wage-gap between the sectors ($W$), the separation rate ($\delta$) and the coverage ratio ($k$). The importance of the rate of labour turnover ($\delta$) arises because a high value of the variable raises the probability of finding a job in the high wage-sector and
therefore encourages greater search-unemployment.

Empirical studies of the effects of minimum wages (e.g., Mincer (1976), Ragan (1977) and Swidinsky (1980)) have typically used a minimum wage variable based on both the minimum-wage rate relative to the average wage and the coverage-ratio of minimum wage laws. In summarising minimum-wage conditions by such an index, account has been taken (at least loosely) of the importance of both \( W \) and \( k \) in the Mincer model. The relationship between the model and the minimum wage index used by these authors will be further discussed in Chapter 4.1 below.

The Effect of Minimum Wages: Empirical Evidence

Our expectation, that the unemployment burden associated with a minimum wage will fall disproportionately on the young, has been supported by several studies. For Canada, one may mention a study of the impact of wage-minima on youth unemployment carried out by Swidinsky (1980). Swidinsky examined the impact of the minimum wage in each of the five standard economic regions of the country. He found that the tendency for the minimum wage to exacerbate teenage unemployment was greater for females than for males and significant only in Ontario, Quebec and the Atlantic Provinces. His estimate for the national economy was that increases in teenage coverage and minimum wages between 1956 and 1975 added 1.2 percentage points to the aggregate teenage unemployment.

Most empirical work on minimum wage laws has however been carried out for the United States. An important contribution was that by T.G. Moore (1971) who carried out an empirical study of the impact of variations in
minimum wages on youth unemployment in the United States. Moore found that, at least for those under twenty years of age, variations in the ratio of the minimum wage to the average wage (as well as variations in the extent of coverage of minimum wage laws) had a statistically significant and unfavourable impact on unemployment-rates. The impact was most severe on non-whites and on teenage girls. However, the impact ceased to be statistically significant for those in the 20-24 age group; presumably, the productivity of the individual workers in their early twenties had commonly risen to an extent such that it was outside the range affected by the legislation.

Another study, one that also used multiple-regression analysis, was that by Douglas Aidie (1973). Like Moore, Aidie found that minimum-wage legislation in the United States had had an adverse effect on teenage unemployment-rates; however, the extent of the impact was different in severity from that reported by Moore. Moore's study suggested that a rise in the U.S. minimum wage from $1.60 to $1.80 in January, 1970, and then from $1.80 to $2.00 in January, 1971, would have raised the unemployment-rate of non-white teenagers (the group apparently most affected by the legislation) by 15.6 percentage points. On the basis of Aidie's reported elasticities, Goldfarb (1974) calculated that a similar pair of minimum-wage increases would have caused only an 8.1 percentage-point rise in non-white teenage unemployment. The two studies were therefore felt by Goldfarb to show a variation in the results that rendered their use unreliable. However, it remains true that each showed statistically significant results qualitatively similar to those predicted by economic theory.
A third study of U.S. data, Kosters and Welch (1972), considered a slightly different question: the effect of minimum wage legislation on the sensitivity of teenage unemployment to variations in total economy-wide employment levels. After establishing that, on average, "teenagers absorb a phenomenally disproportionate share of aggregate employment fluctuations" (p. 108), they went on to show that the existence of minimum wage laws served to exacerbate the situation. In their analysis of data for 1954 to 1968, they defined the "effective minimum wage" as the minimum legal wage divided by the average hourly earnings in manufacturing multiplied by the proportion of workers covered by the legislation. A one per cent increase in the effective minimum wage was then estimated to reduce the share in normal employment of white teenage males by 0.33 percent and to increase their share of "transitional" employment by 2.5 percent. In short, more generous minimum wage provisions not only reduced the proportion of jobs typically taken by teenagers; they also made teenage unemployment more volatile by increasing its responsiveness to variations in the level of aggregate employment.

Unfortunately, the literature does not show complete unanimity amongst researchers with respect to these allegedly adverse effects of minimum-wage legislation. Katz (1970), again using multiple-regression techniques, reported little evidence for the proposition that the law in the United States had raised the rate of unemployment amongst teenagers. Lovell (1972) took Katz's results and attempted to find the reason for the underlying difference between his conclusions and those of Moore. He isolated the crucial difference between them as being
that Moore failed to include a controlling variable that would represent the rate of teenage population growth. This led Lovell to side with Kaitz: he thought that the effects of the teenage population bulge had to be taken into account if a pure measure of minimum wage effects were to be obtained. Lovell therefore rejected Moore's findings.

Lovell's analysis is itself subject to question, however [Aidie and Galloway (1973); Fisher (1973)]. According to economic theory, an increase in the teenage population should not of itself cause an unemployment "problem" to develop. A rightward shift in the supply curve of teenage labour should lead to the market for teenage labour being cleared at a lower real wage; it is only if institutional impediments prevent the wage from being lowered that there then arises a pool of involuntary unemployment. The minimum-wage law is such an institutional impediment: acting in conjunction with population change, it would be predicted to be at the root of the unemployment increase amongst the young. ¹

3. The Influence of Demographic Change on Age-Specific Unemployment-Rates

This section considers in more detail the proposition that an increase in the size of one age-group will raise the unemployment-rate of that group because of 'cohort crowding' (Wachter and Kim (1979)). The proposition rests of course on the idea that workers of different ages

¹ A survey on the effects of minimum wages has been provided by Goldfarb (1974). The link between minimum wages and population change is stressed in Forrest (1982).
constitute in some sense different factors of production. This seems reasonable because in many occupations, distinctly different tasks are performed by 'young' and 'old' workers. In such occupations, 'young' workers could be substituted for 'old' workers only with increasing difficulty; as they were substituted for 'old' workers, they would have to be assigned tasks less and less appropriate to their level of experience. But if workers of different ages do indeed constitute different factors of production, then what determines the degree of elasticity of substitution between them?

Some light may be cast here by considering Welch's career-phase model [Welch (1979)]. Welch viewed a work career as consisting of a series of distinct phases. At each phase, different sets of tasks are performed, depending on how much work-experience has been acquired. At each phase, the activities undertaken are productive, and marginal productivities may be altered by changes in the number of persons engaged in the particular activities of that phase. On this view, an increase in the number of persons entering the first (learner) phase of a career will depress the equilibrium wage for that career-phase. What happens is that an increased supply of a variable factor (youth labour) is added to quantities of other factors (including experienced labour) that are fixed in the short-run. The 'law' of diminishing returns predicts in these circumstances that the marginal productivity of the variable factor will fall and, thus, its relative-wage will have to fall if unemployment is not to result. The required drop in the equilibrium wage for the group whose size has varied will be greater the less the elasticity of substitution between workers in different career-
phases.

Welch points out that at any time a worker will be in transition from one career-phase to the next. There is some possibility therefore of using a worker to perform tasks commonly understood as belonging either to previous career-phases or to the next one. In the latter case, the worker’s progress could be said to be escalated. Now, consider the effect on the market-clearing wage of an increase in age-group size. The effect would be greater for the youngest age-group than it would be if a similar change had occurred for a middle-age-group. This is because the latter group can turn to tasks that would previously have been normal in either earlier or later career-phases. It is harder to substitute young workers into alternative tasks in response to a falling wage-rate since the number of career-phases in which they could perform the required activity is small. Similarly, an increase in the size of the most experienced worker-group would have relatively large effects since those workers are substitutable only backwards into lower-phase positions.

Our prediction, then, is that an increase in the size of an age-group will require a fall in its market-clearing wage. The fall will be greater if the group concerned is a young age-group rather than a middle-age group. To the extent that minimum-wage laws and institutional impediments prevent the required fall in wages from taking place, unemployment will increase as a result of increased group size. The young will again be most prone to suffer unemployment from this source.
because

1) the required fall in equilibrium wage would be greater for the young; and

ii) their mean wages are the lowest of any group—thus their wages are more likely to be in the range where minimum-wage regulations are relevant.

We shall therefore have as two of our maintained hypotheses the propositions that age-specific unemployment-rates are related to age-group size, and that the youngest age-group will be particularly susceptible to this problem of cohort-crowding. In addition, because of their limited substitutability for other groups, the unemployment-rate of the oldest group may also be expected to be sensitive to cohort-size.

A final hypothesis is that female unemployment-rates will be less sensitive to age-group size than will male unemployment-rates. Women tend to be concentrated into a relatively small number of job-categories. These job types often are not of the sort which have well-established career-patterns—the work is frequently unskilled and therefore there would be more substitutability between older and younger workers than there would be in the typical male occupation. As a result, the effects of an influx of new potential workers would be more diffuse, falling not only on the age-group that is expanding in size but also on other female age-groups, members of which are operating in the same labour-market. So within the youngest age-group, we expect the effects of cohort-crowding to impinge more severely on males than on females.
The predictions made above depend upon two assumptions. The first is that the technology is sufficiently inflexible that the changes in wages required to absorb increased supplies of one factor are non-trivial. The second is that wages are insufficiently flexible to vary by enough to maintain full-employment amongst all groups. Whether these assumptions are appropriate will of course depend on whether our maintained hypotheses receive adequate support from the empirical evidence.

**Short-Run versus Long-Run Effects of Demographic Change**

The preceding account was couched in terms of the 'law' of diminishing returns, a theory that considers the effect of varying the supply of one factor of production whilst holding quantities of other factors fixed. In these circumstances, an increased supply of the one variable factor will reduce its marginal productivity and require a lower wage to maintain it in full-employment. This though is only a short-run analysis. In the longer-run, other factors (such as capital) may be varied in quantity; technology may be adapted to suit the characteristics possessed by the factor whose supply has increased; and so, in the longer-run, the fall in wage required to maintain full-employment will be less than in the short-run. In our context, this would imply that any youth-unemployment resulting from wage inflexibility at the time of the increase in the supply of youth labour should become less severe as firms adjust to new labour market conditions. Indeed, to the extent that the ageing of the baby boom cohorts was predictable, some firms should have adjusted their technology in anticipation of enhanced
supplies of unskilled labour and this again should have reduced the impact on youth unemployment of an increased supply of youth labour. Wachter (1976) expressed scepticism over these grounds for optimism about the impact of demographic change. Following his ideas, one might speculate that very substantial changes in relative wages would have been necessary to prevent the increase in the youth population from raising youth unemployment-rates. The short-run problem arises because the organisation of a productive process takes time to adjust if the unskilled wage-rate falls relative to that of other workers and relative to the cost of capital, firms will not at once employ more unskilled workers unless they have anticipated the changes in question and have reorganised the methods of production to take account of them. In fact, the increase in the unskilled labour-force was greater than expected during the nineteen-sixties and nineteen-seventies because of the sharp rise in female participation-rates. Even with appropriate wage flexibility therefore, it would have taken time to absorb the new supplies of young labour. Moreover, even in the medium-term, and even assuming that correct recognition of long-run trends had by then occurred, large firms would not necessarily have found it profitable to restructure their organisation to take advantage of this newly-enhanced source of low-priced labour. In particular, a (costly) investment that would raise the proportion of entry-level jobs may be rendered inappropriate by the expectation that the number of new entrants to the labour-force will be declining in the nineteen-eighties. In all these circumstances, it seems unlikely
that industry would have been restructured sufficiently to absorb all the new entrants of the nineteen-seventies.

A Contrary View

Wachter's ideas have been challenged implicitly in a paper on teenage unemployment in the United States by Feldstein and Ellwood (1979). Feldstein and Ellwood point out that the U.S. labour-market is able to more than double teenage employment between May and July every year. They then argue as follows:

"Employers clearly anticipate a seasonal increase in the supply of teenagers and organise production to take advantage of their availability. We are struck by the contrast between this experience and the claim that much of the current high teenage unemployment-rate is due to the demographic shift that increased teenagers from 7 per cent of the labour force in 1958 to 10 per cent today. If production can adjust so rapidly to the seasonal shift in the demographic composition of the labour force, it would be surprising if it could not adjust to the much slower change in demography over the past two decades. This leads us to believe that too much weight has generically been given to the demographic explanation of the rising teenage unemployment-rate". 

[Feldstein and Ellwood (1979) p. 9]

The point is an interesting one but seems to draw too close an analogy between absorbing seasonal and absorbing secular rises in the number of teenagers available for work. On the labour demand side, certain jobs become available for teenagers in May and June but these are usually explicitly seasonal jobs (e.g., lifeguards, holiday-replacements) that of their nature could not be generated on a permanent basis; indeed school vacations are in part organised explicitly to take account of these temporary opportunities. The creation of permanent jobs is a different kind of problem since there it is entirely new jobs that are needed. On the labour supply side, one may speculate
that summer job-seekers would be less inclined to remain in search-unemployment than those looking for a long-run opportunity and would therefore be more easily absorbed by the labour market.

4. The Influence of Demographic Change on Participation-Rates

In this chapter's analysis so far, the effect of changes in population age-composition has been discussed in the context of an assumption that labour-force participation-rates remain unaffected by demographic influences. This assumption may be challenged on various grounds, discussed below. If it were successfully challenged, then the analysis would have to be modified to the extent that changes in the age-composition of the working-age population could no longer be viewed as inducing proportionate changes in the age-composition of the labour-force.

How might participation-rates be affected by a change in population age-structure? To answer this question by way of illustration, suppose the change consists of an increase in the number of teenagers. With existing age-specific participation-rates held constant, there would be an increase in the proportion of the labour-force accounted for by teenagers and, according to our cohort-crowding hypothesis (discussed in Section 3 above), there would be a consequent increase in teenage unemployment-rates. This increase in unemployment would be predicted to affect participation-rates by the well-known analysis of the discouraged worker effect.

Discouraged workers are those who leave the labour-force when unemployment-rates rise. In the context of search models, the
explanation of their behaviour would be that an increase in the age-group unemployment-rate would be taken as signalling an increase (for members of that age-group) in the mean duration of search necessary to locate a job; this increase in expected search costs would cause some workers to decide that it was not worthwhile to engage in search and these workers would therefore not become members of the labour-force. Such behaviour would tend to lower age-specific participation-rates.

Empirical evidence indicates that discouraged worker effects may be significant for some groups. Duncan (1965) focussed on young people and found that increased youth unemployment tended to raise school-attendance rates\(^1\) in the United States. Perry (1977) found, for most demographic groups, a negative relationship between participation-rates and a measure of unemployment conditions in the U.S. economy (his period of analysis was 1949–75). For Great Britain, Corry and Roberts (1974) examined male and female participation-rates across different regions. The unemployment-rate was not significant in explaining variations in male participation-rates; however, for females, a significant negative relationship between participation- and unemployment-rates indicated the presence of a discouraged worker effect. Corry and Roberts (1974) expected \textit{a priori} a discouraged worker effect for young males also but were unable to test for this because age-specific unemployment-rates are not collected regularly in Britain.

Given the evidence that discouraged worker effects are important,

\(^{1}\text{Full-time school attendance is of course an alternative to full-time labour-force participation.}\)
it is desirable to incorporate them in our model; and indeed, this was done in the simple model outlined in Chapter 2.2. There, a change in population (age-group) proportions affected labour-force proportions \((w_i)\), which then affected unemployment-rates \((u_i)\) (the cohort-crowding hypothesis); a change in the \(u_i\) influenced participation-rates \((PR_1)\), which further modified the \(w_i\).

Our model, then, should incorporate these sorts of discouraged-worker effects; but the relevant equations in our formal model need also to be specified in such a way as to take account of other important determinants of the \(PR_1\). We shall include, as determinants, male and female hourly earnings and the relevant age-group's fertility-rates. These variables are suggested as crucial by a literature that began with Mincer (1962) and Becker (1965), and that views the family as a 'firm' engaged in the 'home production' of utility-yielding commodities; the 'production' of these commodities involves the use of both the time of family members and the purchase of market-provided goods and services. The question therefore arises of how household members should divide their time between household production on one hand and the labour-market on the other - with one possibility being that some household members should not participate in the labour-market at all.

Whether an individual will be a member of the labour-force in a particular period will depend on a comparison between her expected market wage and the shadow price of her time, evaluated at zero hours.

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1. We focus initially on determinants of female participation-rates because it is these that have varied significantly over time.
of market work. The shadow price will reflect the value of her time in the alternative of household production.

If female earnings increase, this will raise the proportion of women for whom the expected market wage exceeds the shadow wage. Thus, we predict that the participation-rate of a female group will depend on female earnings possibilities.

The participation-decision of females will depend also on male earnings possibilities: in the Mincer-Becker framework, the family is the decision-taking unit and male and female labour-supply decisions are made jointly. A common assumption is that husbands' and wives' household time-inputs are gross substitutes. In this case, an increase in male earnings would cause wives to substitute away from market work and towards household production.¹ For a proportion of women, the fall in equilibrium hours of market work would be such as to change their labour-force status from participants to non-participants. Thus, we predict that the participation of a female group will depend on male earnings possibilities.

Finally, we consider fertility-rates. In the literature on family labour supply, some authors (e.g. Mincer (1962)) have taken fertility as exogenous. We shall follow this assumption for the present.² The fertility-rate for a particular age-group is then taken

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1. Kneisner (1976) and Robinson and Tones (1982) report some support for the alternative assumption of complementarity of husbands' and wives' time. On the basis of this assumption, the response of female labour supply to male earnings possibilities could not be signed.

2. Though it will be questioned in Chapter 4.2 below.
as a determinant of the age-specific participation-rate of that group. The proposition arises from considering young children as female time-intensive: the presence of young children will raise the mother's shadow wage and make it less likely that she will participate in the labour-force. The prediction that the participation-rate of a female group will depend on the current fertility-rate of the group then follows only with the assumption that the presence of young children is adequately proxied by the current fertility-rate. Empirical evidence in favour of the prediction is provided by, inter alia, Heckman (1974) and Spencer (1973).

We have suggested that the participation-rate for a female demographic group depends on its unemployment-rate, on its fertility-rate, and on male and female earnings. We have referred here to female groups because it is for female groups that the more substantial variations in participation-rate have occurred. However, in principle, male participation-rates should be written with precisely the same determinants because the labour-supplies of men and women are jointly determined and, like women, men could be used by the household either in household work or in the labour market.

5. A Model of the Effects of Changes in Age Structure on Age-Specific Unemployment-Rates

The preceding analyses are now drawn together to formulate a testable model on the effects of changes in the age-composition of the working-age population. Let \( p_i \) be the proportion of the working-age population in demographic group \( i \); let \( w_i \) be the proportion of the labour-force in demographic group \( i \); and let \( FR_i \) and \( u_i \) be respectively...
the participation-rate and the unemployment-rate of group i. There will be eight demographic groups delineated: following the practice of Chapter II, we have four age-groups and each is now further divided by sex. Males groups are denoted by i=1 to i=4 (corresponding to the groups 15-19, 20-24, 25-44 and 45 years and above); female groups are denoted by i=5 to i=8 (corresponding to the same age-groups as in the male case). In this section, we discuss the model first in the context of the four female groups and then in the context of the four male groups.

The unemployment-rate for a demographic group is to be hypothesised to depend on the relative numbers of that group in the labour-force. It will of course depend also on a set of other variables. Amongst these we include the rate of unemployment of prime-males. This group of 'mainstream' adult workers is the largest demographic group in the labour-force and its unemployment-rate is used to proxy general conditions in the labour-market; variations in the prime-male unemployment-rate will reflect variations in aggregate demand conditions and variations in policies (e.g., employment taxes) that have their specific effects on the labour-market. Other variables included in our equation are a variable representing the minimum-wage conditions for males in the economy (wm) and a variable representing unemployment-insurance conditions in the economy (d). Reasons for including such variables were given in Section 2 of This Chapter and in Chapter 1.3

1. This is our cohort-crowding hypothesis.
2. Prime-males are those in the age-group, 25-44 years.
3. It accounted for 28-37% of the Canadian labour-force over 1953-78.
4. For females in the economy (wf) in the female equations.
respectively. Precise specification of all variables will be discussed in Chapter IV.

We propose then that age-specific unemployment-rates may be explained by relationships of the form

\[ u_i = f_1 (w_i, u_3, w_f, d) \quad i = 5, 6, 7, 8 \] (1)

In turn, the relative size of group \( i \) in the labour-force will depend on both the group's relative size in the population and on rates of labour-force participation. This relationship is described by the identity

\[ w_i = p_i \frac{\text{PR}_i}{\text{PR}} \] (2)

where \( \text{PR} \) is the participation-rate for the whole economy.

In formulating an equation to explain variations in \( \frac{\text{PR}_i}{\text{PR}} \), we must include determinants of both numerator and denominator. In the last section, the participation-rate of group \( i \) (\( \text{PR}_i \)) was viewed as depending on the own-group unemployment-rate (\( u_1 \)), on the own-group fertility-rate (\( f_{1} \)), and on male and female average (real) hourly earnings (\( e_m \) and \( e_f \)). By analogy, the participation-rate of the population as a whole will depend on the unemployment-rate of the population as a whole (\( U \)), on the fertility-rate of the population as a whole (\( f_{A} \)), and on male and female average (real) hourly earnings (\( e_m \) and \( e_f \)). The determinants of \( \frac{\text{PR}_i}{\text{PR}} \) so far discussed are therefore: \( u_1, U, f_{1} \).
fert_A, em and ef. 

We now add two further explanatory variables to the PR_1/PR equation. We enter p_i because the population-proportion may have a direct effect on PR_i insofar as it affects the average age of marriage. We also include wf (the minimum-wage variable) because the minimum wage will affect the probability of finding a job at any particular average wage. Thus, our equation for PR_1/PR reads

\[
\frac{PR_i}{PR} = f_3(u_i, U, fert_i, fert_A, em, ef, p_i, wf)
\]  

No particular functional-form is assumed for equations (1) and (3) at this stage. Substitution of equations (2) and (3) into (1) yields the reduced-form equation

\[
u_i = F_1(p_i, u_3, U, wf, fert_i, fert_A, em, ef) \quad i = 5, 6, 7, 8
\]  

For estimation purposes, this equation would bear the disadvantage that u_3 and U are extremely highly correlated: since prime-males are the largest group in the labour-force, their group unemployment-rate is a major element in constructing the aggregate unemployment-rate. We shall therefore have to omit U from our final version of the estimating equation. For similar reasons (multicollinearity), fert_A will also be omitted. We will therefore have the final reduced-form equation

1. If em and ef affect PR_1 and PR differentially, they will affect the value of PR_1/PR.
2. This relationship is proposed in Easterlin (1980, Chapter 3).
\[ u_i = F_2(p_1, u_3, w_f, d, f_{ert_i}, em, ef) \quad i = 5, 6, 7, 8 \quad (5). \]

We now consider the derivation of a similar set of reduced-form equations for the four male groups (signified by \( i = 1, 2, 3, 4 \)). Instead of the female minimum wage variable, we will now have a male minimum wage variable (\( wm \)). Otherwise, for the non prime-age groups (\( i = 1, 2, 4 \)), the reduced-form equation (again omitting \( U \) and \( f_{ert_A} \)) is the same as for the female demographic groups:

\[ u_i = G[p_1, u_3, wm, d, f_{ert_i}, em, ef] \quad i = 1, 2, 4 \quad (6). \]

The complication for the male case is that we also have to explain how prime-male unemployment is determined. The equation for \( u_3 \) can in fact be specified in the same way as for the other male age-groups except that a new controlling cyclical variable will be required. Writing this as \( g \),\(^1\) we have that

\[ u_3 = H(p_2, g, wm, d, f_{ert_i}, em, ef) \quad (7). \]

In Chapter IV equations (5), (6) and (7) will be estimated by simultaneous methods of regression analysis. Here we ask how our estimation will help us to assess the importance of demographic change as regards its effect on unemployment. From this point of view, the key coefficients in our estimated equations will be those on the demographic variables, \( p_1 \), which have been included to allow us to test the cohort-crowding hypothesis. If the estimated coefficients are found to be significantly greater than zero, this result will constitute

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1. The specification of this variable will be discussed in the next chapter.
support for the cohort-crowding hypothesis. Estimates of those coefficients on $p_i$ would then be used to derive an estimate on the overall impact of changes in the age-structure of the population. To use them in this way requires of course that we appreciate what the coefficients on the $p_i$ represent and, to explain this, we recall now the decomposition exercise of Chapter II (Table 2-2). In this table, the change in the overall unemployment-rate in Canada between each pair of trade-cycles was decomposed into three parts as follows:

i) an element unambiguously attributable to changes in labour-force weights ($\sum_i u_i^0 (w_i^1 - w_i^0)$);

ii) an element unambiguously attributable to changes in age-specific unemployment-rates ($\sum_i w_i^0 (u_i^1 - u_i^0)$); and

iii) an element measuring the interaction or covariance between age-structure and unemployment-rates ($\sum_i (w_i^1 - w_i^0)(u_i^1 - u_i^0)$).

The first element specified was termed the pure weights effect; this will be taken as our estimate of the impact of changes in age-structure in so far as the impact worked via changing labour-force weights.

The second element specified measures the impact on the overall rate of unemployment of variations in age-specific unemployment-rates. Depending on which pair of trade-cycles is considered, this source accounts arithmetically for 87-91% of total changes in the overall unemployment-rate.\(^1\) This is therefore the element in the

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1. Figures derived from Chapter II. (Table 2-2).
unemployment-variation that is likely to be the most fruitful subject for analysis. Our problem of course is to estimate how much of this element was attributable to demographic change rather than to changes in other exogenous variables. Our regression equations are designed to accomplish this task, i.e. to estimate the rates-effect of age-structure change.

The third element specified is the interaction or covariance term. It was argued in Chapter 2.2 that some of this interaction term originates with changes in population age-composition but that some of it must be associated with changes in the other exogenous variables that influence the $u_i$. Our task again is to separate out that part of the term that is associated with changes in our demographic variables.

How will our estimates of reduced-form equation

$$u_i = G(p_i, u_{wm}, d, fert, em, ef) \quad i = 1, 2, 4$$

assist in separating out those parts of components (ii) and (iii) that are attributable to demographic change? In Appendix III, for the simple case of equations in log-linear form, it is demonstrated that the coefficient on $p_i$ in this equation is a composite reflecting the coefficients on $w_i$ in equation (1) and on $u_i$ and $p_i$ in equation (3). The composite coefficient measures the total impact of a change in $p_i$ on the group-specific unemployment-rate, $u_i$: it incorporates both the direct effect of changes in $p_i$ and the effects in so far as they occur.
by modifying the group-specific participation-rates (PRᵢ). Estimates of this coefficient can thus be used to estimate the sum of the rates effect and that part of the covariance term associated with demographic change. Therefore, the parts of elements (ii) and (iii) that we need to separate out can be estimated together by reference to our regression equations. The combined estimate we shall term the augmented rates effect of demographic change.

In Chapter II, we discussed the merits of three possible ways of measuring the weights effect of demographic change. The 'pure weights effect' omitted the interaction term altogether; the 'Kitagawa weights effect' included one-half of the interaction term; the 'full weights effect' included all of the interaction term. The basis for choosing between these three measures is now clear from the foregoing analysis. Our estimate of the augmented rates effect already includes the relevant part of the interaction term. To include any of the interaction term in measuring the weights effect would therefore constitute double-counting. Our choice of measure for the weights effect will therefore be the pure weights effect. To this we shall add estimates of the augmented rates effect. This will permit us to gauge the total effect on overall unemployment of changes in the age-composition of the population.
Chapter IV

The Effect of Changes in Age-Structure on Unemployment

in Canada, 1953-78

1. Methodology

In Chapter II (Table 2-2) the change in the overall unemployment-rate in Canada as between each pair of trade-cycles in the period was decomposed into three parts, as follows:

i) An element unambiguously attributable to changes in labour-force weights \( \left( \sum_i u_i^0 (w_i^1 - w_i^0) \right) \),

ii) An element unambiguously attributable to changes in age-specific unemployment-rates \( \left( \sum_i w_i^0 (u_i^1 - u_i^0) \right) \),

iii) An element measuring the interaction or covariance between age-structure and unemployment-rates \( \left( \sum_i (w_i^1 - w_i^0)(u_i^1 - u_i^0) \right) \).

The first element specified was termed the pure weights effect; this will be taken as our estimate of the impact of changes in age-structure on overall unemployment in so far as the impact worked via changing labour-force weights. It is an estimate because changing weights could be the result of exogenous changes in the participation-rate as well as changes in population age-compositon. However, the bulk of the variation in labour-force age-structure has been associated with corresponding changes in population age-structure (as was
illustrated in Chapter 1.4):

The second element specified measures the impact on the overall rate of unemployment of variations in age-specific unemployment-rates. Depending on which pair of trade-cycles is considered, this source accounts arithmetically for 87-91% of total changes in the overall unemployment-rate. This is therefore the element in the unemployment-variations that is likely to be most deserving of analysis. The problem to be tackled is that some of the changes in age-specific unemployment-rates may be caused by demographic change and some by changes in other exogenous variables (such as unemployment-insurance conditions). Regression equations of the form discussed in Chapter 3.5 will be used here to separate out the impact on overall unemployment of those changes in age-specific unemployment-rates that can be attributed to changes in population age-composition.

The third element specified, the covariance or interaction term, is small but relatively difficult to interpret. It was argued in Chapter 2.2 that some of this interaction term arises from demographic change. It was further argued, in Chapter 3.5, that, in the type of regression equation specified, the coefficient on the demographic variable would reflect not only the rates effect but also that part of the covariance term associated with demographic

1. figures derived from Chapter II, Table 2.2
change. Estimates based on our regression results therefore embrace both the rates effect and the relevant part of the covariance term. Because our measure embraces both of these parts, we will term it the augmented rates effect. The total effect of changes in population age-composition will be estimated by summing estimates of the pure weights effect and the augmented rates effect.

We now turn to our task of deriving an estimate of the augmented rates effect by considering the effect on age-specific unemployment-rates of changes in population age-structure. Earlier studies for the United States have suggested that this effect is potentially important. A time-series analysis by Wachter and Kim (1979) tested what they called the 'cohort overcrowding' hypothesis of youth unemployment. They successfully used a variable (termed R'y in their paper) that represented the proportion of the adult population falling within a specific (youth) age-group. They found that increases in youth-unemployment rates could be attributed in part to increases in the relative size of youth age-groups. A similar finding came out of a cross-section study reported by Freeman and Wise (1980, p. 12). That study concluded that, across geographical areas, an increase in the relative size of the youth population was associated with a reduction in the employment-rate of young people.

The present study utilises regression analysis to examine the effect of changes in group-size on unemployment-rates for eight groups in Canada; the groups are delineated by age (15-19 years,
20-24 years, 25-44 years, 45 years and above) and delineated also by sex. The general system of equations estimated is of the reduced-form derived in Chapter 3.5:

\[ u_i = f_i \left( p_i, u_3, \text{wm, d, fert}_i, \text{em, ef} \right) \quad i = 1, 2, 4 \]

\[ u_3 = f_3 \left( p_3, g, \text{wm, d, fert}_1, \text{em, ef} \right) \]

\[ u_i = f_i \left( p_i, u_3, \text{wf, d, fert}_i, \text{em, ef} \right) \quad i = 5, 6, 7, 8 \]

Here, \( u_i \) is the unemployment-rate for group \( i \); \( p_i \) is the proportion of the 15-69 year old population accounted for by group \( i \); \( u_3 \) is the prime-male unemployment-rate; \( g \) is a variable representing general demand conditions in the economy; \( \text{wm} \) and \( \text{wf} \) are variables representing minimum-wage conditions; \( \text{d} \) is a variable representing unemployment-insurance-conditions; \( \text{fert}_i \) is a fertility variable; and \( \text{em} \) and \( \text{ef} \) are earnings variables.

In the specification of \( p_i \), the cut-off point of 69 years was based on the age-group divisions made in the Labour Force Survey. There, the oldest age-group (45 years and above) has no upper age-limit. However, very few of that group will be 70 years of age or older; 69 years was therefore taken to be an acceptable upper limit for 'nominal' labour-force participation. Changes in numbers of people over 69 years in the population have been assumed to have no labour-force implications.

In the equations specified, \( p_i \) is of course included to allow us to test the hypothesis that cohort-relative-size affects the cohort-unemployment-rate. Note that \( g \) is a variable based on
population figures rather than on labour-force figures. The $p_1$ therefore constitute truly exogenous variables.

We consider again now which other variables it is appropriate to include in the regression equations. The study of Wachter and Kim (1979) used only one variable here, the rate of unemployment for prime-males; the role of this variable was to control for variations in demand conditions in the economy as a whole. This study for Canada also utilizes such a variable but to it has been added variables chosen to reflect the influence of variations in minimum-wage and unemployment-insurance conditions. Because of their influence on participation-rates, we include also variables representing fertility-rates and male and female earnings.  

Choice of Aggregate Demand Variable

As in Wachter and Kim (1979), the variable chosen to represent the impact of general labour-market conditions on age-specific unemployment-rates is the prime-male unemployment-rate; here, prime-males are defined as those aged 25-44 years. This variable will pick up the effects of both variations in aggregate demand and variations in the relative attractiveness of employing labour rather than other factors of production (for example, it would reflect the

1. Chapter 3.3 explained why unemployment-insurance conditions may be predicted to affect unemployment-rates. Chapter 3.2 explained why minimum-wage laws may be expected to affect unemployment-rates.

2. See Chapter 3.5.
effects of the imposition of a payroll tax.¹

Unlike the paper by Wachter and Kim, this study hopes to be able to 'explain' the prime-male unemployment-rate itself as well as unemployment-rates for other groups. Another cyclical variable therefore had to be chosen for inclusion in the prime-male equation. The one chosen was the 'G.N.P. gap', g. This variable, g, was generated by taking a series, 'G.N.P. measured in 1972 dollars' and fitting a log-linear trend to the twenty-six years of data. Residuals were then retrieved and used as the G.N.P. gap variable (with positive values representing shortfalls of real G.N.P. from trend and negative values representing real G.N.P. figures above the fitted trend). The variable may be interpreted as measuring the gap between actual and 'potential' output.

The G.N.P. gap was chosen as the controlling cyclical variable in the prime-male equation in preference to rival candidates such as the real growth rate or the real growth rate lagged. The principal reason for making this choice was that the dependent variable in our equation is a measure of the level of unemployment at a point in time; the G.N.P. gap reflects the pressure of aggregate demand, also measured at a point in time, and is therefore appropriate as an independent variable in this case. The rate of growth of output would be more suitable if the dependent variable

¹ A more general indicator of labour-market tightness would of course have been the overall rate of unemployment (rather than the rate for the group, prime-males). This however could not be used since the dependent variable in each of the regression equations would then have constituted part of the explanatory variable, overall unemployment.
being explained were the rate of change of unemployment over time rather than unemployment measured at a point in time.\(^1\)

One problem with the G.N.P. gap measure may be mentioned here. It arises because the regression equation for prime-males includes also a variable representing the effects of unemployment-insurance payments; estimates of the resultant coefficient in fact proved to be statistically significant. To the extent that increased unemployment benefits raised the unemployment-rate, it seems likely that it would also have dragged real G.N.P. further below its trend than it would otherwise have been. Part of the effect of increased benefit payments would thus be attributed to the cyclical variable rather than to the unemployment-insurance variable. The coefficient on the latter would therefore be an underestimate of the true figure. No attempt is made here to allow for this source of error because the impact of unemployment insurance is not the focus of the study. However, it may be of interest to note that Benjamin and Kochin (1979), in their study of inter-war British unemployment, estimated that the coefficient on the benefits variable picked up only two-thirds of the total unemployment-insurance effect, the other one-third appearing in the coefficient on the cyclical variable.

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1. An additional reason for preferring the G.N.P. gap is that the measurement of G.N.P. is seriously subject to error. The effect of measurement errors would be unnecessarily compounded by using the rate of change of G.N.P.
Choice of Minimum Wage Variable

Some measure of movements in Canadian wage rates was required for the regression analysis because it was expected (on the basis of the argument in Chapter 3.2) that these movements would have different unemployment effects on different age-groups. The measure used is a minimum wage index (compiled separately for males and females) described in Swidinsky (1980). Prof. Swidinsky kindly supplied the indices for the period 1953-75 and these were then updated to 1978 using his methods and data sources. This minimum wage index, whether for males or females, takes account of both the level of the minimum wage (relative to the average wage) in each province and of the rate of coverage of the law in each province. It is similar to the index used for the United States by Kaitz (1970), Lovell (1972), Welch (1974), Mincer (1976) and Ragan (1977). The Canadian index is given by

\[ MWI = \sum_{i=1}^{10} \frac{L_i}{L_c} \frac{M_{W_i}}{AHE_i} C_i \]

where the subscripts \( i \) refer to the ten provinces and the subscript \( c \) refers to Canada. \( MW \) is the minimum hourly wage, \( C \) is the proportion of the provincial non-agricultural labour force to which the law applies, \( L \) is the non-agricultural labour force, and \( AHE \) is average hourly earnings in manufacturing. Appendix IV tabulates this minimum wage index for Canada (for males and females separately) for the period 1953-1978. Fluctuations in the indices

1. Canadian minimum wage rates do not apply to agriculture.
are mainly explained by the fact that the variable AHE has increased annually whereas the legislated MW has tended to change with a frequency much less than once a year. The sharp rise in the male index in 1953 occurred because Ontario extended its general minimum-wage law to males in that year. For all provinces, C_1 has been one in recent years. Note finally that the MW figure used is that applicable to adults. From time to time, provinces have had special regulations on the employment of juveniles but a separate index for juveniles proved impossible to construct given the complexity of the rules and the inadequacy of the data. It is therefore assumed that movements in youth-wage-minima are reflected adequately by variations in the minimum-wage-index for Canada.

It should be noted here that, while (as stated on the previous page) our minimum-wage index is the same in form as that used in several previous studies, it may nevertheless be subject to criticism. Two restrictions are inherent in the specification of this form of minimum-wage-index. First, the coverage-ratio is assumed to enter multiplicatively with the level of the wage minimum. Secondly, the minimum wage is measured relative to average hourly earnings. Neither restriction has been given theoretical justification. Neither, for example, is implied by Mincer's (1976) analysis discussed in Chapter 3.2 above. For example, in Mincer's analysis, it is the percentage absolute rather than the percentage relative gap between covered- and uncovered-sector wage-rates that is significant.
Despite this weakness in its specification, we have nevertheless followed the currently standard method for the construction of a minimum-wage-index.

Choice of Unemployment-Insurance Variable

The next variable to be included in the unemployment equations is intended to represent variations in unemployment-insurance conditions. The Unemployment Insurance Act (1971) radically altered the rules governing unemployment benefit in Canada in the direction of making more generous payments to those without a job. If the replacement-ratio for Canada is calculated, by far the greater part of the variation in its value is accounted for by the introduction of the rules set out in this 1971 Act. It was therefore decided to use a dummy variable to take account of the influence of this legislation. The dummy is set equal to zero for each year from 1953 to 1970, to 0.5 for 1971 (the year

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1. See, for example, Holden and Peel (1979, p. 614).
during which, on July 1st, the Act came into force) and to one for each year from 1972 on. It was felt preferable to include a dummy variable rather than the calculated replacement-ratio because the replacement-ratio was calculable only for unemployed people generally. It may have had very different values across age-groups because of the nature of the eligibility rules and because of variations in the average wage across age-groups. Inclusion of the dummy should enable us to control for the differential impact of the Act on the several demographic groups.

Choice of Fertility Variable

Age-specific fertility-rates were required for inclusion in the model because they were hypothesised to influence group participation-rates and therefore labour-force weights. The fertility-rate for a group was taken as

\[ \text{fert}_i = \frac{\text{number of live births to mothers in age-group } i}{\text{population of females in age-group } i} \]

The relevant numbers of births for our four age-groups were obtained from various issues of the United Nations Demographic Yearbook and the Statistics Canada publication, Vital Statistics. For the 45- and-above age-group, the fertility-rate was always close to zero; it was therefore decided that the variable \( \text{fert}_i \) should be omitted from the equation for that group.

For males, we also required a variable to proxy the presence of young children in the household. In Canada, there are no data for number of births by age of father. For each male age-
group, we thus had to take the relevant fertility-rate as the
fertility-rate for women of the same age-group. There is a degree
to which this is only an approximate solution, since spouses may
be in different age-groups; however, because age-specific female
fertility-rates are highly correlated with each other, this seems
unlikely to be a serious source of error.

Choice of Male and Female Earnings Variables

It was felt desirable to include measures of both male and
female earnings possibilities because the labour-force participation
decision of an individual is predicted to be affected both by own-
and by spouse's earnings rate (Chapter 3.4). Unfortunately, sex-
specific earnings series covering the whole of our period were not
available for Canada. Average hourly earnings by sex were collected
for manufacturing up to 1969 (though not for 1961 and 1962, when
the survey was not carried out). Beyond 1969, only hourly earnings
across manufacturing as a whole (i.e. not by sex) were available.
The construction of reliable series for male and female earnings
thus presented some difficulty.

The sex-specific earnings series (em and ef) were in fact
constructed as follows. For 1953-69, average hourly earnings for
males and females in manufacturing were deflated by the Consumer
Price Index, so that they were expressed in constant (1971)
dollars. For 1980, the data available included both average

1. Source: Statistics Canada (Catalogue Number 72-204), Earnings
and Hours of Work in Manufacturing.
earnings by sex and average earnings of all workers. It was assumed that after 1969 male and female (real) earnings grew at the same percentage rate as (real) earnings for all workers in manufacturing: this was of course equivalent to assuming that male and female wage relativities remained constant after 1969. Missing observations for 1961 and 1962 were filled in by a similar method.

Even for the period up to 1969 (the period of raw rather than constructed data), the simple correlation coefficient for the two variables em and ef was extremely high (+.995): for the period after 1969, male and female earnings were assumed to grow exactly in step with each other, so that the correlation between our series for the whole period is even higher. The two series thus turn out to be almost identical: the co-linearity problem is too severe to allow the inclusion of both em and ef in regression equations. Although in principle em and ef belong in the unemployment equation for all demographic groups, in practice we had to decide to use only em in the male equations and only ef in the female equations.

2. Female Unemployment: Specification of the Model and Regression Results for Canada, 1953-78

Let the four female age-groups, 15-19 years, 20-24 years, 25-44 years, and 45 years and above, be denoted by subscripts $i = 5, 6, 7, \text{ and } 8$ respectively. The model discussed then consists of the four equations.
\[ u_i = f_i (p_i, w_3, w_f, d, d_{fert}, e_f) \quad i = 5, 6, 7, 8, \]

where \( u_i \) is the unemployment rate for group \( i \), \( p_i \) the population-proportion, \( w_f \) the female minimum wage index, \( d \) the unemployment-insurance dummy, \( d_{fert} \) the fertility-rate (excluded for \( i = 8 \)) and \( e_f \) average female earnings.

One problem with estimating these equations as they stand is that the dependent variable is each case is not free to vary between minus and plus infinity as the classical regression model requires. A common solution to such a problem is to undertake a logistic-transformation. This involves replacing, for example, the unemployment-rate by the odds of being unemployed and then taking the log of those odds. Suppose one starts with a percentage unemployment rate, \( u_i \). This is transformed to the new variable

\[ y_i = \log \left( \frac{u_i}{1-u_i} \right). \]

This new variable may in principle vary between minus and plus infinity.

The logit-equations estimated in this case were of the form

\[ y_i = f_i (p_i, w_3, w_f, d, d_{fert}, e_f) \quad i = 5, 6, 7, 8. \]

An alternative was to transform the relevant right hand side variables to the log-of-odds form also. There was no econometric reason for doing or not doing this but it proved to be a marginally
better specification in terms of results (and also made simultaneous estimation of the four equations easier). Therefore, the regression results embodied in this chapter take the form

\[ y_i = \alpha + \beta_1 x_1 + \beta_2 v_3 + \beta_3 w_6 + \beta_4 d + \beta_5 b_{ay}_i + \beta_6 e_{f} \quad i = 5, 6, 7, 8 \]

where \( x_1 = \log \left( \frac{p_i}{1-p_i} \right) \) and \( b_{ay}_i = \log \left( \frac{1}{f_{ert}} \right) \).

Estimation of these equations by ordinary-least squares would have been inefficient because of the likelihood that error terms across the four equations would be correlated. Estimation was therefore by the method of seemingly unrelated regression equations (SURE).\(^1\)

The results of estimation by SURE are given in Table 4-1(a). An immediately puzzling feature of these results was the failure of the fertility variable to contribute to our understanding of female unemployment. Examination of the correlation matrix revealed the likelihood of a co-linearity problem: the simple correlation coefficients between female earnings and the fertility-rates for groups 5, 6 and 7 were -.97, -.96, and -.955, respectively. The equations were therefore re-estimated by SURE, omitting first \( e_f \) (Table 4-1(b)), and then, secondly, replacing \( e_f \) and omitting \( b_{ay} \) instead (Table 4-1(c)). In each case, the omission of the one variable significantly 'improved' the performance of the other, \textit{prima facie} evidence of the importance of the co-linearity problem in this case.

\[ 1. \text{ See Zellner (1962).} \]
<table>
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<tr>
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<th>$y_5$</th>
<th>$y_6$</th>
<th>$y_7$</th>
<th>$y_8$</th>
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(a) Independent variables: $X_1$, $X_2$, $X_3$, $X_4$, $X_5$, $X_6$, $X_7$, $X_8$, $X_9$.
Table 4-1 (continued)

(b)

| dependent variable | constant | $x_1$ | $y_3$ | $w_1$ | $d$ | baby | ef | $R^2$ | D.W. |
|--------------------|----------|-------|-------|-------|-----|-------|    |       |      |
| $y_5$              | 3.21     | 1.85  | 0.91  | 0.72  | 0.007 | -0.65* |    | .97   | 1.71 |
|                    | (2.13)   | (5.10)| (14.07)| (0.71)| (0.09)| (-3.59)|    |       |      |
| $y_6$              | -5.23    | -1.01 | 0.91  | 2.02  | 0.16  | -0.96* |    | .97   | 1.50 |
|                    | (-1.72)  | (-1.06)| (13.70)| (3.04)| (1.62)| (-2.69)|    |       |      |
| $y_7$              | -1.65    | 1.24  | 0.63  | 1.17  | 0.35  | -0.07* |    | .97   | 1.90 |
|                    | (-2.81)  | (1.33)| (11.98)| (1.30)| (5.00)| (-3.31)|    |       |      |
| $y_8$              | -4.90    | -0.97 | 0.53  | 3.79  | 0.27  |         |    | .90   | 1.89 |
|                    | (-1.15)  | (-0.45)| (6.09)| (2.05)| (3.23)|        |    |       |      |
Table 4-1 (continued)

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<th>$e_f$</th>
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<th>D.W.</th>
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</table>

Symbols are explained in the text.
t-statistics appear in brackets.
asterisks indicate statistical significance at the 5% level.
For estimation purposes, it was felt desirable to use a set of equations from which either ef or baby were omitted; this was the only way in which we could avoid the statistical imprecision occasioned by multicollinearity. Should ef be omitted or should baby be omitted? There would be no obvious basis for choice if both were indeed exogenous variables as we have thus far assumed them to be. However, there are theoretical and empirical grounds for unease about treating the fertility variable as exogenous. As Becker put it, "it is not surprising that the number of children a family has is strongly negatively related to the mother's value of time as measured by her wage if she is in the labour force" [Becker (1974), p. 318]. This negative relationship predicted between fertility and female earnings potential has received strong empirical backing (from, e.g., Butz and Ward (1979)). It appears therefore that fertility would be more appropriately treated as an endogenous variable dependent on female earnings in which case it has no proper place in our reduced-form regression equations.

On the basis of the argument outlined here, it was decided to use for estimation purposes the set of results given in panel (c) of Table 4.1. These results form the basis of the comments in the remainder of this section.

---

1. In so far as our principal concern is with the population variable, \( x_1 \), it will be noted that our conclusions would not have been very different had we instead chosen the sets of results given in panels (a) or (b).
Comments on Female Regression Results (Table 4-1(c))

Population Variable

The cohort-crowding hypothesis suggests that the coefficient on the population-proportions variable should be positive. In fact, the estimate of this coefficient is positive and significantly greater than zero (at the 5% level of significance) only for the 15-19 year old group. It proved not to be significantly different from zero for any of the three older age-groups.

In the context of these results, it may be appropriate to reiterate the prediction (based on Welch's (1979) work) of Chapter 3.3 with respect to the relative importance of the cohort-crowding hypothesis for females and males. We predicted that female age-specific unemployment-rates might be less sensitive to cohort-size than corresponding male rates because of greater ease or substitution between workers of different ages in typical female occupations. It will be interesting to note in the following section whether or not this prediction will be supported in terms of stronger results on the population variable, \( x_1 \).

Cyclical Variable

The cyclical variable, \( y_3 \) (the log-of-odds of unemployment for prime-males), is very strongly significant in all four equations. This indicates that the dominant influence on the unemployment-rate of a particular female age-group is the general state of the labour-market.
Minimum Wage Variable

The coefficient on the variable wf is significantly positive only for the oldest group. It is not significantly different from zero (at the 5% level) for any of the other three groups. This result is particularly puzzling for the youngest age-group on which, we predicted, the effects of formal minimum-wage laws would impinge with particular severity. The explanation may be the failure of our index, based on adult female minimum-wage rates, to trace female youth minimum-wage rates with sufficient accuracy.

Unemployment Insurance Variable

The estimated coefficient on the unemployment-insurance variable is not significantly different from zero for the youngest female group. This is not surprising since many teenagers do not participate in the scheme. The estimated coefficient is positive and significant (at the 5% level) for the middle two age-groups. For the oldest age-group, the coefficient is not significantly different from zero.

Female Earnings Variables

To the extent that it raised the participation-rate (and therefore the labour-force weight) of a group, female average earnings were predicted to raise female group-unemployment rates.¹

¹ Changes in ef and em occurred together, but any income effect against female participation, arising from increases in em, was ignored in this prediction of the sign on ef in our regressions.
In these equations, the prediction receives some support since the coefficient on ef is statistically significant in all four equations.

General Comments

Autocorrelation does not appear to be a serious problem in the estimation of the female model. At the 5% level of significance, one cannot reject the hypothesis that there is zero autocorrelation in the \( y_5 \), \( y_6 \) and \( y_7 \) equations; the Durbin-Watson statistic is within the inconclusive range in the \( y_8 \) equation.

The interpretation of coefficients is postponed until after the presentation of male results.

3. Male Unemployment: Specification of the Model and Regression Results for Canada, 1953-78

Let the four male age-groups, 15-19 years, 20-24 years, 25-44 years, and 45 years and above, be denoted by subscripts \( i=1,2,3 \) and 4 respectively. The model to be discussed is then very similar to that for females: the only differences are that a male minimum wage index (wm) and male average earnings (em) are substituted for the corresponding female data and, of course, a separate equation is needed to explain prime-male unemployment. Thus, the model specified in this section (with variables again couched in the log-of-odds form) consists of the four equations:

\[
y_i = f_1(x_i, y_5, \text{wm}, d, \text{baby}_i, \text{em}) \quad i = 1,2,4,
\]

\[
y_3 = f_1(x_3, g, \text{wm}, d, \text{baby}_3, \text{em}) \quad i = 3.
\]
Here, \( g \) is the G.N.P. gap, and baby is omitted for \( i = 4 \).

Because single equation estimation of simultaneous equation models yields inconsistent estimates, the system was estimated using the method of three-stage least-squares (3SLS). This method involves first estimating each equation using two-stage least-squares; the residuals of each equation are then utilised to estimate the cross-equation variances and covariances; the third stage of the procedure consists of obtaining generalised-least-squares parameter estimates. The advantage compared with two-stage least-squares is that three-stage least-squares takes into account cross-equation correlation. There is therefore a gain in efficiency. 3SLS estimates for our system are presented in Table 4–2.

Panels (a), (b) and (c) of Table 4–2 correspond to the panels so labelled in the female results table: in (a), results for the full model are reported; for (b), the average earnings variable was omitted; for (c), it was replaced and the fertility variable dropped instead. As in the female case, the full model was inappropriate to use because of a multicollinearity problem: there was very strong negative correlation between the earnings and fertility variables. Again, and for the same reasons as in the female case, it was felt preferable to choose to drop the latter and so, as between (a), (b) and (c), the results set (c) were preferred.

A further co-linearity problem in this case arose because of

---

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<th>$wm$</th>
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<th>em</th>
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(d).

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<td>(4.24)</td>
<td>(20.00)</td>
<td>(1.21)</td>
<td>(0.21)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_2$</td>
<td>1.60*</td>
<td>0.33*</td>
<td>1.03*</td>
<td>0.22</td>
<td>0.21*</td>
<td></td>
<td></td>
<td></td>
<td>.98</td>
<td>1.69</td>
</tr>
<tr>
<td></td>
<td>(3.22)</td>
<td>(2.09)</td>
<td>(5.27)</td>
<td>(1.53)</td>
<td>(5.90)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_3$</td>
<td>-5.93*</td>
<td>-2.70</td>
<td>6.89*</td>
<td>-2.70*</td>
<td>0.35*</td>
<td></td>
<td></td>
<td></td>
<td>.63</td>
<td>1.08</td>
</tr>
<tr>
<td></td>
<td>(-3.56)</td>
<td>(-1.79)</td>
<td>(5.27)</td>
<td>(-2.64)</td>
<td>(3.81)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_4$</td>
<td>5.02*</td>
<td>3.43*</td>
<td>0.84*</td>
<td>0.36*</td>
<td>-0.08</td>
<td></td>
<td></td>
<td></td>
<td>.93</td>
<td>1.97</td>
</tr>
<tr>
<td></td>
<td>(3.39)</td>
<td>(3.85)</td>
<td>(14.20)</td>
<td>(2.42)</td>
<td>(-1.82)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Symbols are explained in the text.
t-statistics appear in brackets.
asterisks indicate statistical significance of the 5% level.
a positive correlation between the male minimum wage index and male average earnings (the simple correlation coefficient was +.83). The problem was resolved by dropping the \( \gamma \) variable – thus we chose to take the results set (d) as our final preferred set of results for the estimation of male unemployment equations using 3SLS. Fortunately, the results on the population variable (our central concern) were broadly similar across the sets of results reported; the choice between them is therefore not crucially important.

The main problem with all the results sets presented in Table 4-2 – including our 'preferred' set (d) – lies in the equations to explain \( y_3 \). There is some indication of autocorrelation here with a Durbin-Watson statistic (in case (d)) of 1.08. Actually, this lies within the inconclusive range but nevertheless there was cause for concern because of the crucial role played by \( y_3 \) in our system of equations – \( y_3 \), in a sense, 'drives' the other three equations because \( y_3 \) 'explains' the bulk of variations in \( y_1 \), \( y_2 \) and \( y_4 \). Tests were therefore carried out for the presence of first or higher orders of autocorrelation. These suggested the presence of second-order autocorrelation. The \( y_3 \) equation was thus re-estimated as a single equation using a Cochrane-Orcutt transformation procedure for second-order serial correlation. The results appear as Table 4-3.

The new specification of the \( y_3 \) equation seems entirely satisfactory and it appeared desirable to incorporate it within the estimation of the system of four male equations. Unfortunately, a
Table 4.3
Prime-Male Unemployment, CORC2 Estimation

<table>
<thead>
<tr>
<th>dependent variable</th>
<th>constant</th>
<th>$x_3$</th>
<th>$g$</th>
<th>$wm$</th>
<th>$d$</th>
<th>$R^2$</th>
<th>D.W.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$y_3$</td>
<td>-1.79</td>
<td>0.98</td>
<td>7.97*</td>
<td>-0.05</td>
<td>0.34*</td>
<td>.81</td>
<td>1.96</td>
</tr>
<tr>
<td></td>
<td>(-0.96)</td>
<td>(0.63)</td>
<td>(5.96)</td>
<td>(-0.64)</td>
<td>(2.62)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

$RHO_1 = 0.66$

$RHO_2 = -0.10$

Symbols are explained in the text. T-statistics appear in brackets. Asterisks indicate significance at the 5% level.
statistical package allowing correction for autocorrelation within the context of three-stage-least-squares estimation was not available. The following ad hoc procedure was therefore adopted. First, the first and second order coefficients of autocorrelation (RH01 and RH02) were retrieved from the results in Table 4-3. Each variable \(v_i\) in the \(y_3\) equation was then transformed as follows

\[
v_i^* = v_i - (RH01) v_i t-1 - (RH02) v_i t-2
\]

These new variables were then used in the equation for \(y_3\) which was itself estimated as part of the three-stage-least-squares estimation of the whole male unemployment system. The results of this procedure are presented in Table 4-4 which is the basis for the comments in the remainder of this section.

Comments on Male Regression Results (Table 4-4)

Population Variable

The coefficient on the variable \(z_i\) was expected to be positive in accordance with our "cohort-crowding" hypothesis. In fact, the coefficient is positive and significant (at the 5% level) for three of the four male age-groups. For the prime-male group (25-44 years), the coefficient is not significantly greater than zero but this was expected since the group is so large that, relative to its total size, fluctuations in its relative importance in the population have been small.
Table 4-4

Male Unemployment – 3SLS Estimation with Correction for Autocorrelation in the Prime-Male Equation

<table>
<thead>
<tr>
<th>dependent variable</th>
<th>constant</th>
<th>$x_1$</th>
<th>$y_3$</th>
<th>$g$</th>
<th>$wm$</th>
<th>$d$</th>
<th>$y_3^A$</th>
<th>$\bar{R}^2$</th>
<th>D.W.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$y_1$</td>
<td>2.50*</td>
<td>0.63*</td>
<td>0.90*</td>
<td>0.41*</td>
<td>-0.005</td>
<td>.98</td>
<td>2.04</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(4.50)</td>
<td>(3.46)</td>
<td>(26.44)</td>
<td>(1.95)</td>
<td>(-0.24)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_2$</td>
<td>1.79*</td>
<td>0.34*</td>
<td>1.08*</td>
<td>0.28*</td>
<td>0.19*</td>
<td>.99</td>
<td>1.89</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(3.29)</td>
<td>(1.95)</td>
<td>(25.92)</td>
<td>(1.82)</td>
<td>(5.54)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_3$</td>
<td>-1.49</td>
<td>0.71</td>
<td>9.04*</td>
<td>-1.10</td>
<td>0.40*</td>
<td>0.70*</td>
<td>.82</td>
<td>2.03</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-2.00)</td>
<td>(0.42)</td>
<td>(6.90)</td>
<td>(-1.42)</td>
<td>(3.17)</td>
<td>(3.82)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_4$</td>
<td>4.69*</td>
<td>3.28*</td>
<td>0.80*</td>
<td>0.29*</td>
<td>-0.08</td>
<td>.93</td>
<td>1.94</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(3.17)</td>
<td>(3.64)</td>
<td>(16.35)</td>
<td>(1.98)</td>
<td>(-1.69)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Note: This model corresponds to that reported in Table 4-2(d)
In the $y_3$ equation, each independent variable was transformed by the formula

\[
v_i^* = v_i - 0.66 v_{i-1} + 0.10 v_{i-2}.
\]

\[
y_3^{A} = -0.66 y_{3,t-1} + 0.10 y_{3,t-2}.
\]

(see text for explanation)

Symbols are explained in the text
$t$-statistics appear in brackets
asterisks indicate statistical significance at the 5% level.
Cyclical Variable

The variable representing aggregate demand conditions is very strongly significant in each case; variations in prime-male unemployment "explain" most of the variation in the unemployment rates of other groups. The "gap" variable performs strongly in the prime-male equation.

Minimum Wage Variable

Increases in the minimum wage are predicted to affect most adversely the unemployment-rates of low-paid age groups. For "non-prime" equations, the coefficient on \( w_m \) was therefore expected to be positive. For prime-males, there was no a priori expectation as to sign for while some members of the group may lose jobs as a result of the minimum wage, others may benefit because of substitution between them and groups whose price has been artificially raised. In fact, for all groups but prime-males, the coefficient turns out to be positive and significant at the 5% level (one tailed-test); for prime-males, it was negative but not significantly different from zero at the 5% level of significance.\(^1\)

Unemployment Insurance Variable

Teenagers are limited in the extent to which they are able to participate in the Canadian unemployment-insurance scheme; this accounts for the zero coefficient on the unemployment-insurance dummy.

---

1. A caveat in interpreting the minimum-wage results is that the strong correlation between the minimum wage index and male average earnings makes it uncertain which variable's effects are in fact reflected in the coefficient estimate.
in the equation for \( y_1 \). For the other groups, the coefficient on \( d \) was expected to be positive. In fact, it is strongly positive for the 20-24 and 25-44 age groups. However, for "45 and above" it is negative, though not significantly less than zero at the 5% significance level. No explanation is readily apparent for this result in the \( y_4 \) equation.

General Comments

The Durbin-Watson statistics show an absence of first-order serial correlation in the four equations. While the equations in general seem to be satisfactory, the interpretation of the coefficients requires explanation because of the use of the "log of odds" form. The interpretation will be presented in the next section.

4. Interpretation of Male and Female Regression Results

The results presented in the last two sections included equations in the form

\[
\log \frac{u_i}{1-u_i} = \alpha + \beta_1 \log \frac{p_i}{1-p_i} + \beta_2 \log \frac{u_3}{1-u_3} + \beta_k Z_k
\]

\( i = 1, 2, 4, 5, 6, 7, 8 \)

\( (Z) = \) other variables

\( k \) signifies the \( k^{th} \) variable

As they stand, the coefficients in such an equation are hard to interpret. Ideally, we would instead like to have values for expressions such as \( \frac{\delta u_i}{\delta p_i} \) and \( \frac{\delta u_i}{\delta u_3} \). Fortunately, these may be obtained by a little
manipulation. For example, if we totally differentiate the above equation, we can derive the result that

\[
\frac{\delta u_1}{\delta p_1} = \beta_1 \frac{u_1 (1-u_1)}{p_1 (1-p_1)}
\]

Similarly,

\[
\frac{\delta u_1}{\delta u_3} = \beta_2 \frac{u_1 (1-u_1)}{u_3 (1-u_3)}
\]

and

\[
\frac{\delta u_1}{\delta d} = \beta_4 u_1 (1-u_1).
\]

All these expressions were evaluated at the means of variables for each case where the estimate of the relevant regression coefficient was statistically significant (at the 5% level). The results of the exercise are tabulated in Table 4-5. The table also present the corresponding elasticities.

The results indicate, for example, that a one percentage point rise in the proportion of the 15-69 year old population accounted for by 15-19 year old males was associated with a 1.1 percentage-points rise in their group unemployment-rate. Females of the same age were somewhat more prone to suffer from an increase in their numbers. The highest elasticity of all was for the oldest male group. This last group was benefitting from falling numbers; presumably, to the extent that new workers were being absorbed by the labour market, there was a corresponding rise in the demand for complementary inputs (including experienced labour).
### Table 4-5

**Interpretation of Regression Results**

<table>
<thead>
<tr>
<th>Group</th>
<th>Population Variable</th>
<th>Elasticity of Unemployment Rate w.r.t. Pop. Prop.</th>
<th>Cyclical Variable</th>
<th>Elasticity of Unemployment Rate w.r.t. Prime-Male Unemployment</th>
<th>U.I. Variable</th>
<th>Effect on $u_i$ of Setting U.I. Dummy to one</th>
</tr>
</thead>
<tbody>
<tr>
<td>Male 15-19</td>
<td>$\frac{\delta u_i}{\delta p_i}$</td>
<td>1.10</td>
<td>0.50</td>
<td>2.71</td>
<td>0.82</td>
<td>n.s.</td>
</tr>
<tr>
<td>Male 20-24</td>
<td>$\frac{\delta u_i}{\delta p_i}$</td>
<td>0.44</td>
<td>0.34</td>
<td>2.09</td>
<td>1.04</td>
<td>+1.38</td>
</tr>
<tr>
<td>Male 25-44</td>
<td>n.s.</td>
<td>n.s.</td>
<td>-</td>
<td>-</td>
<td>-</td>
<td>+1.51</td>
</tr>
<tr>
<td>Male 45-</td>
<td>$\frac{\delta u_i}{\delta p_i}$</td>
<td>0.93</td>
<td>3.76</td>
<td>0.82</td>
<td>0.80</td>
<td>n.s.</td>
</tr>
<tr>
<td>Female 15-19</td>
<td>n.s.</td>
<td>n.s.</td>
<td>2.20</td>
<td>1.56</td>
<td>1.81</td>
<td>0.74</td>
</tr>
<tr>
<td>Female 20-24</td>
<td>n.s.</td>
<td>n.s.</td>
<td>2.20</td>
<td>1.56</td>
<td>1.81</td>
<td>0.74</td>
</tr>
<tr>
<td>Female 25-44</td>
<td>n.s.</td>
<td>n.s.</td>
<td>0.63</td>
<td>0.56</td>
<td>0.63</td>
<td>0.56</td>
</tr>
<tr>
<td>Female 45-</td>
<td>n.s.</td>
<td>n.s.</td>
<td>0.43</td>
<td>0.50</td>
<td>0.43</td>
<td>0.50</td>
</tr>
</tbody>
</table>

n.s. : not significant
These results are broadly in line with our maintained hypothesis: it was suggested in Chapter 3.3 that the youngest and oldest male groups' unemployment-rates would be more sensitive to relative cohort size than would those of middle-age groups. It was also predicted in Chapter 3.3 that female unemployment-rates would be relatively insensitive to cohort size because of ease of substitutability between workers of different ages in typical female jobs. In the event, this difference was evident for older females but not for teenagers.

Here it may be of interest also to discuss the responsiveness of age-specific unemployment-rates to the rate of unemployment of prime-males. Table 4.5 presents the striking result that, for all females and (to a lesser extent) for the youngest and oldest males, the elasticity of group unemployment with respect to prime-male unemployment is significantly less than one. In other words, unemployment for groups comprised mainly of secondary workers is less sensitive to variations in the trade-cycle than is the unemployment-rate for prime workers. One may only speculate about the cause of this phenomenon. It may, for example, be that the lower rate of unionisation among secondary workers makes for more flexibility in wages - such flexibility in the price of labour would lessen variations in the quantity hired. Another suggestion, by Wachter (1977), is that young workers (like women) are abnormally concentrated in the service sector, which is less cyclical in its level of activity than is manufacturing.

Finally, it will be noted that Table 4-5 attributes quite strong unemployment increases to the introduction of the new unemployment-insurance system in 1971. With a dummy representing this system set
equal to one, the unemployment-rate for three of our demographic groups is raised in each case by more than a full percentage point.

5. Assessment of the Augmented Rates Effect

The previous sections have yielded some indication of the importance of demographic change in each age-sex group. In this section we pull the evidence together to obtain an estimate of the total impact of age-structure change on the rate of unemployment in Canada. The derivation of the estimate begins in Table 4-6. This table presents information for each group for which the coefficient on the demographic variable was significant in the regression equations. Take, as an illustration of the method used, the group consisting of males aged 15-19 years. The first column tells us that a rise of one-percentage point in this group's population-proportion was expected to raise its rate of unemployment by 1.10 percentage-points. The next four columns give the mean values of the group's population-proportion across each of our four Canadian trade-cycles. It will be noted, for example, that 15-19 year old males increased in relative numbers by 1.298 percentage points between the first two cycles. Multiplying this change in \( p_i \) by our estimate of the value of \( \frac{\partial u}{\partial p_i} \) gives us an estimate of the change in group unemployment-rate attributable to a change in group-size. In this case, the change in group-size raised the male teenage unemployment-rate by 1.43 percentage points between the 1954-8 and 1961-8 cycles. Next, the implication of this for the overall unemployment-rate in Canada was obtained by multiplying the figure of 1.43 by the mean proportion of the labour-force taken up by 15-19 year old males over the
### Table 4-6

The Augmented Rates Effect of Age-Structure Over Four Canadian Trade-Cycles

<table>
<thead>
<tr>
<th>Group</th>
<th>$\frac{\delta u_i}{\delta p_i}$</th>
<th>mean values of $p_i$</th>
<th>change in $u_i$ attributed to change in $p_i$</th>
</tr>
</thead>
<tbody>
<tr>
<td>male</td>
<td>15-19 1.10</td>
<td>5.896 7.194 7.628 7.567</td>
<td>+1.43 +0.48 -0.07 +1.48 +0.41</td>
</tr>
<tr>
<td>male</td>
<td>20-24 0.44</td>
<td>5.678 5.701 6.562 6.720</td>
<td>+0.02 +0.38 +0.07 +0.47 +0.45</td>
</tr>
<tr>
<td>male</td>
<td>45-   0.93</td>
<td>16.457 16.502 16.252 15.874</td>
<td>+0.04 -0.23 -0.35 -0.57 -0.58</td>
</tr>
<tr>
<td>female</td>
<td>15-19 2.20</td>
<td>5.756 7.004 7.398 7.298</td>
<td>+2.75 +0.87 -0.22 +3.39 +0.65</td>
</tr>
</tbody>
</table>
Table 4-6 (continued)

The Augmented Rates Effect of Age-Structure Over Four Canadian Trade-Cycles

<table>
<thead>
<tr>
<th>Group</th>
<th>( \frac{\delta u_i}{\delta p_i} )</th>
<th>( \delta u_i )</th>
<th>( \delta p_i )</th>
<th>1954-58</th>
<th>1961-68</th>
<th>1968-72</th>
<th>1972-78</th>
<th>1954-58</th>
<th>1961-68</th>
</tr>
</thead>
<tbody>
<tr>
<td>male 15-19</td>
<td>1.10</td>
<td>+0.08</td>
<td>+0.03</td>
<td>-0.004</td>
<td>+0.11</td>
<td>+0.02</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>male 20-24</td>
<td>0.44</td>
<td>0.00</td>
<td>+0.03</td>
<td>+0.006</td>
<td>+0.04</td>
<td>+0.04</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>male 45-</td>
<td>0.93</td>
<td>+0.01</td>
<td>-0.06</td>
<td>-0.07</td>
<td>-0.13</td>
<td>-0.13</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>female 15-19</td>
<td>2.20</td>
<td>+0.13</td>
<td>+0.04</td>
<td>-0.01</td>
<td>+0.16</td>
<td>+0.03</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Total Augmented Rates Effect | +0.22 | +0.04 | -0.08 | +0.18 | -0.04 |
combined periods 1954-8 and 1961-8. This exercise yields the number of 0.08 in the eleventh column of the table — this then is our estimate of the impact on overall unemployment arising from a 'cohort crowding' effect amongst teenage males. An identical procedure was followed for other demographic groups and for comparisons as between other pairs of Canadian trade-cycles. The total augmented rates effects shown in the table were obtained by aggregation across age-groups.

It will be noted from the results in Table 4-6 that the augmented rates effects had a rather negligible impact on the overall rate of unemployment. Their influence was, however, non-trivial on the unemployment-rates for particular age-groups; teenage girls, for example, suffered significantly higher unemployment in 1961-68 as a result of the influx of new labour market entrants. Older males, on the other hand, benefited from their fall in relative numbers and this partially offset the adverse impact on overall unemployment coming from the rise in the teenage rate. In other words, rates effects were more important in re-distributing the burden of unemployment than in increasing the overall rate.

In this paper, the overall impact of demographic change on the unemployment-rate in Canada has been viewed as the sum of weights effects and rates effects. The former were estimated in Chapter II, and in Table 4-7 these are added to the augmented rates effects estimated in the present chapter. These additions yield the 'total effects' shown in the table. Between 1954-8 and 1961-8, demographic change raised the unemployment-rate by 0.2 percentage points (75% of the overall increase between the two periods). Between 1961-8 and 1968-72, demographic
<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Augmented Rates Effect</td>
<td>+0.22</td>
<td>+0.04</td>
<td>-0.08</td>
<td>+0.18</td>
<td>-0.04</td>
</tr>
<tr>
<td>Pure Weights Effect</td>
<td>+0.02</td>
<td>+0.06</td>
<td>+0.09</td>
<td>+0.12</td>
<td>+0.11</td>
</tr>
<tr>
<td>Total Effect</td>
<td>+0.24</td>
<td>+0.10</td>
<td>+0.01</td>
<td>+0.30</td>
<td>+0.07</td>
</tr>
<tr>
<td>(% of total increase in U)</td>
<td>(75.0%)</td>
<td>(12.5%)</td>
<td>(0.01%)</td>
<td>(11.8%)</td>
<td>(8.75%)</td>
</tr>
</tbody>
</table>
change raised the unemployment-rate by 0.1 percentage points (12.5% of the overall increase between the two periods). Between 1968–72 and 1972–78, demographic change raised the overall unemployment-rate virtually not at all.

Thus, demographic change appears to explain most of the modest increase in unemployment between the nineteen-fifties and the nineteen-sixties; in the late nineteen-sixties, it was still playing some part in raising unemployment; but the evidence attributed no significant role to it in the nineteen-seventies when the sharpest rise in unemployment occurred. This negative conclusion with regard to the nineteen-seventies contrasts with our finding that unemployment-insurance changes induced a sharp rise in unemployment. The effect of unemployment-insurance changes was estimated using a procedure exactly analogous to that used for estimating weights effects. The procedure of taking estimated regression coefficients on the U.I. dummy for each group for which they were significant, and combining these with the appropriate labour-force weights, yielded an estimate that the 1971 Unemployment Insurance Act raised the Canadian unemployment-rate by 0.80 percentage points. ¹

The results indicate that one must be very sceptical about claims that the entry into the labour-force of unusually large cohorts of teenagers was responsible for large changes in Canadian unemployment in the nineteen-seventies. Examining the possible channels by which

¹. This estimate is broadly similar to that of other studies of the effect of the 1971 Act – see Chapter 1.3.
this could have occurred, we found no evidence for the hypothesis. The impact of this demographic change impinged more upon the structure of unemployment than upon the overall rate of unemployment.

6. Results Using Disaggregated Population Variables: A Supplementary Analysis

While the results presented above are those chosen as the basis of our assessment of the effects of demographic change, it was thought interesting to include in addition a section tentatively examining the subject in the context of somewhat more disaggregated population variables. In previous sections, the population variable in an equation to explain the unemployment-rate of a group \( i \) was based on the size of that group relative to the rest of the working-age population. The implicit assumption was that members of group \( i \) constituted one factor of production; other age-groups (combined) constituted a second factor of production. Given less than perfect substitutability between the two factors, and less that perfect wage-flexibility, the unemployment-rate of one group then depended partly on its size relative to the other group.

The hypothesis examined here is that the degree of substitutability, and the degree of relative-wage flexibility, will be different on the one hand between group \( i \) and some age-group \( j \), and, on the other hand, between group \( i \) and a third age-group \( k \). Little is known about the elasticities of substitution between age-groups but we try here to at least attempt some measurement of the problem.

In principle the unemployment-rate of group \( i \) should depend on its
size relative to each other group (as well as on other relevant independent variables). In estimating such an equation, the inclusion of so many variables would be a problem because of the loss of degrees of freedom. We therefore take the more moderate course of specifying only two population variables - population of i relative to younger age groups (of the same sex) and population of i relative to older age groups (of the same sex). To make the results stand comparison with those in previous sections, we include all the same additional explanatory variables as in our estimating equations above; we also define our new disaggregated population variables in the familiar log-of-odds form. Thus, for each age-group i, we now have two variables

\[ \text{RELPPOP}_i A = \log \frac{\text{odds of being a member of group } i}{\text{odds of being a member of a younger age group}} \]

\[ \text{RELPPOP}_i B = \log \frac{\text{odds of being a member of group } i}{\text{odds of being a member of an older age group}} \]

For the youngest age-group (15-19 years), the comparison in variable A is with 20-24 year olds and in B with 25-64 year olds. For the oldest age-group, the comparison in variable A is with 15-24 year olds and in B with 25-44 year olds.

The results of the exercise are presented in Table 4-8. In the case of the results for males, a correction for second-order autocorrelation was made in the equation for \( y_3 \); the procedure was identical to that described in Section 3 of this chapter, with \( RHO 1 = .67 \) and \( RHO 2 = -.09 \).

In all cases where the population variable had a coefficient not significantly different from zero, the new population variables carry
Table 4-8
Regression Results with Disaggregated Population Variables
(a)

Female Unemployment - Estimation by SURE

<table>
<thead>
<tr>
<th>dependent variable</th>
<th>constant</th>
<th>RELPOP A</th>
<th>RELPOP B</th>
<th>y3</th>
<th>wf</th>
<th>d</th>
<th>ef</th>
<th>R²</th>
<th>D.W.</th>
</tr>
</thead>
<tbody>
<tr>
<td>y5</td>
<td>1.54</td>
<td>0.007</td>
<td>1.17</td>
<td>0.80*</td>
<td>0.85</td>
<td>-0.07</td>
<td>0.44*</td>
<td>.97</td>
<td>1.81</td>
</tr>
<tr>
<td></td>
<td>(1.00)</td>
<td>(0.17)</td>
<td>(2.75)</td>
<td>(11.90)</td>
<td>(0.70)</td>
<td>(-0.69)</td>
<td>(2.95)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>y6</td>
<td>-2.35*</td>
<td>-0.007</td>
<td>0.02</td>
<td>0.73*</td>
<td>0.86</td>
<td>0.15*</td>
<td>0.71*</td>
<td>.98</td>
<td>1.94</td>
</tr>
<tr>
<td></td>
<td>(-7.44)</td>
<td>(-0.12)</td>
<td>(0.42)</td>
<td>(12.69)</td>
<td>(1.26)</td>
<td>(1.93)</td>
<td>(6.42)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>y7</td>
<td>-2.30*</td>
<td>-0.08</td>
<td>-0.01</td>
<td>0.54*</td>
<td>-0.06</td>
<td>0.28*</td>
<td>0.47*</td>
<td>.97</td>
<td>2.09</td>
</tr>
<tr>
<td></td>
<td>(-3.40)</td>
<td>(-1.02)</td>
<td>(-0.03)</td>
<td>(9.73)</td>
<td>(-0.06)</td>
<td>(3.49)</td>
<td>(3.87)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>y8</td>
<td>-4.13</td>
<td>-0.21</td>
<td>-0.80</td>
<td>0.44*</td>
<td>2.68</td>
<td>0.009</td>
<td>0.47*</td>
<td>.93</td>
<td>2.57</td>
</tr>
<tr>
<td></td>
<td>(-3.70)</td>
<td>(-1.50)</td>
<td>(-1.06)</td>
<td>(4.83)</td>
<td>(1.62)</td>
<td>(0.07)</td>
<td>(2.35)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>
Table 4-8 (continued)
(b)
Male Unemployment - Estimation by 3SLS with
Correction for Autocorrelation

<table>
<thead>
<tr>
<th>dependent variable</th>
<th>constant</th>
<th>RELPOP A</th>
<th>RELPOP B</th>
<th>$y_3$</th>
<th>g</th>
<th>wm</th>
<th>d</th>
<th>$y_3$</th>
<th>$R^2$</th>
<th>D.W.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$y_1$</td>
<td>2.26*</td>
<td>-0.20*</td>
<td>0.61*</td>
<td>0.91*</td>
<td>0.25</td>
<td>0.05</td>
<td></td>
<td></td>
<td>0.98</td>
<td>2.28</td>
</tr>
<tr>
<td></td>
<td>(5.11)</td>
<td>(-1.64)</td>
<td>(3.52)</td>
<td>(31.00)</td>
<td>(0.97)</td>
<td>(-1.93)</td>
<td></td>
<td></td>
<td>0.98</td>
<td>2.28</td>
</tr>
<tr>
<td>$y_2$</td>
<td>0.76</td>
<td>0.50*</td>
<td>-0.03</td>
<td>1.11*</td>
<td>6.74*</td>
<td>0.16*</td>
<td></td>
<td></td>
<td>0.99</td>
<td>2.55</td>
</tr>
<tr>
<td></td>
<td>(1.34)</td>
<td>(2.42)</td>
<td>(-0.15)</td>
<td>(29.08)</td>
<td>(2.26)</td>
<td>(4.94)</td>
<td></td>
<td></td>
<td>0.99</td>
<td>2.55</td>
</tr>
<tr>
<td>$y_3$</td>
<td>-1.96*</td>
<td>-0.07</td>
<td>-0.87*</td>
<td>0.96*</td>
<td>1.02*</td>
<td>0.41*</td>
<td>0.34</td>
<td></td>
<td>0.65</td>
<td>1.95</td>
</tr>
<tr>
<td></td>
<td>(-5.56)</td>
<td>(-0.08)</td>
<td>(-0.34)</td>
<td>(6.76)</td>
<td>(-1.67)</td>
<td>(2.06)</td>
<td>(1.83)</td>
<td></td>
<td>0.65</td>
<td>1.95</td>
</tr>
<tr>
<td>$y_4$</td>
<td>-0.06</td>
<td>0.58*</td>
<td>1.79*</td>
<td>0.79*</td>
<td>-0.40</td>
<td>-0.11*</td>
<td></td>
<td></td>
<td>0.94</td>
<td>1.87</td>
</tr>
<tr>
<td></td>
<td>(-0.25)</td>
<td>(1.81)</td>
<td>(3.93)</td>
<td>(18.61)</td>
<td>(-1.15)</td>
<td>(-2.49)</td>
<td></td>
<td></td>
<td>0.94</td>
<td>1.87</td>
</tr>
</tbody>
</table>

Note: the equation for $y_3$ incorporates a correction for autocorrelation (see text).

Symbols are explained in the text
$t$-statistics appear in brackets
asterisks indicate statistical significance at the 5% level.
just the same zero coefficients. For the other groups, the following principal points emerge:

(a) For teenagers of either sex, it was their population relative to those aged 25+ that was shown to be important: their unemployment-rate was not dependent on their population relative to the adjacent 20-24 age-group. A possible interpretation is that RELPOP B is proxying the old population variable, p2, (the two are very highly correlated because 20-24 is a relatively small age-group). The teenage population would then be seen as competing with all other workers with fine age-gradations unimportant - "15-19 versus the rest" might then be echoing a "trainees versus trained" dichotomy in patterns of substitutability between groups.

(b) For 20-24 year old males, their cohort-size only affected their unemployment-rate when it was measured relative to the size of the 15-19 cohort (not when it was measured relative to the 25+ group).

(c) For males, 45 and above, the coefficient RELPOP A and RELPOP B were each significant (on a 5% one-tailed test), though the more crucial variable appeared to be B (population size relative to the size of the adjacent group, 25-44).
Chapter V

The Effects of Changes in Age-Structure on

Unemployment in the United States, 1948-78

1. Introduction

Although the chief empirical focus of this study is the case of Canada, it was felt that an analysis of another country would be useful as supplementary evidence on the effects of demographic change. The United Kingdom seemed an interesting choice because its 'baby boom' had a rather different time-profile from those that occurred in North America. However, adequate time-series of age-specific unemployment-rates were not available for Britain. The United States was eventually selected for further study solely on the grounds of there existing an adequate data base there. Even so, for reasons outlined below, the analysis had to be less detailed than for Canada though the conclusions nevertheless seem fairly robust.

2. Decomposition of Changes in the Unemployment-Rate in the U.S.A.

In calculating the components of changes in the American unemployment-rate, the elements of change were defined in the same way as for Canada. Thus, changes were broken down into

i) an element of change unambiguously attributable to changes in labour-force weights \[ \sum_{i} u_{i}^{0} (w_{i}^{1} - w_{i}^{0}) \];

ii) an element of change unambiguously attributable to changes in age-specific unemployment-rates \[ \sum_{i} w_{i}^{0}(u_{i}^{1} - u_{i}^{0}) \]; and
iii) an element measuring the interaction or covariance between labour-force weights and age-specific unemployment-rates

\[ \sum_i (w_i^1 - w_i^0) (u_i^1 - u_i^0) \].

Here again, the \( u_i \) refer to age-specific unemployment-rates and the \( w_i \) to proportions of the labour-force accounted for by each group \( i \). Superscripts 0 and 1 distinguish between the base time-period and the terminal time-period.

This exercise of decomposition involved dividing the labour-force into groups according to age. The categories differ slightly from those made for Canada; the differences were dictated by the way in which the official labour-force statistics were presented. There were five age-groups delineated for the U.S.A., these being:

- 16-19 years
- 20-24 years
- 25-44 years
- 45-64 years
- 65 years and above.

The other preliminary task of data-organisation was to select time-periods over which comparisons of unemployment-rates could be made. Unemployment-peaks were less clear-cut in the American than in the Canadian case but the following three 'trade-cycles' were eventually chosen for study:

- 1954 - 1958
- 1963 - 1971
- 1971 - 1975
Each of these periods begins and ends with a year in which the American unemployment-rate was higher than in both the preceding and the following year. Table 3-1 shows the mean unemployment-rate in each of these three trade-cycles.

The results of the decomposition of changes in the overall unemployment-rate between each pair of trade-cycles is presented in Table 5-2. As in the Canadian case, we take the figures in the column headed \( \sum_{i} u_i^{10} (w_i^1 - w_i^0) \) as estimates of our pure weights effects.

Table 5.3 puts the size of these weights effects in the context of the total changes in unemployment-rate experienced by the United States as between each of our three periods. Between 1954-58 and 1963-71, the mean unemployment-rate fell by 0.46 percentage-points; without the pure weights effect, the fall would have been 0.24 percentage points greater (i.e. approximately half as great again). Between 1963-71 and 1971-75, the mean unemployment-rate increased by 1.54 percentage points; the weights-effect accounted for 15% of this increase. Although differences in the delineation of time-periods and age-groups make an exact comparison with Canada impossible, it is nevertheless plain from these figures that demographic change, working via the weights-effect channel, was pushing the unemployment-rate upwards more strongly in the United States than in Canada.

1. As in the corresponding table for Canada (Table 2-2), columns (1) and (6) differ slightly because of rounding errors in the calculation of the three components of change.
Table 5-1

Mean Unemployment-Rates Across Three American Trade-Cycles

<table>
<thead>
<tr>
<th>Years</th>
<th>Percent</th>
</tr>
</thead>
<tbody>
<tr>
<td>1954-58</td>
<td>5.02</td>
</tr>
<tr>
<td>1963-71</td>
<td>4.56</td>
</tr>
<tr>
<td>1971-75</td>
<td>6.10</td>
</tr>
</tbody>
</table>

Table 5-2

Decomposition of Changes in the Unemployment-Rate Across Three American Trade-Cycles

(percentage points)

<table>
<thead>
<tr>
<th>Trade Cycles</th>
<th>Total Change (\sum u_i^0(w_i - w_i^0))</th>
<th>(\sum w_i^0(u_i^1 - u_i^0))</th>
<th>(\sum (u_i^1 - u_i^0)(w_i^1 - w_i^0))</th>
<th>Sum of Columns</th>
</tr>
</thead>
<tbody>
<tr>
<td>1954-8 to</td>
<td>-0.46</td>
<td>+0.24</td>
<td>-0.75</td>
<td>+0.04</td>
</tr>
<tr>
<td>1963-71</td>
<td>+1.08</td>
<td>+0.40</td>
<td>+0.38</td>
<td>+0.19</td>
</tr>
<tr>
<td>1971-75</td>
<td>+1.54</td>
<td>+0.20</td>
<td>+1.19</td>
<td>+0.06</td>
</tr>
<tr>
<td>1963-71 to</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1971-75</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Period</td>
<td>Pure Weights Effect</td>
<td>Pure Weights Effect as a Percentage of the Total Change in Unemployment</td>
<td></td>
<td></td>
</tr>
<tr>
<td>-----------</td>
<td>---------------------</td>
<td>------------------------------------------------------------------------</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1954-58</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1963-71</td>
<td>+0.24</td>
<td>-51.1</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1971-75</td>
<td>+0.40</td>
<td>+41.2</td>
<td></td>
<td></td>
</tr>
<tr>
<td>1963 to</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>1971-75</td>
<td>+0.20</td>
<td>+13.8</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>
There were two reasons for the weights-effect being more important in the United States than in Canada. First, the increase in the proportion of the labour-force accounted for by the unemployment-prone teenage group was larger in the American case. Secondly, this teenage-group whose weight in the labour-force was increasing was relatively more prone to unemployment in the United States than in Canada. Comparing the two countries over 1953-78, the mean unemployment-rate for the 25-44 age group was almost exactly the same in each case (4.17% in the U.S.A., 4.13% in Canada). However, for teenagers, the American mean unemployment-rate was much higher than that for Canada (14.65% compared with 11.43%), the difference arising mainly from higher unemployment amongst teenage girls in the U.S.A. ¹ A given increase in the share of teenagers in the labour-force would therefore increase the overall unemployment-rate by a greater amount in the United States than in Canada.

3. Estimating the Augmented Rates Effect of Age-Structure Change in the U.S.A.

In estimating augmented rates-effects (as defined in Chapter III), regression analysis was utilised to examine the effect of changes in group-size on unemployment-rates for eight groups delineated by age and sex; the age-groups were 16-19 years, 20-24 years, 25-44 years and 45-64 years. No analysis was carried out for the 65 years- and-

¹ Mean unemployment-rates for teenagers over 1953-78 were: males, 14.59% in the U.S.A., 13.03% in Canada; females, 14.72% in U.S.A., 9.55% in Canada.
above group whose members in 1978 accounted for less than 3% of the labour-force; with a group unemployment-rate of only 4%, the 65 years-
and-above group added only one-tenth of one percentage-point to the national unemployment-rate and variations in their group unemployment-rate would therefore have been of no importance to the overall unemployment figures.

The general reduced-form equations\footnote{The reduced-forms can be derived in an exactly analogous fashion to the derivation for Canada – see Chapter 3.5 and Appendix III.} tested in the analysis, were of the form

\[ u_i = f(p_i, z_i). \]

Here, \( u_i \) is the unemployment-rate for group \( i \), \( p_i \), included to test for the presence of the 'cohort crowding' effect, is the proportion of the 16-64 year old population accounted for by group \( i \); \( z_i \) is a vector of other variables thought to influence \( u_i \). To facilitate an inter-country comparison it was desirable that the variables included in \( z_i \) were similar to those included in the Canadian equations of Chapter IV; exact similarity was not possible however.

Choice of Cyclical Variable

As in the case of Canada, the equation used to explain the prime-male unemployment-rate includes a cyclical variable called 'G.N.P. gap'. This variable, \( g \), was generated by taking a series 'G.N.P. measured in 1972 dollars' and fitting a log-linear trend to the thirty-one years of data (1948-78); residuals were then retrieved...
and used as the G.N.P gap variable (with positive values representing shortfalls of G.N.P from trend and negative values representing real G.N.P. figures above the fitted trend).

Again, following the practices of Chapter III, the unemployment equations for other age-groups included the rate of unemployment for prime-males (25-44 years) as the explanatory cyclical variable.

Choice of Minimum-Wage Variable

The minimum wage index used for the United States was similar to the indices used for Canada in the sense that it reflected a variety of wage minima with weights reflecting the proportion of the labour-force covered by each legislated minimum-rate. The difference from Canada is that for most of the period there were different minimum-wage rates in different industries; in Canada, variations in the minimum wage were principally across provinces rather than across industries as in the U.S.A. The American index for 1948-77 was kindly supplied by Prof. Peter Mattila of Iowa State University and updated to 1978 using figures in the Statistical Abstract of the United States. It was calculated from the formula

\[ E \left[ \frac{E_t}{E} \right] \left[ \left( \frac{MP}{AHE} \right) \left( OB \right) + \left( \frac{MN}{AHE} \right) \left( CN \right) \right] \]

where

- \( E = \) payroll employment,
- \( AHE = \) gross average hourly earnings
- \( MP = \) basic minimum wage weighted by number of months in effect
MN = minimum wage for "newly covered" workers (as defined in U.S. legislation)

CB = proportion of non-supervisory employees covered by the basic minimum

CN = proportion of non-supervisory employees covered by the rate applicable to "newly covered" workers

i = major industry division (wholesale and retail trade treated as separate divisions)

t = total non-farm economy.

The index was not available separately for each age-sex group and it has therefore to be assumed that variation in the index correspond adequately to variations in minimum wage conditions as faced by any individual demographic group.

Choice of Unemployment-Insurance Variable

For Canada, it was possible to represent variations in unemployment-insurance provision by using a dummy variable set equal to one for the period after the introduction of new legislation in 1971. This was possible because the great majority of the variation in both benefits and coverage was associated with the introduction of the 1971 Act. No single change of this magnitude occurred in the unemployment-insurance system of the United States and it was therefore necessary to construct a continuous series to represent unemployment-insurance conditions. Our unemployment-insurance index was equal to the retention-ratio multiplied by the coverage of unemployment-insurance:
Unemployment-Insurance Index = \frac{\text{average weekly benefit}}{\text{average weekly earnings}} \times \frac{\text{U.I. covered employment}}{\text{total employment}}

The upward trend in this unemployment-insurance index owed much more to the extension of coverage than to changes in the retention-ratio. In the results reported here, the unemployment insurance index is used as a variable in the regression equations; work was also done with coverage and retention-ratios entered separately but the results of this exercise were not very different from those reported here.

The use of an unemployment-insurance index based on aggregate data is unfortunate in our context of trying to explain unemployment-rates for individual sub-groups of the population. As an illustration, suppose the index rose sharply because unemployment-insurance schemes were extended to married women. This rise in the index would have little relevance to the unemployment position of teenage boys and the 'fit' of any equation using the index to 'explain' teenage male unemployment would worsen as a result. The problem with using the index is thus a serious one but, unfortunately, more disaggregated data on unemployment-insurance were not available.

Choice of Male and Female Earnings Variables

The same problem arose as in the Canadian case - consistent time-series were not available for hourly earnings by sex for the whole period of the analysis. However, there were figures for average hourly earnings by industry\(^1\) and (having been first deflated by the 1971 Con-

---

sumer Price Index), these were used in the generation of the series, em and ef.

The proxy taken for male earnings potential was the series on average hourly earnings of non-supervisory workers in private manufacturing. In 1978, women still only accounted for 30.4% of the labour-force in U.S. manufacturing, so that male earnings are being proxied by average earnings in a male-dominated industry. Similar reasoning indicated the choice of average hourly earnings in wholesale and retail trade as a proxy for female earnings potential - this was one of the most female-intensive of the various industry groups with women accounting for 43.7% of its employees in 1978.

The two series, em and ef, were highly correlated. Therefore, each is used without the other, in the male and female equations respectively.

4. Unemployment: Specification of the Model for the U.S.A.

Initially, estimates were made of systems of equations with the following form (in the case of males)

\[ y_i = f_i (x_i, y_3, w, u_i, em) \quad i = 1, 2, 4, \]

\[ y_i = f_i (x_i, g, w, u_i, em) \quad i = 3, \]

where

\[ y_i = \text{log of odds of unemployment for group } i, \]

\[ x_i = \text{log of odds of being a member of group } i, \]

\[ g = \text{G.N.P. gap,} \]

\[ w = \text{log of the minimum wage index,} \]

\[ u_i = \text{log of the unemployment-insurance index,} \]

\[ em = \text{male average hourly earnings series.} \]
i = 1: male 16-19
i = 2: male 20-24
i = 3: male 25-44
i = 4: male 45-64.

Systems such as these were chosen because of their close comparability with those used in the analysis of unemployment-rates in Canada. Results for the various specifications tried were however unsatisfactory. The coefficient on the unemployment-insurance variable was frequently of the "wrong" (i.e. negative) sign; that on the minimum wage variable was seldom significantly different from zero; and autocorrelation was always a serious problem. Table 5-4 presents a sample result from this exercise. Estimation was by three-stage least-squares and used annual data from 1948-78. It will be noted that the population variable \( x_1 \) appears with a coefficient significantly greater than zero in all four equations. Coefficients on the excess demand variable are similarly of the expected positive sign. However, the coefficients on the minimum-wage variable are nowhere significantly different from zero. For the unemployment-insurance variable, the coefficient is negative and significant (5% level) in the case of the youngest group (the "wrong" sign) and insignificantly different from zero in the other three equations. Autocorrelation is indicated as a serious problem in the second equation by virtue of a low Durbin-Watson statistic.

The results embodied in Table 5-4 are typical of those that were obtained from estimating several such systems with varying specifications. The principal problem was with the unemployment-insurance index; this problem was not resolved by entering unemployment-insurance
Table 5-4

Male Unemployment - Estimation by 3SLS

<table>
<thead>
<tr>
<th>dependent variable</th>
<th>constant</th>
<th>$x_1$</th>
<th>$y_3$</th>
<th>$g$</th>
<th>$w$</th>
<th>$u_i$</th>
<th>$e_m$</th>
<th>$R^2$</th>
<th>D.W.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$y_1$</td>
<td>2.87*</td>
<td>1.08*</td>
<td>0.66*</td>
<td>0.12</td>
<td>-0.41*</td>
<td>0.03</td>
<td>.95</td>
<td>1.69</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(3.12)</td>
<td>(5.08)</td>
<td>(17.51)</td>
<td>(1.12)</td>
<td>(-2.76)</td>
<td>(0.32)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_2$</td>
<td>2.01*</td>
<td>0.88*</td>
<td>1.03*</td>
<td>-0.09</td>
<td>-0.55</td>
<td>0.14</td>
<td>.91</td>
<td>0.87</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(2.24)</td>
<td>(2.97)</td>
<td>(13.46)</td>
<td>(-0.45)</td>
<td>(-1.39)</td>
<td>(1.27)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_3$</td>
<td>-2.87*</td>
<td>2.78*</td>
<td></td>
<td>10.08*</td>
<td>-0.02</td>
<td>-0.62</td>
<td>0.72*</td>
<td>.81</td>
<td>1.40</td>
</tr>
<tr>
<td></td>
<td>(-2.35)</td>
<td>(3.66)</td>
<td></td>
<td>(9.46)</td>
<td>(-0.04)</td>
<td>(-1.50)</td>
<td>(3.81)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_4$</td>
<td>3.49*</td>
<td>1.65*</td>
<td>0.89*</td>
<td>0.14</td>
<td>0.07</td>
<td>-0.31*</td>
<td>.96</td>
<td>2.14</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(1.84)</td>
<td>(2.20)</td>
<td>(22.28)</td>
<td>(1.29)</td>
<td>(0.03)</td>
<td>(-3.52)</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Symbols are explained in the text.
t-statistics appear in brackets.
asterisks indicate statistical significance at the 5% level.
coverage and the retention-ratio separately. Why the index should fail to explain any of the variation in unemployment is not clear. A possible explanation is that an aggregate unemployment-insurance index fails to reflect adequately variations in the unemployment-insurance conditions faced by individual demographic groups (a similar remark could be made with reference to the minimum-wage index). This possibility could of course be tested if disaggregated data were used but such data were not available to the author. It was therefore decided to provide a less detailed analysis of unemployment-rates in the United States and to estimate only simple equations of the sort used by Wachter and Kim (1979). Wachter and Kim "explained" teenage-unemployment using only an excess-demand variable and a population-proportion variable. We add to these only a male or female earnings series (included because of its impact on participation-rates). Further differences between their analysis and our's are that we extend the analysis to cover other age-groups and enter the variables in the logistic form explained in Chapter IV. The period of analysis here is 1948-78. As in the Canadian chapter, we first analyse regression results for the female demographic groups.

5. Female Unemployment: Regression Results for the U.S.A., 1948-78

The four female age-groups whose unemployment-rates were to be analysed were 16-19, 20-24, 25-44 and 45-64 years; these groups are denoted by subscripts 5, 6, 7 and 8 respectively. The model estimated consisted of the four equations

\[
\hat{y}_i = \alpha_i + \beta_1 x_1 + \beta_2 y_3 + \beta_3 \text{ef} \quad i = 5, 6, 7, 8.
\]

Here \(x, y_1\) and ef have exactly the same connotation as they were given in Chapter IV.
Again, as in Chapter IV, estimation of the model by single-equation ordinary-least-squares would have been inappropriate because of the likelihood that error terms across the four equations would be correlated. Estimation was therefore by the method of seemingly unrelated regression equations (SURE).

The results of estimation by SURE are given in Table 5-5. In the first equation, the Durbin-Watson statistic was just low enough that we had to reject (at the 95% confidence level) the hypothesis that there was zero autocorrelation. Some attempt was therefore made to correct for autocorrelation in estimating the system of equations. Unfortunately, a statistical package allowing correction for autocorrelation within the context of SURE was not available. As an ad hoc measure, the following procedure was therefore adopted. Each equation was estimated by a single-equation Cochrane-Orcutt procedure. RHO was statistically significant for the equations for \( y_1 \) and \( y_2 \). The values of RHO estimated for the equations for \( y_1 \) and \( y_2 \) were therefore retrieved. Each independent \((v_i)\) in the first and second equations was then normalised, using the relevant value of RHO, in the following manner:

\[
\begin{align*}
v_i^* &= v_i - (RHO \, v_i)_{t-1}
\end{align*}
\]

The resulting normalised equations were then re-estimated using the seemingly-unrelated-regression-equations method. The results are presented in Table 5-6 and form a basis for subsequent analysis in this chapter.

Comments on Female Regression Equations

As in the Canadian case, the results for females provide only limited support for the cohort-crowding hypothesis.
### Table 5-5

**Female Unemployment—Estimation by SURE**

<table>
<thead>
<tr>
<th>dependent variable</th>
<th>constant</th>
<th>$x_1$</th>
<th>$y_3$</th>
<th>$ef$</th>
<th>$R^2$</th>
<th>D.W.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$y_5$</td>
<td>0.34</td>
<td>0.59*</td>
<td>0.43*</td>
<td>0.38*</td>
<td>0.94</td>
<td>1.01</td>
</tr>
<tr>
<td></td>
<td>(0.38)</td>
<td>(2.59)</td>
<td>(11.10)</td>
<td>(4.03)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_6$</td>
<td>-0.90*</td>
<td>0.23*</td>
<td>0.61*</td>
<td>0.44*</td>
<td>0.96</td>
<td>1.34</td>
</tr>
<tr>
<td></td>
<td>(-2.62)</td>
<td>(2.42)</td>
<td>(20.42)</td>
<td>(13.47)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_7$</td>
<td>-1.79*</td>
<td>-0.56</td>
<td>0.61*</td>
<td>0.09</td>
<td>0.85</td>
<td>1.67</td>
</tr>
<tr>
<td></td>
<td>(-5.01)</td>
<td>(-1.24)</td>
<td>(12.28)</td>
<td>(0.84)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_8$</td>
<td>-2.06*</td>
<td>-0.43</td>
<td>0.60*</td>
<td>0.04</td>
<td>0.80</td>
<td>1.66</td>
</tr>
<tr>
<td></td>
<td>(-2.79)</td>
<td>(-1.11)</td>
<td>(10.70)</td>
<td>(0.74)</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

*Symbols are explained in the text. t-statistics appear in brackets. Asterisks indicate statistical significance at the 5% level.*
Table 5-6
Female Unemployment - Estimation by SURF with
Correction for Autocorrelation

<table>
<thead>
<tr>
<th>dependent variable</th>
<th>constant</th>
<th>$x_1$</th>
<th>$y_3$</th>
<th>$y_1^A$</th>
<th>ef</th>
<th>$R^2$</th>
<th>D.W.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$y_5$</td>
<td>0.99</td>
<td>0.73*</td>
<td>0.42*</td>
<td>-1.36*</td>
<td>0.26*</td>
<td>.94</td>
<td>1.32</td>
</tr>
<tr>
<td></td>
<td>(1.01)</td>
<td>(2.37)</td>
<td>(10.54)</td>
<td>(-2.67)</td>
<td>(1:99)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_6$</td>
<td>-0.72</td>
<td>0.21</td>
<td>0.60*</td>
<td>-1.17*</td>
<td>0.43*</td>
<td>.94</td>
<td>1.32</td>
</tr>
<tr>
<td></td>
<td>(-1.78)</td>
<td>(1.60)</td>
<td>(17.69)</td>
<td>(-4.41)</td>
<td>(7.85)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_7$</td>
<td>-2.04*</td>
<td>-1.00</td>
<td>0.61*</td>
<td>-0.03</td>
<td>.84</td>
<td>.88</td>
<td>1.78</td>
</tr>
<tr>
<td></td>
<td>(-5.46)</td>
<td>(-1.99)</td>
<td>(12.60)</td>
<td>(-0.22)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$y_8$</td>
<td>-2.14*</td>
<td>-0.48</td>
<td>0.59*</td>
<td>0.04</td>
<td>.79</td>
<td>1.64</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-2.79)</td>
<td>(-1.18)</td>
<td>(10.36)</td>
<td>(0.71)</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Note: In the $y_5$ equation, each independent variable was transformed by the formula

$$v_i^* = v_i - 0.15 v_{i-1} \quad \text{and} \quad y_5^A = -0.15 y_{5-1}$$

In the $y_6$ equation, each independent variable was transformed by the variable

$$v_i^* = v_i - 0.17 v_{i-1} \quad \text{and} \quad y_6^A = -0.17 y_{6-1}$$

(see text for explanation).

Symbols are explained in the text.

T-statistics appear in brackets.

Asterisks indicate statistical significance at the 5% level.
The coefficient on the demographic variable is significantly greater than zero (at the 1% level) for the teenage group. In the other equations, the coefficient on the demographic variable is not significantly different from zero on a two-tailed test (5% level of significance). This relative failure of the cohort-crowding hypothesis in the case of females could be considered disappointing. However, it may not prove out of line with our maintained hypothesis (from Chapter 3.3) that female unemployment-rates will be less sensitive to age-groups size than will male unemployment-rates.

As would be expected, the estimate of the coefficient on the cyclical variable, $y_3$, is positive and significant (5% level) in all four equations. The coefficient on the female earnings variable is positive and significant at the 5% level (one-tailed test) in the equations for the two youngest age-groups and insignificantly different from zero for the two eldest age-groups.

Further comment on the female results is postponed until after the male results have been presented.

6. Male Unemployment: Regression Results for the U.S.A., 1948-78

The four male groups (15-19, 20-24, 25-44 and 45 years and above) are here denoted by subscripts $i = 1, 2, 3$ and 4 respectively. The model estimated consisted of the four equations.

$$\hat{y}_1 = a + \beta_1 x_1 + \beta_2 y_3 + \beta_3 em$$

$$i = 1, 2, 3$$

$$\hat{y}_3 = a + \beta_1 x_3 + \beta_2 g + \beta_3 em$$
Estimation, as in the Canadian case, was by the method of Three-Stage
Least-Squares (3SLS).

The results of estimation by 3SLS are given in Table 5-7. Unfortunately, severe autocorrelation was indicated in the equation for $y_2$.
Tests indicated a coefficient of first-order autocorrelation not significantly different from one. The second equation was therefore couched in terms of first-differences and the following system estimated by 3SLS:

\[
\begin{align*}
\hat{y}_{1t} &= a + \beta_1 x_{1t} + \beta_2 y_{3t} + \beta_3 em_t \\
\hat{y}_{2t} &= a + \beta_1 (x_{2t-1} - x_{2t}) + \beta_2 (y_{3t-1} - y_{3t}) + \beta_3 (em_{t-1} - em_t) + \beta_4 y_{2t-1} \\
\hat{y}_{3t} &= a + \beta_1 x_{3t} + \beta_2 e_t + \beta_3 em_t \\
\hat{y}_{4t} &= a + \beta_1 x_{4t} + \beta_2 y_{3t} + \beta_3 em_t
\end{align*}
\]

The results of the exercise are tabulated in Table 5-8 and form a basis for subsequent analysis in this chapter.

Comments on Male Regression Results

As in the Canadian case, the results for male unemployment provide stronger support for the cohort-crowding hypothesis than do the results for female unemployment. The coefficient on the demographic variable, $x_1$, is significant (at the 5% level; one-tailed test) in three of the four equations. Thus, in each country, the variable was significant for the youngest and oldest groups; in Canada (but not the U.S.A.) it was significant for the 20-24 age-group; in the U.S.A. (but not Canada), it was significant for the prime-age group (though it would not in this case have been adjudged significant on a two-tailed test at the 5% level).
Table 5-7

Male Unemployment - Estimation by 3SLS

<table>
<thead>
<tr>
<th>dependent variable</th>
<th>constant</th>
<th>(x_1)</th>
<th>(y_3)</th>
<th>(g)</th>
<th>(e)m</th>
<th>(R^2)</th>
<th>D.W.</th>
</tr>
</thead>
<tbody>
<tr>
<td>(y_1)</td>
<td>2.56*</td>
<td>0.85*</td>
<td>0.62*</td>
<td>0.04</td>
<td>.94</td>
<td>.94</td>
<td>1.30</td>
</tr>
<tr>
<td></td>
<td>(2.95)</td>
<td>(3.88)</td>
<td>(17.86)</td>
<td>(0.51)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(y_2)</td>
<td>2.27*</td>
<td>0.53*</td>
<td>1.00*</td>
<td>0.02</td>
<td>.90</td>
<td>.90</td>
<td>0.67</td>
</tr>
<tr>
<td></td>
<td>(3.23)</td>
<td>(2.76)</td>
<td>(16.24)</td>
<td>(0.40)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(y_3)</td>
<td>-1.50*</td>
<td>2.59*</td>
<td>9.41*</td>
<td>0.52*</td>
<td>.81</td>
<td>.81</td>
<td>1.37</td>
</tr>
<tr>
<td></td>
<td>(-2.18)</td>
<td>(3.32)</td>
<td>(10.30)</td>
<td>(3.81)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(y_4)</td>
<td>2.68*</td>
<td>1.47*</td>
<td>0.88*</td>
<td>-0.25*</td>
<td>.97</td>
<td>.97</td>
<td>1.96</td>
</tr>
<tr>
<td></td>
<td>(4.64)</td>
<td>(4.31)</td>
<td>(25.81)</td>
<td>(-10.12)</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Symbols are explained in the text.

t-statistics appear in brackets.

Asterisks indicate statistical significance at the 5% level.
Table 5-8

Male Unemployment - Estimates by 3SLS with Correction for Autocorrelation in the \( y_2 \) Equation

<table>
<thead>
<tr>
<th>dependent variable</th>
<th>constant</th>
<th>( x_1 )</th>
<th>( y_3 )</th>
<th>( g )</th>
<th>( e_m )</th>
<th>( y_2^A )</th>
<th>( R^2 )</th>
<th>D.W.</th>
</tr>
</thead>
<tbody>
<tr>
<td>( y_1 )</td>
<td>2.27*</td>
<td>0.78*</td>
<td>0.61*</td>
<td>0.05</td>
<td></td>
<td>( y_2^A )</td>
<td>.94</td>
<td>1.33</td>
</tr>
<tr>
<td></td>
<td>(2.16)</td>
<td>(2.92)</td>
<td>(18.70)</td>
<td>(0.57)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( y_2 )</td>
<td>-0.42*</td>
<td>0.37</td>
<td>0.94*</td>
<td>-0.07</td>
<td>0.83*</td>
<td></td>
<td>.96</td>
<td>1.53</td>
</tr>
<tr>
<td></td>
<td>(-4.46)</td>
<td>(0.68)</td>
<td>(23.22)</td>
<td>(-0.29)</td>
<td>(21.89)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( y_3 )</td>
<td>-2.26*</td>
<td>1.57*</td>
<td>10.10*</td>
<td>0.33*</td>
<td></td>
<td></td>
<td>.84</td>
<td>1.46</td>
</tr>
<tr>
<td></td>
<td>(-3.19)</td>
<td>(1.87)</td>
<td>(11.62)</td>
<td>(2.16)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>( y_4 )</td>
<td>2.50*</td>
<td>1.32*</td>
<td>0.87*</td>
<td>-0.28*</td>
<td></td>
<td></td>
<td>.97</td>
<td>2.05</td>
</tr>
<tr>
<td></td>
<td>(4.76)</td>
<td>(4.17)</td>
<td>(28.13)</td>
<td>(-11.38)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Note: In the equation for \( y_2 \), all independent variables have been transformed to the first-difference form, \( v_1^* = v_i - v_{i-1} \)

\( y_2^A = y_{2-t} \)

Symbols are explained in the text.

\( t \)-statistics appear in brackets. Asterisks indicate statistical significance at the 5% level.
Further comparisons between the Canadian and American results will be made in the next section.

With respect to the other variables included in the model, it will be noted that the coefficient on the cyclical variable is always strongly significant. The male earnings series, $em$, has a positive sign in the $y_3$ equation and a negative sign in the $y_4$ equation. This last result is counter to our prediction that high earnings would raise unemployment via an increase in participation-rates. The explanation may be that another effect has been omitted from our model—the tendency of high earnings to reduce unemployment via an increase in the costs of search unemployment.

7. Interpretation of Male and Female Regression Results

The results presented in the last two sections included equations in the form

$$\log \frac{u_1}{1-u_1} = \alpha + \beta_1 \log \frac{p_1}{1-p_1} + \beta_2 \log \frac{u_3}{1-u_3} + \beta_3 em$$

As they stand, the coefficients in such an equation are hard to interpret. Ideally, we would like to have values for expressions such as $\frac{\delta u_1}{\delta p_1}$ and $\frac{\delta u_1}{\delta u_3}$. The way in which such expressions can be derived from our equations was explained in Chapter 4.4. The evaluation of these expressions for the American case is presented in Table 5-9, which also gives the corresponding elasticities. To facilitate inter-country comparison, we add also the appropriate elasticity-estimates derived from our analysis of the Canadian data.
Table 5-9
Interpretation of Regression Results

<table>
<thead>
<tr>
<th>U.S.A.</th>
<th>Population Variable</th>
<th>Cyclical Variable</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\frac{\delta u_i}{\delta p_i}$ elasticity of unemployment rate w.r.t. population proportion</td>
<td>$\frac{\delta u_i}{\delta \mu}$ elasticity of unemployment rate w.r.t. cyclical variable</td>
</tr>
<tr>
<td></td>
<td>male 16-19</td>
<td>1.98</td>
</tr>
<tr>
<td></td>
<td>male 20-24</td>
<td>n.s.</td>
</tr>
<tr>
<td></td>
<td>male 25-44</td>
<td>0.31</td>
</tr>
<tr>
<td></td>
<td>male 45-64</td>
<td>0.32</td>
</tr>
<tr>
<td></td>
<td>female 16-19</td>
<td>1.72</td>
</tr>
<tr>
<td></td>
<td>female 20-24</td>
<td>n.s.</td>
</tr>
<tr>
<td></td>
<td>female 25-44</td>
<td>n.s.</td>
</tr>
<tr>
<td></td>
<td>female 45-64</td>
<td>n.s.</td>
</tr>
</tbody>
</table>

n.s.: not significant

Comparative Elasticities for Canada (from Table 4-5)

<table>
<thead>
<tr>
<th>Group</th>
<th>elasticity of unemployment rate w.r.t. population proportion</th>
<th>elasticity of unemployment rate w.r.t. cyclical variable</th>
</tr>
</thead>
<tbody>
<tr>
<td>male 15-19</td>
<td>0.59</td>
<td>0.82</td>
</tr>
<tr>
<td>male 20-24</td>
<td>0.34</td>
<td>1.04</td>
</tr>
<tr>
<td>male 25-44</td>
<td>n.s.</td>
<td>-</td>
</tr>
<tr>
<td>male 45 and above</td>
<td>3.76</td>
<td>0.80</td>
</tr>
<tr>
<td>female 15-19</td>
<td>1.56</td>
<td>0.74</td>
</tr>
<tr>
<td>female 20-24</td>
<td>n.s.</td>
<td>0.74</td>
</tr>
<tr>
<td>female 25-44</td>
<td>n.s.</td>
<td>0.56</td>
</tr>
<tr>
<td>female 45 and above</td>
<td>n.s.</td>
<td>0.50</td>
</tr>
</tbody>
</table>
An exact comparison between the Canadian and American results is difficult because the empirical analysis has been less detailed for the latter country. However, the following points stand out:

i) In each country, the teenage group was the only female group for which the coefficient on the population variable was significant.

ii) The coefficient on the population variable was also significant in each country for the two male groups, teenagers and 45-and-above.

iii) For the three demographic groups just cited, the elasticity of the unemployment rate with respect to the population proportion was greatest for males, 45-and-above, in both Canada and the U.S.A. However, for the teenage groups, while in the U.S. girls and boys suffered similarly from a bulge in their population, in Canada the bulge impinged more severely on girls than on boys.

iv) A difference between the countries was that the coefficient on the population variable in the male, 20-24, group was significant only in the Canadian case. This was a surprising result.

v) Another difference was that the coefficient on the prime-male population variable was significant only in the American case. However, it 'passed' the significance test only narrowly (it would have been insignificantly different from zero if judged on a two-tailed rather than a one-tailed test (5% level)).

---

1. Table 5-7 shows that the coefficient on $x_2$ was significant - but this result was 'changed once a correction had been made for auto-correlation in the $y_2$ equation.'
Table 5-9 also enables us to compare the responsiveness of group unemployment-rates to cyclical factors. The relative magnitude of the elasticity measures across age-groups were similar as between the two countries to the extent that the degree of responsiveness of the group unemployment-rate to the prime-male unemployment-rate was generally greater for males than for females and less for the youngest male group than for the 20-24 male group.

In conclusion, the fact that we have had to omit some potentially important variables from the analysis of U.S. employment makes estimates of augmented rates effects appear tentative. However, the similarity of results between Canada and the U.S.A. is sufficient to strengthen our conclusion that demographic effects were playing some role in the American labour market.

8. Assessment of the Augmented Rates Effect

We now use the estimates of $\frac{\delta u}{\delta p}$ derived in the previous section to assess the overall effect of demographic change on the U.S. unemployment-rate. Table 5-10 begins the assessment by deriving estimates of the magnitude of the augmented rates effect as between each of our three periods, 1954-58, 1963-71 and 1971-75. The lay-out of this table follows that of Table 4-6 in Chapter IV and was explained in Section 5 of that chapter.

Table 5-10 indicates that the augmented rates effect played no role in determining changes in American unemployment-rates as between each of these three periods. The result is however sensitive with respect to the estimates of $\frac{\delta u}{\delta p}$ (the prime-male group). Had this been zero (as it
### Table 5-10

The Augmented Rates Effect of Age-Structure Change Over Three American Trade-Cycles

<table>
<thead>
<tr>
<th>Group</th>
<th>$\frac{\delta u_1}{\delta p_1}$</th>
<th>Mean values of $p_1$</th>
<th>Change in $u_1$ attributed to change in $p_1$</th>
<th>Effect on overall $U$</th>
</tr>
</thead>
<tbody>
<tr>
<td>male 16-19</td>
<td>1.98</td>
<td>4.57</td>
<td>6.07</td>
<td>6.37</td>
</tr>
<tr>
<td>male 25-44</td>
<td>0.31</td>
<td>22.82</td>
<td>20.01</td>
<td>19.62</td>
</tr>
<tr>
<td>male 45-64</td>
<td>0.32</td>
<td>16.44</td>
<td>16.49</td>
<td>15.91</td>
</tr>
<tr>
<td>female 16-19</td>
<td>1.72</td>
<td>4.46</td>
<td>5.94</td>
<td>6.21</td>
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<tr>
<td>Augmented Rates Effect</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>
would have been with the application of a two-tailed test), the augmente
ted rates effect would have been indicated as having added one-quarter 
of a percentage-point to the unemployment-rate between 1954-8 and 1963-
71. This would have been the same estimate as for the augmented rates 

As in the Canadian case, a small estimate of the rates effect on 
overall unemployment disguises quite large effects on individual groups. 
Teenagers suffered particularly from the effects of demographic change; 
over three percentage-points was added to the unemployment-rates of both 
sexes between 1954-8 and 1971-5. Older males experienced lower unemploy-
ment as a result of demographic change; although the fall in their un-
employment-rate was relatively small, these groups were large enough 
that the net rates effect of demographic change on overall unemployment 
was neutral.

Finally, Table 5-11 adds together the pure weights and augmented 
rates effects estimated for the United States. Between 1954-58 and 
1971-75, the mean U.S. unemployment-rate increased by 1.08 percentage 
points. Demographic change is estimated to have accounted for 31.5% 
of this increase. 1

Our conclusion is that demographic change had only a small absol-
ute effect on the American unemployment-rate (though it appears quite 
large once an adjustment has been made for cyclical changes). The impor-
tant effect of demographic change in the United States was upon the

1. With the $\frac{\delta u}{\delta p_3}$ estimate at zero, this would have been 58.3% of the 
increase.
Table 5.11

The Total Effect of Age-Structure Change Over
Three American Trade-Cycles

<table>
<thead>
<tr>
<th>Effect on Overall Unemployment-Rate</th>
<th>1954-58 to 1963-71</th>
<th>1963-71 to 1971-75</th>
<th>1954-58 to 1971-75</th>
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</thead>
<tbody>
<tr>
<td>Augmented Rates Effect</td>
<td>-0.04</td>
<td>-0.02</td>
<td>-0.06</td>
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<tr>
<td>Pure Weights Effect</td>
<td>+0.24</td>
<td>+0.20</td>
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<td></td>
<td></td>
<td></td>
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</tbody>
</table>
structure, rather than upon the overall-rate, of unemployment: the evidence indicates that the increase in the size of teenage cohorts raised their relative rate of unemployment.
Chapter VI
Summary and Conclusions

1. Summary

The paper began by drawing attention to the rises that occurred in unemployment-rates in Canada and other countries during the nineteen-seventies. In Canada, the mean unemployment-rate across trade-cycles increased from 4.54% in 1961-8, to 5.34% in 1968-72 and to 6.77% in 1972-8. By 1978, the rate stood at a post-war record and, as in several other countries, the position was the subject of considerable public debate. The paper has attempted to illuminate such debate by assessing the extent to which there was a link between the high rate of unemployment and the increasing share of young people within the working-age population.

Two channels were identified by which the increase in the relative size of the youth population could have contributed to a rising unemployment-rate. The first channel arose because young people usually experience unemployment more frequently than older persons (perhaps as the result of a desire to sample several different jobs in a short time); because of this, high frequency of unemployment, they have (in Canada at least) always experienced higher-than-average unemployment-rates. As a result, any shift in the composition of the population towards the (unemployment-prone) youth group would tend to increase the overall unemployment-rate even if the age-specific unemployment-rates were to stay the same for every individual demographic group. The first channel by which demographic change could have influenced the overall unemployment-rate was termed the weights-effect because it arose from changes in labour-force weights.
The second channel of influence for demographic change was identified as the rates effect. This effect would operate if there were any way in which a change in the age-structure of the population acted so as to change particular age-specific unemployment-rates. *A priori*, it was thought likely that the entry into the labour-force of large cohorts of young people (born in the post-war baby boom) would probably have served to raise youth unemployment-rates. This prediction was based on the assumption that young workers and old workers are imperfect substitutes for one another. As an approximation, in the short-run, experienced workers could be considered as a factor of production in fixed supply; if young workers are considered as a variable factor, increasing numbers of them could be accommodated fully only if their relative wage dropped by enough to compensate for the fall in their marginal product that would be predicted by the law of diminishing returns. In practice, there is evidence, from the United States, that youth relative earnings indeed fell under pressure of greater numbers in the nineteen-seventies [Welch (1979)]; however, there is no guarantee that youth wages would have fallen by enough to avoid unemployment-rate increases because there were institutional impediments (such as minimum wage laws) that made wage-rates less than fully flexible in the downward direction. Hence, it was thought worthwhile to attempt to assess whether youth-unemployment-rates rose under the pressure of increasing numbers and to investigate whether there were any knock-on effects on the unemployment-rates of other groups.
The importance of weights-effects and rates-effects in Canada were assessed using annual data on the Canadian labour-force for the 1953-78 period. Weights effects were quantified on the basis of decomposing the changes in unemployment-rate between each pair of four distinct trade-cycles. The decomposition was into three components: one was the unemployment-change that would have occurred if weights alone had varied (with age-specific unemployment-rates held constant); the second was the unemployment-change that would have occurred had age-specific unemployment-rates alone varied (with labour-force weights held constant); and the third was an interaction term measuring the covariance between age-structure and unemployment-rates. The first component was used as our estimate of the weights effect.

The second and third components of the decomposition accounted for the bulk of variations in the national unemployment-rate. To establish what proportion of the two components was associated with demographic change (rather than with changes in other exogenous variables), multiple regression techniques were utilised. A measure we called the 'augmented rates effect' (reflecting those parts of the second and third elements of the decomposition that were attributable to changes in population age-composition) was estimated on the basis of eight regression equations. Each equation sought to explain the unemployment-rate of a particular demographic group. Each regression equation included an explanatory variable representing the proportion of the working-age population accounted for by the particular demographic group; this variable was used to examine the impact of relative numbers on group
unemployment-rates. Other variables were included to control for cyclical factors, for variations in unemployment-insurance provisions and for variations in the real minimum wage; a real earnings variable was also included, because of its potential impact on participation-rates and therefore on labour-force weights.

Pure weights-effects in the Canadian case were found to be quite small. Comparing the trade-cycle for which the proportion of young people was lowest (1954-8) with that for which it was highest (1972-8), the weights effect was calculated to have added 0.21 percentage points to the national unemployment-rate between the two periods. This was less than one-twentieth of the observed rise in the unemployment-rate between the two periods.

When augmented rates-effects were estimated for Canada, they were also found to be of limited significance. Demographic change, working via this channel, was estimated to have added 0.22 percentage points to the unemployment-rate between 1954-8 and 1961-8 but thereafter (in the period when unemployment was rising most rapidly) the effects were minimal.

Weights-and-rates-effects were added together to derive an estimate of the overall impact of demographic change on the rate of unemployment. Between 1954-8 and 1972-8, three-tenths of a percentage point was added to the Canadian unemployment-rate by changes in the age-composition of the population; however over two-thirds of the impact had already been felt by the end of the 1961-8 trade-cycle so that relatively little was added to our understanding of the causes of the
subsequent increase in unemployment after 1972.

While the impact of demographic change on the overall unemployment rate was found to be relatively small, the impact was decidedly non-trivial on the rates for particular demographic groups. Teenage girls suffered most when their relative numbers increased; the impact on 20-24 year olds was much less than on teenagers of either sex; and older males benefited from somewhat lower unemployment as their relative numbers declined. Detailed results were presented in Chapter IV.

An empirical analysis was also carried out for the American case, though it was less detailed than for Canada. The aggregate effects of the baby-boom generation reaching working-age were similar to the effects in Canada, with a broadly similar pattern across age groups also; again the most unfavourable impact was on teenagers, with the beneficiaries of change being older males. The results were presented in Chapter V.

For both Canada and the United States, an incidental finding of the analysis was that female-unemployment-rates and teenage-unemployment-rates were relatively insensitive to cyclical factors (in the sense that the responsiveness of group-unemployment-rates to prime-male unemployment-rates was inelastic). For Canada, large effects on unemployment were attributed to the introduction of the 1971 Unemployment Insurance Act. Since 1978, the unemployment-insurance scheme has been made less generous; in future empirical work, it will be interesting therefore to observe whether investigators find that these new legislative changes have had a corresponding downward influence on
the rate of unemployment in Canada.

Although light has been cast in the preceding chapters on such matters as the effect of unemployment-insurance arrangements, this paper's primary aim was to examine the effect of demographic change. Therefore, it is on our findings with respect to the relationship between unemployment and population age-composition that the remainder of the concluding chapter is focussed.

2. The Outlook for Youth Unemployment

The present study suggests that the labour market in Canada has had difficulty in absorbing the large cohorts of new entrants that have been a feature of the last decade: youth-unemployment-rates have increased as a result. However, in the next decade, the adjustment to population change will need to be in the opposite direction. Births in Canada reached a peak in 1959 and fell steadily thereafter, particularly in the years following 1964.1 From now (1983) until almost the end of the century, there will therefore be year-by-year falls in the number of young people attaining working-age. On the basis of our results, this decrease in the potential supply of youth labour should bring some relief to the currently high rates of unemployment amongst the young. However, even with this favourable demographic background, the young will continue to experience unemployment-rates much above the average for Canadians generally. Even in 1954 when the proportion of teenagers in

1. In 1964 the number of live births was 452,915; by 1973, the number had fallen to 343,373.
the population was at its lowest, this group still had an observed un-
employment-rate double that of the population as a whole. If Canada
returns to that relative unemployment position, and if that rate of
youth-unemployment is still considered as a problem, the various levels
of government would need to formulate positive labour-market policies
to deal with it. One line of approach would be to follow training
policies that would make young people more substitutable for older
workers in the process of production; at any given relative-wage, that
would make them relatively more employable.

While the fall in the numbers of young people is predicted to
alleviate youth-unemployment, the consequent fall in the overall unem-
ployment-rate will be very modest. Because the proportion of teenagers
in the population is quite small, changes in teenage-unemployment con-
tribute little to changes in overall unemployment. Just as the demo-
graphic effects of increasing numbers were to raise youth unemployment-
rates significantly whilst leaving the overall rate only slightly higher,
so the effects of falling numbers would be expected to lower youth un-
employment-rates significantly whilst only slightly helping overall
rates to fall.

A small favourable effect on the overall unemployment-rate is
indicated from the preceding discussion of likely trends in teenage
unemployment. However, could these favourable effects not be cancelled
out by increased unemployment elsewhere as the products of the post-war
baby-boom reach older age-groups? This is an important question. We
need to consider whether, having been penalised in the youth labour-
market, baby-boom cohorts will continue to suffer as they pass through later stages of the labour-market life cycle.

3. The Outlook for the Baby-Boom Cohorts

An increase in cohort-size can raise the cohort's relative unemployment-rate because workers of different ages are imperfect substitutes for one another. Age, here, seems likely to be a proxy for amount of experience: workers with different levels of experience often perform different tasks and are therefore in some sense different factors of production. A cohort of unusually large size would experience an unusually high unemployment-rate in its teenage period because new entrants have no work experience and are therefore very poor-substitutes for prime workers; to the extent that the relative wage fails to adjust fully to the changed supply position, teenage-unemployment will rise.

By the time our large cohort reaches the 20-24 age-group, its position will not be as unfavourable because most of its members will have acquired some work-experience.¹ The elasticity of substitution between 20-24 year-olds (with some experience) and mainstream workers is greater than between teenagers (with no work experience) and mainstream workers. As a consequence, the fall in relative wages required to ensure the full-employment of the large cohort is correspondingly less severe and the scope for wage inflexibility to generate unemployment correspondingly less great. This prediction, of a lessening over

¹ This is true even though they have experienced a high group-unemployment-rate during their teenage years - the proportion of the rate accounted for by long-term youth-unemployed is small.
time of the disadvantage suffered by members of large cohorts, is supported by the regression results, which indicate a lower elasticity of group-unemployment-rate with-respect-to population-proportion for the 20-24 group than for the 15-19 group.  

As the large cohort passes through the prime age-group, 25-44 years, its members will have reached an experience level near to the average for the population as a whole. At this stage, it should experience few adverse effects from cohort-crowding since workers of average experience can be substituted not only for workers of a higher experience-level but also for workers with a lower experience-level. Substantial wage adjustments should not therefore be required and indeed the regression results show no effect from variations in relative population-size amongst this group.

Later still, in its progression through the life-cycle, our cohort could again suffer from its large size when it becomes part of the "45 and above" age-group. This group will have the highest average-experience-level of any in the labour market and will therefore once again constitute a group distinct from the mainstream workers. Barring any sudden increase in births in the next few years, baby-boom cohorts will reach this peak-experience level at a time when the supply of younger workers will still be relatively small. Crowding-effects could then be anticipated to arise once again. An increasing supply of the

1. see Table 4-5.

2. or some radical change in immigration policy.
factor of production "highly experienced labour" would be being added to an (approximately) fixed supply of "inexperienced" labour with the result that the relative wage of the former group would need to adjust downward. The downward adjustment could occur either through falls in the wage paid for particular tasks or through a fall in the average-skill-level of tasks performed by older workers. If the downward adjustment failed to materialise in adequate measure, unemployment amongst older workers would surely rise. In fact, a rise is predicted by our regression results, which attributed significant unemployment effects to relative-cohort-size in the oldest male group.

While these problems await the baby-boom generation in later life, our results seem to indicate that the penalty of being part of a large cohort will not be felt as keenly by its members during the next decade as during the last decade. Thus, at least as far as unemployment-experience is concerned, Easterlin's prognosis that "year of birth marks a generation for life" seems too pessimistic.

4. Some Suggestions for Further Research

For the United States, Welch (1979) examined the effects of cohort-size on cohort-average-earnings by schooling-group. Such a division of the labour-force by years-of-schooling would also add interest to any further study of the impact of cohort-size on unemployment-rates. Following Welch, we would predict that the effect of cohort-size on unemploy-

ment-rates should be greater the more highly educated the group under examination. Thus, university-graduates tend to enter occupations with more distinct career patterns than those typically entered by non-graduates; the elasticity of substitution seems likely to be greater between experienced and inexperienced labourers than between experienced and inexperienced lawyers. As a result, the wage-adjustments required by the appearance of a large cohort of labour-market entrants should be greater in graduate occupations than in non-graduate occupations. If the degree of wage-inflexibility were similar in each occupational category, the unemployment consequences of increased cohort-size would then be greater amongst graduates than amongst non-graduates. Testing this hypothesis would not only increase our understanding of labour markets but would also contribute to the current debate on education policy. Recent rises in graduate-unemployment (relative to unemployment amongst non-graduates of the same age) have been taken as a sign that the university-system has been over-expanded, but they could in fact be the result of temporary cohort-crowding effects: the peak period for entry into the labour-force by baby-boom cohorts has now passed for the non-graduate group but is now being experienced for the graduate group, who perhaps tend to suffer relatively more than their less highly qualified contemporaries.

Another obvious extension to this study would be to extend the analysis to other countries, preferably to those with a different time-profile of births from Canada and the U.S.A. The author had hoped to carry out a study for the U.K. However, data on age-specific unemployment-rates were unavailable for Britain.
Finally, we have spoken rather loosely here of elasticities of substitution between different groups of workers, distinguished by experience-level. It would be of interest to obtain estimated values of such elasticities by estimating an aggregate production function with workers in different age-groups entered as separate factors of production.
Appendix I

Canadian Labour Force Data:

Construction of Series

All the empirical analysis for Canada in this paper has been based on a series of labour force statistics constructed at McMaster University by a group in the Department of Economics. This appendix sets out the methodology used to generate the data. Since the data have not been published, those figures used in the analysis have been reproduced as Appendix II.

The construction of a series was necessary in order to allow coverage of the whole of the 1953-78 period. The Statistics Canada Labour Force Survey was inadequate for this purpose because of two breaks in the coverage that made its 1953-78 series internally inconsistent. The more important of these breaks occurred in 1975 when the Labour Force Survey changed the questions that were being used in the sample survey to classify a respondent by labour-force status. The questions used previously had been believed to be too indirect to yield accurate results. For example, the 1971 Census results had suggested that the old Survey was omitting from the unemployment total some women whose job-search activity was, in their minds, secondary to keeping house [Kempster (1973)]. When the error in questioning was corrected in 1975, the result was that, for subsequent years, the rate of female unemployment rose above the male.

1. The work was carried out by Frank Denton, Margaret Derrah, Christine Reaver, Leslie Robb and Byron Spencer.
rate. Previously Canadian statistics had shown the reverse relationship between male and female unemployment—a contrast with the situation reported in most other countries. The new Survey also changed slightly the relationships between the unemployment-rates of the various age-groups and reduced slightly the overall unemployment-rate for Canada.

The problems caused by the break in the Labour Force Survey in 1975 were eased by Statistics Canada having collected labour force data for that year, using both the old and new survey methods. For each age-group, the proportion of the "new" labour-force size to the "old" labour-force size was observed in each month. This proportion was then applied to each corresponding month's labour-force size over the 1953 to 1974 period. The old Labour Force Survey results were thus adjusted on a monthly basis to correspond to the definitions and concepts of labour force participation introduced in the New Labour Force Survey. These adjusted figures were then used to generate the annual series reported in Appendix II. The method was applied in the same way to the number-of-unemployed series. The unemployment-rate series was calculated from the number-of-unemployed series and the labour-force size series for each age-group.

The other "break" in the coverage of the Labour Force Survey occurred in 1961-2 when the youngest group in the labour force was redefined as consisting of 15-19 year olds rather than 14-19 year olds: i.e., fourteen-year-olds were excluded after 1961-2. The problem was to adjust figures prior to 1961 so that they were comparable with the post-1961 figures in which fourteen-year-olds were unrepresented.
Fortunately, Statistics Canada collected figures on both bases for each month from April, 1961, to March, 1962. From these twelve months of figures, one calculated twelve proportions of 14-year-olds to 14-19-year-olds. The April, 1961 proportion was applied to each April statistic from 1953 to 1960 to find the number of 14-year-olds in the particular month; this number was then subtracted from the number of 14-19 year olds; a similar adjustment was carried out for each other month (May, June, and so on). The result was a monthly series for "the estimated number of 15-19 year-olds in the labour force" for the whole period from 1953. This method of adjustment was applied to cover the number-of-unemployed also.

1. All the calculations and revisions cited were based on Labour Force Survey data originating in "CANSIM", which is the registered Trade Mark for Statistics Canada's machine-readable data base.
## Appendix II: Unpublished Canadian Data Used in Empirical Analysis

### Table AII-1

<table>
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<th>Year</th>
<th>All Ages</th>
<th>Male 15-19</th>
<th>Male 20-24</th>
<th>Male 25-44</th>
<th>Male 45+</th>
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APPENDIX III

Model on the Determination of Male Unemployment-Rates

Derivation of the Reduced Form

\[ \log u_i = a_1 + b_1 \log w_i + b_2 \log u_3 + b_3 \log wm + b_4 d \]  

(1)

\[ \log w_i = \log p_i + (\log PR_i - \log PR) \]  

(2)

\[ (\log PR_i - \log PR) = a_3 + b_5 \log u_i + b_6 \log U + b_7 \log fert_i + 
   b_8 \log fert_A + b_9 \log em + b_10 \log ef + b_11 \log p_i + 
   b_{12} \log wm \]  

(3)

Substitute (2) into (1)

\[ \log u_i = a_1 + b_1 \log p_i + b_1 (\log PR_i - \log PR) + b_2 \log u_3 + 
   b_3 \log wm + b_4 d \]  

(4)

Substitute (3) into (4)

\[ \log u_i = a_1 + b_1 \log p_i + b_1 a_3 + b_1 b_5 \log u_i + b_1 b_6 \log U + 
   b_1 b_7 \log fert_i + b_1 b_8 \log fert_A + b_1 b_9 \log em + 
   b_1 b_{10} \log ef + b_1 b_{11} \log p_i + b_1 b_{12} \log wm + b_2 \log u_3 + 
   b_3 \log wm + b_4 d \]  

(5)
Re-arrange (5)

\[(1-b_1 b_5) \log u_1 = (a_1 + b_4 a_3) + b_1 (1 - b_{11}) \log p_1 + b_1 b_6 \log U + b_1 b_7 \log \text{fert}_i + b_1 b_8 \log \text{fert}_A + b_1 b_9 \log \text{em} + b_1 b_{10} \log \text{ef} + b_1 b_{12} \log \text{wm} + b_2 \log u_3 + b_3 \log \text{wm} + b_4 d \]  

(6)

\[
\log u_1 = \frac{a_1 + b_4 a_3}{1 - b_1 b_5} + \frac{b_1 (1 - b_{11})}{1 - b_1 b_5} \log p_1 + \frac{b_1 b_6}{1 - b_1 b_5} \log U + \frac{b_1 b_7}{1 - b_1 b_5} \log \text{fert}_i + \frac{b_1 b_8}{1 - b_1 b_5} \log \text{fert}_A + \frac{b_1 b_9}{1 - b_1 b_5} \log \text{em} + \frac{b_1 b_{10}}{1 - b_1 b_5} \log \text{ef} + \frac{(b_3 + b_1 b_{12})}{1 - b_1 b_5} \log \text{wm} + \frac{b_2}{1 - b_1 b_5} \log u_3 + \frac{b_4}{1 - b_1 b_5} d
\]

(7)

i.e., \( u_1 = f(p_1, U, \text{fert}_i, \text{fert}_A, \text{em}, \text{ef}, \text{wm}, u_3, d) \)

\[
\frac{d \log u_1}{d \log p_1} = \frac{b_1 (1 - b_{11})}{1 - b_1 b_5}
\]
## APPENDIX IV

Minimum Wage Series for Canada used in Empirical Analysis

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Source: See Chapter IV
REFERENCES


Green, Christopher and Jean-Michel Cousineau (1976) Unemployment in Canada: The Impact of Unemployment Insurance, Economic Council of Canada (Ottawa).


