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Résumé de l'article

On accepte souvent comme une évidence le fait que plus le taux de chômage global est élevé, plus la durée des périodes de chômage est longue. Les données agrégées sur la durée des périodes interrompues de chômage confirment certainement cette tendance. La théorie économique ne présente toutefois pas d'explications concernant ce lien. Par ailleurs, on a récemment soulevé la question de sa validité empirique. Darby, Haïtiwanger et Plant (1986) soutiennent que la durée moyenne du chômage global augmente pendant une récession parce que le coefficient de pondération de la composition des chômeurs est plus élevé en raison des particuliers qui vivent « normalement » de longues périodes de chômage, et non parce que les périodes de chômage sont plus longues qu'elles ne le seraient normalement.

L'objet de la présente recherche est d'examiner l'une des répercussions de ce point de vue, notamment celle selon laquelle l'étude (selon des caractéristiques individuelles) de la probabilité de quitter le chômage est invariable pendant la durée du cycle économique. Bien que maintes études canadiennes portent sur la durée des périodes de chômage, peu d'entre elles comportent un horizon temporel suffisamment long pour que l'on puisse découvrir de nouveaux éléments. Les données employées proviennent de l'Enquête annuelle sur l'activité pour les périodes allant de 1978 à 1980 et de 1982 à 1985. Pour examiner les périodes de chômage selon chaque année, on a recours à des modèles de durée de vie accélérée appliqués à une variété d'hypothèses relatives à la distribution.

Il en ressort trois résultats importants. Tout d'abord, l'hypothèse selon laquelle les probabilités de quitter le chômage soient constantes pendant le cycle est rejetée; la direction que ces probabilités prennent n'est cependant pas très claire. Tout laisse croire que la durée des périodes de chômage a augmenté entre 1980 et 1982 avec le début de la récession, et qu'elle a diminué avec la reprise de l'économie. Toutefois, les résultats révèlent également un accroissement de la durée des périodes de chômage lorsque le taux de chômage était à la baisse, surtout entre 1984 et 1985. En second lieu, la courbe varie selon les caractéristiques individuelles, et surtout selon l'âge. Les résultats révèlent une détérioration importante et permanente de la position relative des chômeurs plus âgés. La durée moyenne de la période de chômage chez les personnes plus âgées n'a pas seulement augmenté au début de la récession de 1981-1982, elle a continué de progresser au cours de chacune des années qu'a duré la reprise économique subséquente. La situation est tout à fait différente en ce qui concerne les personnes dans la force de l'âge et les plus jeunes.

Enfin, du point de vue méthodologique, on constate que les résultats de l'estimation sont robustes pour plusieurs formes fonctionnelles, mais non dans le temps. Les chercheurs qui emploient des méthodes paramétriques pour analyser la durée des périodes de chômage devraient savoir que la conjoncture observée au moment de la collecte des données peut influer grandement sur la grandeur et les signes de leurs coefficients estimatifs.

Ces résultats permettent de faire des inductions quant à la nature du processus d'ajustement dans le secteur de la main-d'œuvre. Les ajustements qui suivent de graves bouleversements peuvent raisonnablement être caractérisés comme un processus évolutif au sein duquel les travailleurs plus âgés sont, d'une façon ou d'une autre, placés dans une retraite semi-permanente ou permanente, tandis que les travailleurs plus jeunes sont réengagés à un rythme plus rapide. Les résultats démontrent également le besoin de poursuivre l'analyse de la durée des périodes de chômage avec des données plus récentes et des covariables plus élaborées.

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Cyclical Variations in the Duration of Unemployment Spells

Miles Corak

The purpose of this paper is to examine one implication of the view that the duration of unemployment is invariant over the course of the busines cycle. The data used are derived from the Annual Work Patterns Survey for the years 1978-80 and 1982-85. Accelerated life-time models under a variety of distributional assumptions are used to examine unemployment spells from each year of data.

It is often accepted as a truism that the higher the economy-wide unemployment rate, the longer are the durations of individual unemployment spells. A cursory look at the aggregate data on the duration of interrupted unemployment spell lengths from the *Labour Force Survey* could certainly suggest as much. Economic theory, however, does not offer unambiguous predictions concerning this relationship, and recently its empirical validity has come into question. Darby, Haltiwanger, and Plant (1986:1) argue, on the basis of an analysis of U.S. data, that the "main proximate determinant of changes in the unemployment rate is variations in the level and distribution of inflows into unemployment. Since the probability of leaving unemployment [that is, the duration of an unemployment spell] is primarily determined by the characteristics of those becoming unemployed and is little affected by the business cycle, outflows from unemployment and hence the actual changes in the unemployment rate are primarily determined by the inflows." In other words, the economy-wide average duration of unemployment rises during recessions

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because the composition of the unemployed becomes more heavily weighted with individuals that ''normally'' experience long unemployment spells, not because individual spells are any longer than they normally would be.

Their view is developed in the context of a search theoretic model of unemployment with rational expectations. As the business cycle alters the rate of job offers and the nature of the wage offer distribution, individuals respond by altering their search intensity and reservation wage to the extent that the duration of their unemployment spells is unchanged. This leaves the movement in the unemployment rate, and any movements in the overall average spell length, to be determined by the characteristics of the individuals becoming unemployed. Darby and his co-authors suggest that their results lend support to the Lilien (1982) hypothesis that the dynamics of the unemployment rate are determined by fluctuations in the natural rate of unemployment resulting form increases in the extent of structural adjustments.

The purpose of the research reported in this paper is to examine one implication of this view, namely that the individual probability of leaving unemployment is invariant over the course of the business cycle. While there are many existing Canadian studies that examine the duration of unemployment spells, few have adopted a long enough time horizon to shed any light on this issue. Beach and Kaliski (1987) is one exception, but their major concern is with the distribution of the burden of time spent unemployed, and as a result they do not explicitly model the determinants of the duration of unemployment spells. Picot and Wannell (1987) do touch upon the relationship between unemployment durations and cyclical shocks, but their analysis is focused upon the experience of individuals suffering permanent lay-offs during the 1981-82 recession.

The main contribution of the present research is the explicit modelling of unemployment spell durations over a seven-year period that includes both cyclical peaks and troughs. In methodology the approach is probably closest to the work of Dynarski and Shefferin (1987, 1990) for the United States. The data used are derived from the Annual Work Patterns Survey (AWPS) for the years 1978 to 1980 and 1982 to 1985. The AWPS is a retrospective data set containing information on the labour force status of individual respondents for each month of a given year, a different sample of individuals being interviewed each January. Thus, the seven years of available data represent a series of one year windows through which to view the dynamics of individual labour force activity. For the purposes of this study an unemployment spell is defined to begin when a transition from employment to unemployment is made, and it is defined to end when a transition back to employment from either unemployment or not in the labour force occurs.

The first section of the paper describes the data in greater detail and offers a justification for this particular definition of an unemployment spell. The data are potentially subject to problems of truncation from both the left and the right. The former is dealt with by considering only unemployment spells that have begun during the year, while the latter is explicitly addressed by the estimation procedure. The methodology underlying this procedure and the results obtained are the subject of the second section. Each year of data is analyzed separately by parametric models of spell duration. The validity of the results are limited by the use of explicit functional forms, but an attempt is made to assess their robustness by using as many different distributions as possible. In particular, accelerated life-time models under the assumption that the errors take an exponential, Weibull, log-logistic, and log-normal distribution are all estimated, as is Cox's proportional hazards model, which is a distribution free method.

At least three major results are obtained. First, the analysis tends to reject the hypothesis that individual exit probabilities are constant over the cycle, but the direction in which they move is not clear-cut. There is strong evidence to suggest that the duration of individual unemployment spells increased between 1980 and 1982 — when the economy-wide unemployment rate increased from 7.5 per cent to 11.0 per cent — and then decreased as the recovery took hold. However, the results also reveal increases in spell durations when the unemployment rate was falling, most notably between 1984 and 1985. Second, the pattern of change varies across individual characteristics, most notably across age. The results document an important and long-lasting deterioration in the relative position of older unemployed individuals. The average spell duration of the old, those aged 55 to 69 years of age, not only increased during the onset of the 1981-1982 recession, but it continued to increase during each year of the subsequent recovery. This contrasts sharply with the experience of individuals in the prime-aged, and younger age groups. Finally, from a methodological stand point, it is found that the estimation results are robust across functional forms, but not over time. Researchers employing parametric methods to analyze the duration of unemployment spells should be aware that the magnitude and signs of their estimated coefficients may be very sensitive to the cyclical state of the economy at the time their data was collected. It is also found that the functional form assumed for the analysis does have an important bearing on measures of location: the estimate obtained for the average duration of unemployment for any given year depends very much upon functional form.

A final section summarizes these results, underscores some important caveats, and attempts, in a speculative manner, to draw some implications for the way in which the labour sector responded to the business cycle of the 1980s. In particular, it is suggested that the labour sector adjusted to the adverse conditions of the early 1980s in an evolutionary manner by encouraging the permanent or semi-permanent retirement of older workers from jobs in declining firms and sectors, while at the same time promoting the hiring of younger workers at increasing rates in the new and expanding sectors. The direct substitution of older workers from declining sectors to expanding sectors was of secondary importance.

DATA DESCRIPTION AND A PRELIMINARY ANALYSIS

The Annual Work Patterns Survey contains retrospective information on the timing and length of periods that an individual spends in unemployment (U), employment (E), and not in the labour force (N). The Survey is administered to 5 of the 6 rotation groups in the January Labour Force Survey and addresses their labour force experience over the past year. Data are available for 1977 to 1980 and 1982 to 1985. The 1977 data are not included in this study because of a change in the questionnaire. The Survey was not administered for 1981, and it was discontinued in 1986.

Individuals are assigned a labour force status for each month of the year, and a particular algorithm is used to arrange these individual monthly states into a sequence for the entire year. The algorithm, which is described in Statistics Canada (1982), serves to maximize the length of time that an individual spends in employment. The basic unit of measure is a so-called "part-month," which may essentially be thought of as one-half month.

The analysis is conducted with the micro data types provided by Statistics Canada in which the individual records are survey respondents. A spell file was created from these tapes making the unit of observation an unemployment spell. Thus, information on whether or not an individual experienced multiple spells is not used. A spell is defined to begin with an E-U transition, and to continue until a transition is made back to employment. After the first period of unemployment any transitions between unemployment and not-in-the-labour-force are ignored so that spells may end with a U-E transition or with a N-E transition. Thus, the analysis applies strictly to those individuals leaving their jobs, for whatever reason.

The justifications for using this definition of spell duration are at least three in number. First, and most importantly, permitting an unemployment spell to end when an individual left the labour force would simply imply that the sequencing algorithm was being retraced. The resulting spell durations would be artificial constructions from this algorithm rather than reflections of the actual transitions undertaken. Second, the difference between unemployment and not-in-the-labour-force is very soft in this data set suggesting that there may not be a great behaviourial distinction between the two states, and that it is therefore appropriate to aggregate them into a single state. Job search at any *one* point during the month is sufficient to lead to an unemployment classification for the *entire* month. An individual cannot by the definitions used in the survey be both unemployed and out of the labour force during the same month even though most individuals are not likely to have searched for work continuously throughout the month.[1] Further, the requirement in the constructed spell file that the individual have an unemployment classification immediately after the employment separation will, to some extent, have the effect of removing those individuals not able or willing to work from the analysis, while the lack of distinction between subsequent periods of U and N will explicitly permit discouragement to be reflected in the data.[2] Finally, the definition of unemployment permits the analysis to be conducted as a two-state model. If an unemployment spell is permitted to end with an exit to employment or an exit into non-participation a competing risks model would be necessary. To avoid this some researchers have aggregated employment with not-in-the-labour-force: examples of such work include Darby, Haltiwanger, and Plant (1985, 1986), and much of the work of Hasan and deBroucker (1982, 1985). If the analysis must be restricted to a two-state model it seems preferable that unemployment and not-in-the-labour-force be aggregated together. Kiefer (1985), Dynarski and Shefferin (1987) and some of the work of Clark and Summers (1979) are examples of studies that have used such an aggregation.

A number of limitations in the data should be noted. First, the definition of unemployment is not equivalent to the traditional search-definition used in the *Labour Force Survey*. Any direct comparisons should be made cautiously, and the interpretation of any results should recognize the broader nature of the definition being used. The AWPS data may also be subject to a recall bias. Statistics Canada (1982), and Beach and Kaliski (1984) have examined the extent of this bias, and find, among other things, that there is an under-reporting of unemployment for periods early in the year. This bias may also be reflected in the results. Second, the time horizon of the analysis is restricted to one year. Any projections of spell durations beyond this limit are therefore questionable. Third, the co-variates available are limited to demographic, educational, and regional variables. Potentially important determinants of spell

1 The possibility that the distinction between U and N is soft is also supported by the fact that gross flows data from the *Labour Force Survey*, which uses a definition of unemployment broadly similar to the AWPS, reveal significant flows from N to E. If one subscribed strictly to a search theoretic interpretation of unemployment such movements are not theoretically possible. Their prevalence in reality suggests that there is little value in restricting the definition of unemployment in these terms.

2 In effect this definition of unemployment accepts the claim by Clark and Summers (1979) that many of the transitions between U and N are spurious and do not reflect changes in behaviour. It should be noted, however, that Flinn and Heckman (1983) challenge this view.

duration such as reason for job separation, reservation wages, industry/occupation, and receipt of unemployment insurance payments are not available.

Table 1 summarizes the number of completed and truncated spells in each year of the available data. Roughly 60 per cent of the spells are right truncated and there are, by construction, no left truncated spells. The large percentage of incomplete spells is due to the one year horizon of the survey. Most spells beginning in the later part of the year will, unless they are rather short, be in progress at the time of the survey in January of the following year. Further, the one year horizon in conjunction with the definition of how a spell may be initiated makes it impossible to observe spells longer than 11.5 months.

TABLE 1

Annual Work Patterns Survey Spell Data: 1978-80, 1982-85

Year	*Total Number of Spells*	*Completed Spells*	*Truncated Spells*
1978	9 627	3 722 (38.7)	5 905 (61.3)
1979	9 522	3 592 (37.7)	5 930 (62.3)
1980	9 877	3 843 (38.9)	6 034 (61.1)
1982	14 265	5 789 (40.6)	8 476 (59.4)
1983	12 156	4 857 (40.0)	7 299 (60.0)
1984	12 168	4 593 (37.8)	7 575 (62.2)
1985	11 371	4 483 (39.4)	6 888 (60.6)

( ) indicates row per cent.

Table 2 presents the sample proportions for each of the available co-variates by year. The co-variates may be divided into four broad groups: demographic, educational, regional, and the date at which the spell started. The sample proportions are relatively stable throughout the period, but there are some notable exceptions. The fraction of household heads, of males, and of married individuals increases slightly in 1982 at the outset of the recession but returns rather quickly to pre-recession levels in the following years. The major change with respect to age concerns the fact that the proportion of spells accounted by prime aged individuals (those 25-44 years of age) increases over the years while that accounted by younger individuals (15 to 19 years of age) falls. These changes coincide with the 1982 recession. The proportion of spells due to individuals in the oldest age category stays relatively constant. This is also the case for the proportion of spells accounted by individuals that claimed to be students at some point during the year. There is, however, a tendency for a greater fraction of spells to be accounted for by the more educated over time.

TABLE 2

Sample Proportions of Co-Variates: AWPS Spell Data, 1978-80 1982-85

	1978	*1979*	*1980*	*1982*	*1983*	*1984*	*1985*
1) *Demographic*							
Household head	50.2	51.0	51.5	52.8	51.1	50.5	50.7
Not head	49.8	49.0	48.5	47.2	48.9	49.5	49.3
Male	63.9	63.3	65.2	66.8	64.8	63.0	61.9
Female	36.1	36.7	34.8	33.2	35.2	37.0	38.1
Married	58.6	56.7	55.4	59.0	56.2	56.9	56.0
Single	36.5	38.0	39.0	35.4	37.7	36.8	37.2
Other	4.9	5.3	5.6	5.6	6.1	6.3	6.8
35-44 years	14.6	14.6	15.0	17.1	16.5	17.6	17.5
15-19 years	18.6	18.3	18.4	13.7	13.9	11.3	12.6
20-24 years	22.3	22.8	23.0	22.1	23.3	22.9	22.5
25-34 years	27.5	27.5	27.6	29.5	29.8	32.2	31.7
45-54 years	10.6	10.5	10.1	11.4	10.2	10.2	9.4
55-69 years	6.4	6.3	5.9	6.2	6.3	5.8	6.3
2) *Educational*							
Not student	86.4	87.2	87.5	88.7	87.9	88.2	87.5
Student	13.6	12.8	12.5	11.3	12.1	11.8	12.5
Some high school	57.2	59.3	58.6	59.1	58.7	58.0	58.8
None or elementary	23.5	21.9	21.6	18.8	17.1	15.5	15.1
Some postsecondary	6.9	6.5	6.9	7.7	8.6	9.2	8.9
Postsecondary grad	7.8	7.9	8.3	9.7	10.4	11.1	11.4
University	4.6	4.4	4.6	4.7	5.2	6.2	5.8
3) *Regional*							
Ontario	17.7	18.6	19.4	19.4	17.4	16.6	15.7
Newfoundland	8.8	7.5	7.2	6.6	7.8	7.4	8.2
P.E.I.	3.5	3.6	3.1	3.2	3.7	3.2	3.1
Nova Scotia	7.3	6.8	6.9	6.8	7.1	7.3	7.6
New Brunswick	9.3	9.8	9.9	8.5	9.9	9.5	9.2
Quebec	17.3	17.5	17.0	15.9	15.2	15.3	16.5
Manitoba	6.6	6.1	6.4	6.7	6.5	5.4	5.5
Saskatchewan	5.8	5.6	6.3	5.8	7.2	7.4	7.5
Alberta	9.6	10.4	10.5	11.9	11.8	15.2	14.1
B.C.	14.1	14.1	13.3	15.2	13.4	12.7	12.6
4) *Start Date*							
3rd quarter	30.7	28.4	30.0	32.8	32.2	31.7	32.0
1st quarter	12.0	11.5	12.3	12.0	12.5	11.8	11.8
2nd quarter	19.0	19.1	20.7	21.3	20.1	19.2	20.8
4th quarter	38.3	41.0	37.0	33.9	35.2	37.3	35.4

The proportion of spells due to those with none or elementary education falls from a high of 23.5 per cent in 1978 to a low of 15.1 per cent in 1985, while that due to those with some post-secondary education, post-secondary graduation, or some university rises.

METHODOLOGY AND ESTIMATION RESULTS

Accelerated failure time models are described by Cox and Oakes (1984: 62-70, 85-87), and Kalbfleisch and Prentice (1980: 33-35, 54-62), and have been employed in studies of unemployment spell durations by *inter alia* Addison and Portugal (1987). Proportional hazards models have, at least since the work of Lancaster (1979), been the preferred method for such studies, but the two methods are in fact equivalent when the exponential or Weibull distributions are assumed, which is more often than not the case.[3] Briefly, it is assumed that spell duration is influenced multiplicatively by the explanatory variables, or equivalently that the natural logarithm of spell duration is linearly related to them. If t indexes spell duration then

$$t = \exp(X\beta)\, t_0^{\sigma}$$

$$\ln t = X\beta + \sigma\varepsilon \qquad (1)$$

Where ε may be thought of as an error with probability density function g( ), σ is a scale parameter, $t_0 = \exp(\varepsilon)$, a random draw from the baseline hazard, and X is a vector of co-variates the first element of which signifies a constant. The main concern in the present analysis is to obtain estimates of β, which, with a change of sign, may also be thought of as parameters of the hazard function. When the data set consists of truncated spell lengths least squares estimation of (1) will be biased and inefficient and resort must be made to maximum likelihood. In such cases the contribution to the likelihood of a completed unemployment spell will be the density $g((\ln t - X\beta)/\sigma)$ while that of a truncated spell will be a probability given by the survivor function $G((\ln t - X\beta)/\sigma)$. The survivor function is defined as one minus the cumulative distribution function.

The above sources offer details on the estimation procedure. In what follows the estimation is conducted by year for the exponential, Weibull, log logistic, and log normal distributions. Characteristics of these distributions and their representation in log-linear models of spell duration are catalogued in Kalbfleisch and Prentice (1980: 21-30). Cox's proportional hazards linear

3 Cox and Oakes (1984:71) prove that with constant explanatory variables accelerated failure time models and proportional hazards models coincide under the assumption of a Weibull distribution for the baseline hazard.

regression model is also employed (Cox 1972). This is a distribution free method based upon the ranking of spell durations.[4]

For the sake of brevity the results for each of the five models for 1980 are summarized in Table 3. The results for all distributions over all of the years are available from the author. The omitted categories in these models are given by the variables listed first in each of the groupings of Table 2, thus the reference category is: household head, male, married, 35-44 years of age, not currently a student, some high school education, Ontario resident, with an unemployment spell that began in the third quarter of the year. The parameter estimates, whenever they prove to be significant, are robust over the distributions, the exceptions being the intercept term and the dummy variables associated with the quarter in which the spell started. This suggests that the models will yield different estimates of location. All of the other significant variables have the same sign and are within one standard deviation of each other in magnitude.[5] This result is in contrast with those obtained by Addison and Portugal (1987) in a study of unemployment durations from a U.S. survey of displaced workers drawn from the Current Population Survey, but in broad agreement with those of Dynarski and Shefferin (1987).

For the purposes of summarizing the results over the seven years under study attention may be focused upon a single distribution. On the basis of the results reported in a preliminary analysis of the empirical hazard rates the log logistic distribution is chosen.[6] Table 4 presents the results of the estimation of this model for each year. Many of the parameters are subject to change in

4 The parameters were established by maximum likelihood using a Newton-Raphson algorithm by the S.A.S. version 5 procedure LIFEREG. Standard errors are given by the inverse of the observed information matrix. In actual fact the estimation procedure was much broader than indicated in the text and included the estimation of fully interactive models by gender for each of the distributions. The likelihood values for these models were not substantially greater with the result that the Akaike Information Criterion suggested the choice of the models with no interactions. An attempt was also made to estimate fully interactive models by province but convergence was never attained regardless of the year or distribution used. Finally, the gamma and generalized gamma distributions were also employed but convergence was much more difficult and for the majority of years under study not attained in spite of the fact that several different starting values for the estimates were used. Lawless (1982) suggests that such difficulties are often the case because of the flatness of the likelihood function. The results for these distributions are not reported in the present paper. Cox's proportional hazards model is estimated by the S.A.S. procedure PHGLM.

5 It should be noted that the important significance test concerning the scale parameter in the Weibull and log logistic is whether it differs from one, since a Weibull model with scale equal to one reduces to an exponential and a log logistic model with a scale greater than one indicates that the hazard is non-monotonic, first rising and then falling. Finally, for the Weibull model the scale parameter, σ, is equivalent to $1/\alpha$ of earlier parlance so that the results suggest that the hazard is declining with spell duration.

6 See Corak (1990a) and the related appendix, that is available from the author.

TABLE 3

Maximum Likelihood Estimates: AWPS Spell Data, 1980

	Functional Form				
Model	*Exponential*	*Weibull*	*Log-Logistic*	*Log Normal*	*Proportional Hazards*
Intercept	1.933*	2.030*	1.499*	1.556*	
	(0.06209)	(0.07243)	(0.07269)	(0.06957)	
Not head	0.1296*	0.1453*	0.1898*	0.1652*	-0.1203*
	(0.04174)	(0.04832)	(0.04989)	(0.04681)	(0.04191)
Female	0.04472	0.04834	0.05021	0.04267	-0.03416
	(0.03999)	(0.04625)	(0.04741)	(0.04438)	(0.04012)
Single	-0.1188*	-0.1319*	-0.1137*	-0.09098*	0.08933*
	(0.04497)	(0.05200)	(0.05371)	(0.05025)	(0.04505)
Other marital	-0.005861	-0.005751	0.02791	0.02287	-0.0005781
	(0.07135)	(0.08254)	(0.08578)	(0.08118)	(0.07162)
15-19 years	0.2368*	0.2676*	0.2895*	0.2413*	-0.2053*
	(0.07475)	(0.08653)	(0.08892)	(0.08352)	(0.07501)
20-24 years	0.2247*	0.2483*	0.2258*	0.1787*	-0.1704*
	(0.05988)	(0.06927)	(0.07155)	(0.06680)	(0.05996)
25-34 years	0.04584	0.05359	0.07983	0.05980	-0.04801
	(0.05114)	(0.05914)	(0.06161)	(0.05815)	(0.05122)
45-54 years	-0.03448	-0.03471	0.01219	0.004673	0.01165
	(0.06479)	(0.07488)	(0.07722)	(0.07326)	(0.06474)
55-69 years	0.1587*	0.1742*	0.1762*	0.1513*	-0.1300*
	(0.08074)	(0.09337)	(0.09598)	(0.08954)	(0.08082)
Student	-0.1701*	-0.1879*	-0.2263*	-0.1974*	0.1314*
	(0.05893)	(0.06824)	(0.06993)	(0.06595)	(0.05924)
Elementary	-0.03841	-0.04163	-0.05819	-0.05563	0.02592
	(0.04550)	(0.05263)	(0.05393)	(0.05062)	(0.04558)
Postsecondary	0.004935	0.006982	-0.02360	-0.02117	-0.002043
	(0.06524)	(0.07543)	(0.07812)	(0.07296)	(0.06528)
Postsec grad	-0.1547*	-0.1700*	-0.1459*	-0.1283*	0.1133*
	(0.05824)	(0.06735)	(0.06982)	(0.06658)	(0.05826)
University	0.003110	0.0007567	-0.05768	-0.04973	0.01211
	(0.07542)	(0.08720)	(0.09169)	(0.08579)	(0.07545)
Newfoundland	0.1945*	0.2089*	0.1877*	0.1624*	-0.1460*
	(0.07104)	(0.08213)	(0.08436)	(0.07847)	(0.07103)
P.E.I.	0.1438	0.1527	0.1292	0.1180	-0.1138
	(0.1020)	(0.1180)	(0.1200)	(0.1117)	(0.1021)
Nova Scotia	0.2703*	0.2970*	0.3062*	0.2612*	-0.2151*
	(0.07471)	(0.08641)	(0.08748)	(0.08117)	(0.07473)

(continued)

TABLE 3 (concluded)

Maximum Likelihood Estimates: AWPS Spell Data, 1980

	Functional Form				
Model	*Exponential*	*Weibull*	*Log-Logistic*	*Log Normal*	*Proportional Hazards*
New Brunswick	0.03548	0.03522	0.04092	0.03404	-0.03182
	(0.06190)	(0.07155)	(0.07351)	(0.06946)	(0.06186)
Quebec	0.4137*	0.4598*	0.4954*	0.4329*	-0.3464*
	(0.05634)	(0.06528)	(0.06572)	(0.06135)	(0.05634)
Manitoba	0.01701	0.01381	-0.001669	-0.0006485	0.0005766
	(0.07207)	(0.08331)	(0.08601)	(0.08103)	(0.07205)
Saskatchewan	0.02256	0.02380	0.03412	0.03811	-0.02231
	(0.07540)	(0.08717)	(0.08830)	(0.08344)	(0.07543)
Alberta	0.01390	0.01239	0.003927	-0.001020	0.005222
	(0.05314)	(0.06973)	(0.07198)	(0.06791)	(0.06034)
B.C.	-0.04343	-0.05284	-0.07245	-0.06799	0.04784
	(0.05459)	(0.06311)	(0.06569)	(0.06223)	(0.05458)
1st quarter	-0.2849*	-0.4096*	-0.6558*	-0.5720*	0.5764*
	(0.04563)	(0.05421)	(0.05720)	(0.05448)	(0.04630)
2nd quarter	-0.2394*	-0.3310*	-0.4640*	-0.4333*	0.4448*
	(0.04074)	(0.04799)	(0.04924)	(0.04684)	(0.04112)
4th quarter	0.2021*	0.3721*	0.3063*	0.2961*	-0.5691*
	(0.05389)	(0.06476)	(0.05622)	(0.05120)	(0.05468)
Scale parameter	1.0	1.156	0.8859	1.457	
		(0.01614)	(0.01194)	(0.01827)	
ln likelihood	-10 364.8	-10 307.5	-9 966.73	-9 752.94	

(.) indicates standard error.
* indicates that parameter differs from zero with at least 10 per cent significance.

their significance levels and even their sign. In some years individuals that were not household heads suffered shorter spells than household heads, sometimes longer spells, and sometimes spells that were not significantly different. In this data set gender does not appear to play a significant role in determining spell duration, except perhaps in 1978 when females experienced longer spells and 1982 when they experienced shorter spells. There are also no simple statements to be made about the role of marital status except for the fact that during 1982 at the height of the recession married individuals tended to have shorter spell durations than their counterparts. Likelihood ratio tests for parameter constancy across all of the sample years were conducted and strongly rejected.[7] The exit probabilities cannot be considered constant across this seven year period.

One significant result presented in Table 4 concerns the importance of age in determining spell duration. Before and during 1982 both the young and the old, the 15 to 24 and the 55 to 69 year old groups, had significantly longer unemployment spells than the prime age group, the 35-44 year olds. The coefficients for 15-19 year olds and 20-24 year olds rivals and often is larger than that of the 55-69 year old group. After 1982 this pattern changes. The young no longer experience spells any different in length than the prime age group, while the situation of the old becomes worse as their coefficient values rise to reach the highest levels for all age groups over all years. Indeed, this is the only coefficient, with the exception of that for 25-34 year olds in 1983, that is statistically significant during the post 1982 period.

In one sense the fact that the 55-69 year olds experience significantly longer spells throughout the period under study should not come as too great a surprise. The definition of unemployment that is being used would permit retirement to be captured as a part of the unemployment spell. However, this retirement comes only after the individual spent some time searching for another job after the initial job separation so that in one sense it might be thought of as the result of discouragement or involuntary in nature rather than of disability or voluntary. Further, this possibility applies mostly to the individuals approaching or older than 65 years of age, a group that represents a very small fraction of the data set. What is important is the change in the environment faced by this group over the course of the sample period. The probability of exit from unemployment for this group fell sharply vis-à-vis other groups after the 1982 recession. Finally, it should be stressed that this discussion refers only to the oldest of the age categories, the behaviour of those 45 to 54 years old, a group often cited as requiring targeted assistance, is never

7 The proper way to proceed in this regard is to obtain the log likelihood value from an estimation of the model over all the available data without distinguishing the year. This forms the restricted model while the models estimated for each year separately, which together are equivalent to a fully interactive model by year, form the unrestricted model. Unfortunately, convergence could not be attained for a model estimated on the aggregate data for the exponential distribution, which is the simplest distribution of those entertained. This avenue was not therefore pursued. Each year of data was used to estimate a log logistic model in which the parameters were restricted to those of 1980 and a likelihood ratio test was conducted in which the unrestricted models are given by the results of Table 4. The likelihood ratio test statistics for 1978 through 1985 were: 104, 116, 232, 155, 184, 151. This statistic is distributed as χ^2 with 28 degrees of freedom. The critical χ^2 value at 1 per cent significance is 48.3, thus the null that the parameters are the same as those for 1980 may be rejected. A more complete analysis might consider employing the estimates of each year in turn as the basis for the null hypothesis.

TABLE 4

Maximum Likelihood Estimates: Log Logistic Model, 1978-80 1982-85

Model	*1978*	*1979*	*1980*	*1982*	*1983*	*1984*	*1985*
Intercept	1.607*	1.664*	1.499*	1.524*	1.591*	1.844*	1.747*
	(0.07471)	(0.07783)	(0.07269)	(0.05853)	(0.06656)	(0.06877)	(0.07230)
Not head	-0.1173*	-0.03724	0.1898*	0.1886*	0.1751*	0.01406	0.03627
	(0.04989)	(0.05111)	(0.04989)	(0.04194)	(0.04453)	(0.04363)	(0.04503)
Female	0.1475*	0.02967	0.05021	-0.07841*	-0.03649	0.04091	0.04648
	(0.04715)	(0.04825)	(0.04741)	(0.04057)	(0.04205)	(0.04196)	(0.04300)
Single	-0.01811	0.03956	-0.1137*	0.1487*	0.01440	0.04284	0.04784
	(0.05427)	(0.05531)	(0.05371)	(0.04557)	(0.04855)	(0.04823)	(0.05005)
Other marital	-0.02837	0.1579*	0.02791	0.1245*	0.1758*	-0.1358*	0.08942
	(0.09031)	(0.09156)	(0.08578)	(0.07095)	(0.07599)	(0.07365)	(0.07466)
15-19 years	0.2524*	0.05216	0.2895*	0.1349*	0.03012	0.03680	-0.08601
	(0.08810)	(0.09141)	(0.08892)	(0.07593)	(0.08440)	(0.08685)	(0.08825)
20-24 years	0.2246*	0.1376*	0.2258*	0.07352	0.07084	0.07224	0.007955
	(0.07096)	(0.07329)	(0.07155)	(0.0709)	(0.06341)	(0.06336)	(0.06484)
25-34 years	0.04289	0.06628	0.07983	-0.01093	0.1317*	0.03328	0.05798
	(0.06165)	(0.06455)	(0.06161)	(0.04787)	(0.05281)	(0.05234)	(0.05395)
45-54 years	-0.02613	0.002554	0.01219	-0.05294	0.05730	-0.1083	0.001885
	(0.07673)	(0.07763)	(0.07722)	(0.05924)	(0.06763)	(0.06814)	(0.07291)
55-69 years	0.1762*	0.1823*	0.1762*	0.1406*	0.2911*	0.3143*	0.2376*
	(0.09278)	(0.09589)	(0.09598)	(0.07475)	(0.08283)	(0.08875)	(0.08532)
Student	-0.1935*	-0.1369*	-0.2263*	-0.2755*	-0.09540	0.1984*	0.04347
	(0.06764)	(0.06994)	(0.06993)	(0.05970)	(0.06401)	(0.06417)	(0.06852)
None or elem	0.06431	-0.03977	-0.05819	-0.1004*	-0.05882	0.03773	0.06243
	(0.05205)	(0.05469)	(0.05393)	(0.04540)	(0.05148)	(0.05483)	(0.05707)
Some postsec	0.01356	-0.04100	-0.02360	0.1461*	-0.1483*	-0.03597	-0.04780
	(0.07515)	(0.07932)	(0.07812)	(0.06176)	(0.06173)	(0.06181)	(0.06494)
Postsec grad	-0.003029	0.08298	-0.1459*	-0.08538	-0.1661*	-0.1060*	-0.07444
	(0.07204)	(0.07602)	(0.06982)	(0.05412)	(0.05684)	(0.05670)	(0.05822)
University	-0.08036	0.08504	-0.05768	0.1744*	-0.16997*	-0.06818	-0.05280
	(0.09048)	(0.09557)	(0.09169)	(0.07698)	(0.07683)	(0.07309)	(0.07689)
Newfoundland	0.3673*	0.3282*	0.1877*	0.3235*	0.06881	0.2798*	-0.04625
	(0.08199)	(0.08901)	(0.08436)	(0.07527)	(0.07569)	(0.08178)	(0.07916)
P.E.I.	-0.09498	0.06242	0.1292	-0.006061	-0.03612	0.02187	0.07766
	(0.1073)	(0.1121)	(0.1200)	(0.09519)	(0.1004)	(0.1065)	(0.1144)
Nova Scotia	0.1597*	0.08813	0.3062*	0.1220*	-0.05432	-0.04829	-0.05082
	(0.08356)	(0.08825)	(0.08748)	(0.07083)	(0.7720)	(0.07814)	(0.07885)
New Brunswick	0.1441	0.06534	0.04092	-0.07926	-0.2526*	-0.1794*	-0.04815
	(0.07836)	(0.07624)	(0.07351)	(0.06359)	(0.06682)	(0.06964)	(0.07423)
Quebec	0.2874*	0.3253*	0.4954*	0.1485*	0.1964*	0.1720*	0.1554*
	(0.06481)	(0.06650)	(0.06573)	(0.05368)	(0.06159)	(0.06373)	(0.06398)
Manitoba	-0.1264	-0.06674	-0.001669	0.05103	-0.05797	0.08482	-0.02158
	(0.08456)	(0.08815)	(0.08601)	(0.07074)	(0.07872)	(0.08794)	(0.08814)

(continued)

TABLE 4 (concluded)

Maximum Likelihood Estimates: Log Logistic Model, 1978-80 1982-85

Model	*1978*	*1979*	*1980*	*1982*	*1983*	*1984*	*1985*
Saskatchewan	-0.02554	-0.1904*	0.03412	0.1330*	-0.07496	-0.08029	0.05643
	(0.09125)	(0.09187)	(0.08830)	(0.7666)	(0.07572)	(0.07710)	(0.08012)
Alberta	-0.07286	-0.07142	0.003927	0.04433	0.1079*	-0.03890	-0.007412
	(0.07438)	(0.07378)	(0.07198)	(0.05806)	(0.06511)	(0.06174)	(0.06522)
B.C.	-0.1783*	-0.05206	-0.07245	-0.1195*	-0.1558*	-0.2491*	-0.08841
	(0.06430)	(0.06721)	(0.06569)	(0.05219)	(0.06092)	(0.06289)	(0.06663)
1st quarter	-0.7297*	-0.8521*	-0.6558*	-0.4391*	-0.5078*	-0.6058*	-0.7645*
	(0.05698)	(0.05959)	(0.05720)	(0.04798)	(0.05131)	(0.05316)	(0.05427)
2nd quarter	-0.3695*	-0.5466*	-0.4640*	-0.2405*	-0.2561*	-0.4311*	-0.5358*
	(0.04966)	(0.05140)	(0.04924)	(0.04006)	(0.04444)	(0.04560)	(0.04601)
4th quarter	0.1904*	0.1586*	0.3063*	0.2621*	0.3256*	0.08533*	0.2165*
	(0.05308)	(0.05423)	(0.05622)	(0.04437)	(0.04886)	(0.04820)	(0.05113)
Scale	0.8648	0.8735	0.8858	0.9005	0.8994	0.8950	0.9014
	(0.01187)	(0.01217)	(0.01194)	(0.009967)	(0.01086)	(0.01109)	(0.01132)
In likelihood	-9 628.45	-9 352.76	-9 966.73	-15 251	-12 781.3	-12 234.2	-11 811.5

(.) indicates standard error.
* indicates that parameter differs from zero with at least 10 per cent significance.

significantly different from the 35-44 year old group.[8] The coefficients associated with the dummy variables indicating the quarter in which the unemployment spell began are as a group the most significant in magnitude, and indicate that spell length increases the later the start date. This appears to be so for all specifications over all years. Dynarski and Shefferin (1987) also report such a finding. The most likely explanation is that these variables capture the recall bias in the sample. The earlier in the year that the spell began the shorter it is reported to be. This would be consistent with the analysis presented in Statistics Canada (1982), that, as noted, documents significant under-reporting of spells during the early part of the year. It is also possible that these results capture the influence of uncontrolled heterogeneity possibly associated with industry effects, reason for job separation, and seasonality. Heterogeneity has been an important theme in the analysis of unemployment spell durations and if left uncontrolled has potentially serious consequences for the properties of the estimates and their standard errors, and biases the results toward a finding of negative duration dependence (Kiefer 1988: 671-673; Lancaster 1979, 1985). As a result the estimates of the scale parameter may not be entirely

8 These results also suggest that researchers should pay particular attention to the manner in which the age variable is entered into their models. If actual data on age, as opposed to categories, is available a non-linear function such as the quadratic or a step function is preferable.

dependable, but in spite of the fact that unobserved heterogeneity is not explicitly controlled for a certain confidence can be expressed in the parameter estimates themselves since they are similar across the functional forms, including Cox's non-parametric procedure.[9]

Such, however, is not the case for the estimates obtained of the conditional location of the distribution. Table 5 presents the estimates of the survivor function implied by the results of Table 4. These are calculated by first employing the method of Suits (1984) to transform the dummy variables so as to represent an average member of the sample rather than the arbitrary reference category and then by making use of the functional form for the log logistic survivor function.[10] For this distribution the median and the mean are the same and are therefore both given by the 50th percentile. For the average member of the sample mean spell duration lies between four and five months, which is considerably longer than that reported elsewhere in the literature. Beach and Kaliski (1987: 263), for example, report a mean duration of about 2.5 months for the 1978 AWPS. The difference stems from the alternative definition of an unemployment spell used in the present study. In fact, the estimates presented in Table 5 lie closer to the average annual unemployment experience of individuals than to previously reported average spell durations. Corak (1990b) reports estimates that vary from 3.3 to 4.6 months for the average annual unemployment experience of the unemployed. The fact that the present figures tend to be slightly above those estimates is exactly as expected since the underlying definition of an unemployment spell will in many cases cause instances of multiple spells for an individual to be added together along with the intervening time spent out of the labour force. However, similar estimates derived from the results of the Weibull distribution are even longer since this distribution has a much greater part of the density in the outer tail. For example, the estimate for the median for 1978 is 5.12 months, almost a month longer than the estimate of 4.2 months of the log logistic distribution.[11] Thus, the estimates of Table 5 are not robust across functional forms.

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9 Lancaster (1985: 162-163) shows that the estimates of elasticities of mean spell duration are correct even if heterogeneity is not controlled for. Strictly speaking, however, his result applies to the Weibull model in which there are no truncated data.

10 Using the earlier notation, since $G(t) = 1/(1+\gamma t^{\alpha})$ is the survivor function for the log logistic distribution then t may be found for given values of G (t) by recognizing that $\gamma = \exp(-X\beta)$ and $\alpha = 1/\sigma$ where β is given by the estimated parameters of Table 4 and σ by the estimate for the scale parameter. The calculations from this procedure and from the transformations of Suits (1984) are available from the author.

11 The Weibull estimate of the mean is even longer since the mean and the median do not coincide for this distribution, the mean being the 63.2 percentile.

TABLE 5

Estimated Log Logistic Survivor Functions: 1978-80 1982-85, Sample Average and by Selected Age Groups

	Survival Quantile	*1978*	*1979*	*1980*	*1982*	*1983*	*1984*	*1985*
1)	*Unemployment Rate*							
		8.3	7.4	7.5	11.0	11.9	11.2	10.5
2)	*Sample Average*							
	0.90	0.63	0.65	0.62	0.66	0.64	0.61	0.68
	0.75	1.63	1.70	1.63	1.78	1.72	1.63	1.84
	0.50	4.22	4.43	4.31	4.78	4.61	4.37	4.96
	0.25	10.90	11.60	11.40	12.90	12.40	11.70	13.40
	0.10	28.20	30.20	30.20	34.60	33.30	31.20	36.00
3)	*15-19 Years*							
	0.90	0.73	0.64	0.72	0.72	0.60	0.60	0.61
	0.75	1.88	1.66	1.91	1.94	1.61	1.60	1.63
	0.50	4.85	4.33	5.05	5.22	4.32	4.28	4.39
	0.25	12.60	11.30	13.40	14.00	11.60	11.40	11.80
	0.10	32.50	29.50	35.30	37.70	31.10	30.60	31.80
4)	*35-44 Years*							
	0.90	0.56	0.60	0.54	0.63	0.58	0.58	0.66
	0.75	1.46	1.58	1.43	1.70	1.56	1.54	1.78
	0.50	3.77	4.11	3.78	4.56	4.19	4.12	4.78
	0.25	9.75	10.70	10.00	12.30	11.30	11.00	12.90
	0.10	25.20	28.00	26.50	33.00	30.20	29.50	34.70
5)	*55-69 Years*							
	0.90	0.67	0.72	0.64	0.73	0.78	0.79	0.84
	0.75	1.74	1.89	1.70	1.95	2.09	2.11	2.25
	0.50	4.50	4.94	4.51	5.25	5.60	5.64	6.07
	0.25	11.60	12.90	11.90	14.10	15.10	15.10	16.30
	0.10	30.10	33.70	31.60	38.00	40.40	40.30	44.00

Table entries for panels 2) — 5) are months of unemployment.

Even so, the movements that they display over the period, which is the main concern of this study, may be valid. Also included in the table is the economy-wide unemployment rate. The estimates in Table 5 offer some evidence that, controlling for the characteristics of individuals, suggests spell durations move counter-cyclically. This is certainly the case as the economy moved into recession from 1980 to 1982, but not as clear-cut during the recovery. Mean spell duration is roughly constant during 1978 to 1980, but spells that began in 1982 are estimated to be roughly one half-month longer than those that began in 1980. This figure then falls with the course of recovery, with the significant exception of 1985 when it increases again. The results presented in Table 4 reveal that many of the co-variates lose significance for this year.

The survivor functions for 15-19, 35-44, and 55-69 year olds are also presented. These results control for all other characteristics of these groups. The counter-cyclical movement in average spell durations is evident for the two younger groups, but not for the older age category. For this category average spell durations increased between 1980 and 1982, but also for every year afterward. This is an attendant consequence of the movement in the parameters reported in Table 4.

Finally, it should be noted that all of these patterns appear to hold for the entire survivor function, not just the 50th percentile. Overall about 25 per cent of an incoming cohort of unemployed individuals can be expected to spend about one year unemployed before finding another job, while 10 per cent may be expected to stay unemployed for 28 to 36 months. This latter figure, however, is questionable. The horizon of the data is limited to one year and it is therefore not strictly valid to entertain projections outside of this sample limit. Even so the movement in the entire distribution and particularly its outer tail is, like the mean spell duration, counter-cyclical. The oldest group is also an exception to this pattern.

CONCLUSION

The present study summarizes the results of research on the determinants of unemployment spell durations of individuals experiencing job separations in each year from 1978 to 1980 and from 1982 to 1985. Accelerated failure time models that incorporate explicit assumptions concerning the functional form of the baseline hazard are estimated for each year, and for a variety of functional forms. Cox's proportional hazards procedure which does not involve parametric assumptions is also employed. The main contribution of the paper lies in the use of seven years of data that cover a period of both positive and negative shocks. Most previous studies of spell durations are restricted to limited points on the business cycle. The results obtained are robust to the functional form assumed, but not necessarily to the year of data used. It is found that the average duration of an unemployment spell increased significantly as the economy moved into recession during the early 1980s, and that, for the most part, it decreased during the subsequent recovery. However, even though the aggregate unemployment rate fell between 1984 and 1985 the average duration of an unemployment spell increased. There is some tendency for spell durations to move counter-cyclically, but this should in no way to be understood as an iron-clad law. Indeed, it is also observed that for some demographic groups economic recovery does little to reduce unemployment spell durations. Most notable in this regard is the change in the relationship between age and spell duration. Before the 1982 recession both the young and the old were subject to longer unemployment spells than the prime aged groups. During the subsequent recovery age no longer played .a role for the young but it became more significant for the old. The spell durations of individuals aged 55 to 69 not only increased at the onset of the recession, but also during each year of the subsequent recovery. Thus, the often stated assertion that older workers have a low incidence of unemployment, but a longer duration is too simplistic: their spells are longer and increasingly so over time.

This result permits some inferences as to the nature of the adjustment process in the labour sector. Adjustment to severe negative shocks might reasonably be characterized as an evolutionary process in which older workers are, in one way or another, moved into semi-permanent or permanent retirement, while younger workers are rehired at increasing rates. More detailed enquiries are required. In particular, there is a need for an examination of the retirement decision, and an assessment of the search behaviour of the old. Murphy and Topel (1987) have examined this line of research for the United States, but there have been no similar analyses of Canadian data.

The present results also call for continued analysis of spell durations for more recent data, and with more extensive co-variates. The observation that the parameter estimates are not robust over time may reflect the possibility that important co-variates are omitted from the estimation. Reason for separation, unemployment insurance eligibility, industry, and occupation all suggest themselves as important influences of spell durations. The main conclusion of this study, that unemployment spell durations are not constant over the cycle should be tempered by the possibility that individual characteristics are not completely controlled, and hence that omitted variables are the cause of the instability in parameter estimates.

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Variations cycliques dans la durée des périodes de chômage

On accepte souvent comme une évidence le fait que plus le taux de chômage global est élevé, plus la durée des périodes de chômage est longue. Les données agrégées sur la durée des périodes interrompues de chômage confirment certainement cette tendance. La théorie économique ne présente toutefois pas d'explications concernant ce lien. Par ailleurs, on a récemment soulevé la question de sa validité empirique. Darby, Haltiwanger et Plant (1986) soutiennent que la durée moyenne du chômage global augmente pendant une récession parce que le coefficient de pondération de la composition des chômeurs est plus élevé en raison des particuliers qui vivent « normalement » de longues périodes de chômage, et non parce que les périodes de chômage sont plus longues qu'elles ne le seraient normalement.

L'objet de la présente recherche est d'examiner l'une des répercussions de ce point de vue, notamment celle selon laquelle l'étude (selon des caractéristiques individuelles) de la probabilité de quitter le chômage est invariable pendant la durée du cycle économique. Bien que maintes études canadiennes portent sur la durée des périodes de chômage, peu d'entre elles comportent un horizon temporel suffisamment long pour que l'on puisse découvrir de nouveaux éléments. Les données employées proviennent de l'Enquête annuelle sur l'activité pour les périodes allant de 1978 à 1980 et de 1982 à 1985. Pour examiner les périodes de chômage selon chaque année, on a recours à des modèles de durée de vie accélérée appliqués à une variété d'hypothèses relatives à la distribution.

Il en ressort trois résultats importants. Tout d'abord, l'hypothèse selon laquelle les probabilités de quitter le chômage soient constantes pendant le cycle est rejetée; la

direction que ces probabilités prennent n'est cependant pas très claire. Tout laisse croire que la durée des périodes de chômage a augmenté entre 1980 et 1982 avec le début de la récession, et qu'elle a diminué avec la reprise de l'économie. Toutefois, les résultats révèlent également un accroissement de la durée des périodes de chômage lorsque le taux de chômage était à la baisse, surtout entre 1984 et 1985.

En second lieu, la courbe varie selon les caractéristiques individuelles, et surtout selon l'âge. Les résultats révèlent une détérioration importante et permanente de la position relative des chômeurs plus âgés. La durée moyenne de la période de chômage chez les personnes plus âgées n'a pas seulement augmenté au début de la récession de 1981-1982, elle a continué de progresser au cours de chacune des années qu'a duré la reprise économique subséquente. La situation est tout à fait différente en ce qui concerne les personnes dans la force de l'âge et les plus jeunes.

Enfin, du point de vue méthodologique, on constate que les résultats de l'estimation sont robustes pour plusieurs formes fonctionnelles, mais non dans le temps. Les chercheurs qui emploient des méthodes paramétriques pour analyser la durée des périodes de chômage devraient savoir que la conjoncture observée au moment de la collecte des données peut influer grandement sur la grandeur et les signes de leurs coefficients estimatifs.

Ces résultats permettent de faire des inductions quant à la nature du processus d'ajustement dans le secteur de la main-d'oeuvre. Les ajustements qui suivent de graves bouleversements peuvent raisonnablement être caractérisés comme un processus évolutif au sein duquel les travailleurs plus âgés sont, d'une façon ou d'une autre, placés dans une retraite semi-permanente ou permanente, tandis que les travailleurs plus jeunes sont réengagés à un rythme plus rapide. Les résultats démontrent également le besoin de poursuivre l'analyse de la durée des périodes de chômage avec des données plus récentes et des covariables plus élaborées.