The effect of educational attainment on welfare dependence: Evidence from Canada

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Abstract

The impact of educational attainment on the duration of welfare spells is analysed using a unique data set derived from the administration of the social assistance program in Canada over the period 1986–1993. The empirical analysis includes controls for demographic characteristics, program parameters, labour market conditions and unobserved individual characteristics. It is found that educational attainment has a much greater impact on the welfare exit rate for women than for men. Additionally, the welfare exit rate for women is more sensitive to family status and program benefits, but less sensitive to the unemployment rate, relative to that for men. © 2000 Elsevier Science S.A. All rights reserved.

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1. Introduction

The social welfare system has come under increasing strains in many western countries. Growing welfare caseloads and program costs have led governments to consider major reforms. A primary objective of the reform proposals has been to encourage the transition from welfare to work and promote economic independence. Consequently, one important component of the reforms considered in Canada and the United States has been the increase in financial assistance for, and even direct provision of, educational services that improve the human capital of welfare recipients. Although such policies are premised on the potential gains from
improving the general skills of recipients, relatively little is known about the relationship between educational attainment and the dynamics of welfare participation.

The objective of this study is to analyse the relationship between educational attainment and the duration of time spent on welfare. The empirical analysis is based on a unique data set drawn from the administrative records of the welfare program in the province of New Brunswick (N.B.), Canada. This analysis represents the first step in gauging the potential effectiveness of policies that encourage schooling and the acquisition of general skills in reducing welfare dependence and program costs over time.

There is an established literature analysing the dynamics of welfare spells in the United States (see Moffitt (1992) for a review). O’Neill et al. (1987) used the National Longitudinal Survey of Young Women to examine AFDC dynamics over the 1968–82 period. The analysis examined the influence of a large array of covariates, including years of education, on spells of AFDC receipt. An extra year of schooling was associated with an increase in the exit rate of 12%. Blank (1989) examined the dynamics of AFDC participation by female headed households over the 1971–1976 period using monthly data from the control group of the Denver and Seattle Income Maintenance Experiment. Blank also controlled for educational attainment by including the number of years of completed education among the set of covariates. An extra year of education was found to increase the exit rate by between 5 and 8%.

There are few studies of the dynamics of welfare participation in Canada. Researchers have recently begun to examine the dynamics of welfare participation using administrative data from several provinces. Bruce et al. (1996) presented descriptive information on individuals and families who returned to the welfare program in British Columbia (BC) within 2 years of an initial exit. The paper showed that welfare returnees were a very diverse group, a majority of whom were single men and women (without children) rather than mostly lone parents. Barrett and Cragg (1998) examined the duration of welfare spells in BC using the same data source and found that the large majority of spells were relatively short (less than 6 months long). However, as the BC data did not contain a measure of schooling, these studies did not consider the relationship between education and welfare dependence. Fortin and Lacroix (1997) examined the impact of program benefits on the duration of welfare spells experienced by single men and women (without children) in Quebec. The authors found that the level of benefits had a significant effect on the length of welfare spell among younger individuals. They also found that years of completed schooling had a significant impact on the exit for single men and especially single women.

Although the results of previous studies indicate the importance of schooling in explaining the dynamics of welfare participation, this has not been a primary focus of the analyses. Indeed, the common treatment of educational attainment as a
single continuous variable measuring years of schooling is very restrictive. For example, the effect of an extra year of schooling for an individual with an incomplete elementary education is constrained to be equal to that for an individual who has a post-secondary qualification. To obtain a fuller understanding of the relationship between schooling and welfare use it is important to allow for more general, non-linear effects of educational attainment. Likewise, it is informative to allow for differential impacts according to whether an individual completed a given course of education; that is, to allow for ‘qualification’ effects.

The empirical analysis presented in this paper uses a large longitudinal data set derived from the administrative records of 10% of individuals who received welfare benefits in N.B. during the period 1986–1993. These data have a number of properties which make them suitable for the analysis of the relationship between education and welfare dependence. First, the data contain very detailed information on recipients’ educational attainment. Secondly, the data contain monthly information on program participation, which is the time unit by which the program is administered, enabling the precise length of welfare spells to be determined and thereby avoiding the problems of time aggregation experienced with annualised data. Additionally, since the data were generated from the computerised case records of the N.B. welfare program they provide reliable information on the time pattern of program participation. This is an important advantage of the data for they are not subject to the problems of systematic nonresponse, recall error or ‘seam bias’ as experienced in retrospective survey data.

However, the data have several important limitations. Like most administrative data, there is no information on individuals when they are not participating in the program. Consequently it is not possible to distinguish between destination states, such as marriage or employment. Furthermore, the administrative data contain only a limited amount of socio-economic information on recipients. In particular, the data do not record recipients’ market wage or any information on their employment history, which may be important in determining the dynamics of welfare participation. These limitations are addressed by implementing estimation techniques which allow for the effects of unobserved characteristics (unobserved heterogeneity).

In the empirical analysis educational attainment is treated as exogenous. There is a substantial literature dealing with the implications of individual’s self-selection of education when evaluating the effect of education on earnings and other labour market outcomes. It is not possible to control for the endogeneity of education in

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1An exception is Ellwood (1986) who used dummy variables for educational attainment and found that single mothers who were high-school dropouts were at greater risk of longer AFDC spells.

2Card (1998) provides an indepth survey of recent work analysing the causal effect of education on earnings.
the present analysis due to the lack of potential instruments in the NB longitudinal data. Although the analysis provides new information on the relationship between education and welfare exit rates, the empirical relationships uncovered should not be interpreted as causal effects nor used as the basis for policy formulation.

It is found that men with elementary schooling had significantly lower welfare exit rates. Education beyond elementary high school was not associated with a significantly higher exit rate. However, higher levels of educational attainment were associated with a progressively higher exit rate from welfare for women. The important differences by gender suggest that education plays a different role in generating labour market opportunities for men and women. It is also found that demographic characteristics, such as the number of children and the presence of a spouse, are more important in explaining women’s welfare participation than that of men. The exit rate for men was relatively more sensitive to the unemployment rate and the duration of past welfare spells. Finally, an important methodological finding of the analysis is that the estimated shape of the baseline hazard function may be sensitive to the treatment of the unobserved heterogeneity distribution even when a flexible baseline estimator is implemented.

The paper is organized as follows. In the next section the institutional features of the N.B. welfare program are briefly laid out and a dynamic model of welfare participation is discussed. Section 3 outlines the methods used in the analysis. The data set is described in more detail in Section 4 and the results of the empirical analysis are presented in Section 5. Concluding comments are presented in the final section.

2. Program structure and theoretical model

Over the period examined in this study the provincial social assistance programs in Canada operated under the Canada Assistance Plan (CAP). The primary objective of CAP was to provide financial assistance to all individuals and families in need. Under CAP the federal government set broad guidelines for the implementation of the eligibility test and shared with the provinces in the funding of the programs. The provinces administered the programs and had considerable discretion in determining the rules and benefit structures of their programs.

In establishing eligibility, the individual’s (or family’s) employability status was first assessed, which determined the asset exemption level applicable to the individual. If assets and current income were below the maximum allowable levels, the individual qualified for assistance. The actual amount of assistance paid depended on employability status, whether a spouse was present and the number of dependent children. The structure of the social assistance program in New

\[\text{National Council of Welfare (1993) provides a comparison of the maximum annual welfare benefits payable to several different family types in each province and territory of Canada.}\]
Brunswick corresponded to the ‘basic welfare program’ with a small earnings disregard, as examined by Ehrenberg and Smith (1994:195–203).

In considering the dynamics of program participation assume for simplicity that the individual faces two discrete alternatives; working full-time or receiving welfare. The choice of state is based on the comparison of utility under each alternative. The individual’s utility when working full-time will be a function of their demographic characteristics, their wage rate and recurrent (per period) costs of full-time employment, such as child care. Assume that the offered wage is stochastic and increasing in education (due to the accumulation of productivity-enhancing general skills) and the length of time in employment (due to the accumulation of skills and experience when employed) and decreasing in the time previously spent on welfare (due to human capital atrophy or employer screening). The presence of either human capital atrophy or employer screening will lead to duration dependence in the welfare exit rate.\(^4\)

An individual’s utility when participating in the welfare program will be a function of personal characteristics, the level of welfare benefits and the length of the current welfare spell. The length of the current welfare spell will directly affect an individual’s utility if participation in the program in itself alter an individuals’ income–leisure trade-off or changes the ‘stigma’ (nonmonetary) costs of participation. These latter effects are further potential sources of state dependence in welfare participation.

The individuals’ decision process can be considered in the context of a search model. New values of the offered wage arrive at random intervals and when this information arrives the individual (myopically) chooses the preferred alternative which maximises current period utility. This stylised model suggests the following reduced-form specification for the exit rate from welfare, \( h \):

\[
h = h(X_i,K_i,B_i,d_i,UR_i,T,S)
\]

where \( X_i \) is a vector of personal characteristics, \( K_i \) is the individuals’ educational attainment, \( B_i \) is the level of benefits, \( d_i \) denotes per-period costs of employment, \( UR_i \) represents labour market variables, such as the unemployment rate, which affect the arrival rate of job offers, \( T \) is the length of time on welfare in the current spell and \( S \) represents duration of time on welfare prior to the current spell.

It is straightforward to derive the following predictions:

\[
\frac{\partial h}{\partial K_i} > 0, \quad \frac{\partial h}{\partial B_i} < 0, \quad \frac{\partial h}{\partial d_i} < 0, \quad \frac{\partial h}{\partial UR_i} < 0, \quad \frac{\partial h}{\partial T} \leq 0, \quad \frac{\partial h}{\partial S} \leq 0
\]

The higher is an individuals’ educational attainment, the higher is the offered wage.

\(^4\)By positing that human capital accumulates while employed, the model also generates predictions regarding off-welfare spell durations. If the offered wage increases with the time off-welfare then the off-welfare hazard (or the welfare re-entry rate) will be decreasing in the duration of the off-welfare spell. That is, the off-welfare hazard rate will also exhibit negative duration dependence.
and the greater is the relative attractiveness of employment over welfare and hence the higher the exit rate. Higher benefits increase the disincentive for supplying labour, thereby lowering the welfare exit rate. Greater (recurrent) costs of employment decrease the relative attractiveness of employment and hence reduce the exit rate. An increase in the unemployment rate is equivalent to a decrease in the arrival rate of job offers which also lowers the exit rate. Furthermore, as the length of a spell increases the effects of human capital atrophy, employer screening, changes in preferences over income–leisure or in stigma costs of participation may lead to a decreasing welfare hazard rate. By similar reasoning, longer prior welfare spells may also reduce the exit rate. The latter two predictions correspond to negative duration dependence and negative lagged duration dependence as discussed in Heckman and Borjas (1980).

3. Methods

The empirical analysis is conducted in a hazard function framework. The analysis proceeds by first estimating empirical survival probability functions by educational status. The estimate of the survival probability at month $T$ is simply the proportion of spells that are at least $T$ months in duration. The empirical survival probability function is a convenient way to summarise the distribution of welfare spell lengths which provides a natural measure of the proportion of spells which are relatively ‘short’ and ‘long’.

A limitation of the empirical survival probability estimator is that it treats the population as homogeneous. Spell lengths potentially vary according to the characteristics of the recipient and with labour market conditions. To control for such covariates the proportional hazard duration model is used whereby

$$h_i(t) = h_0(t) \exp[z_i(t)' \beta]$$

where $h_i(t)$ is the hazard for person $i$, $h_0(t)$ is the baseline hazard common to all individuals, $z_i(t)$ is a vector of observable characteristics (which may vary with $t$) and $\beta$ is a parameter vector to be estimated. For different values of $z_i(t)' \beta$, the hazard function for individual $i$ is shifted proportionally up or down relative to the baseline.

5The empirical survival probability function provides equivalent information to the empirical hazard function (the latter reports the number of spells that end in month $T$ as a proportion of spells that are least $T$ months long). Since the survival functions are naturally smoother than the hazard functions they provide a clearer picture of the differences in spell length by education category.
The estimation approach followed in this paper is an extension of Prentice and Gloeckler (1978) and is detailed in Meyer (1988, 1990) and Lancaster (1990:172–208). The baseline hazard is estimated as a piece-wise constant function. The time axis is divided into a finite number of intervals and a separate baseline hazard parameter is estimated for each interval. This approach is a very flexible method for estimating the baseline hazard function and avoids the imposition of a parametric functional form on the baseline. This is an important advantage of the specification for it has been shown that misspecifying the baseline hazard is a major source of error in drawing inferences concerning both the presence of duration dependence (Manton et al., 1986; Blank, 1989) and the impact of covariates (Heckman and Singer, 1985; Dolton and van der Klaauw, 1995).

The log likelihood function for this model with a sample of \( N \) welfare spells is given by:

\[
L(\beta, \gamma) = \sum_{i=1}^{N} \delta_i \log(1 - \exp(-\exp[\gamma(k_i) + z_i(t_i)'\beta])) - \sum_{t=0}^{k_i-1} \exp[\gamma(t) + z_i(t)'\beta]
\]

where \( k_i \) is the observed length of the \( i \)th welfare spell, \( \delta_i \) equals one if the spell terminates before being right censored and \( \delta_i \) is zero if the spell is censored. In maximizing the log likelihood the \( \gamma(t) = \log \int_{0}^{t+1} h_i(u) \, du \) are treated as parameters to be estimated.\(^6\)

The proportional hazard model may be extended to allow for unobserved individual characteristics. Assuming that the unobserved heterogeneity takes a multiplicative form, the hazard rate is given by

\[
h_i(t) = \theta_i h_o(t) \exp[z_i(t)'\beta]
\]

where \( \theta_i \) is a non-negative random variable assumed to be independent of \( z_i(t) \).\(^7\)

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\(^6\)The average baseline hazard rate over a given interval, \([t_i, t_{i+1}]\), is defined as \( \exp(\gamma(t))/\left( t_{i+1} - t_i \right) \). In implementing this model spell lengths were divided into 21 intervals. The intervals corresponded to months: 1, 2, 3, 4, 5, 6, 7, 8, 9–10, 11–12, 13–14, 15–18, 19–22, 23–26, 27–30, 31–36, 37–42, 43–51, 52–60, 61–72 and 73+.

\(^7\)This estimator is equivalent to a random-effects panel estimator. It is not feasible to implement a fixed-effects (or difference) estimator in the present context due to the presence of right-censored spells and time-varying covariates. If it were feasible to use a fixed-effects estimator it would be limited to a select sample of repeat users of welfare.
Maximum likelihood estimates of the parameter vector and baseline hazard are then obtained by conditioning the likelihood function on $\theta_i$ and then integrating over the distribution of $\theta_i$. This approach requires specifying a distribution for $\theta_i$.

One popular distribution is the unit gamma with variance $\sigma^2$ (a parameter to be estimated) which leads to the log likelihood given by

$$L(\beta, \gamma, \sigma^2) = \sum_{i=1}^{N} \log \left\{ \left( 1 + \sigma^2 \sum_{t=0}^{k_i-1} \exp[\gamma(t) + z_i(t)'] \beta \right)^{-\sigma^2} \right\}$$

$$- \delta_i \left( 1 + \sigma^2 \sum_{t=0}^{k_i} \exp[\gamma(t) + z_i(t)'] \right)^{-\sigma^2}$$

(3.2)

Alternatively, the unobserved heterogeneity distribution can be approximated nonparametrically by a discrete multinomial distribution as suggested by Heckman and Singer (1984). The masspoints, $\mu_j$, and the corresponding probabilities, $p_j$, are estimated with the log likelihood function given by

$$L(\beta, \gamma, \mu, p) = \sum_{i=1}^{N} \log \left\{ \sum_{j=1}^{J} p_j \delta_i \mu_j \exp \left( - \sum_{t=0}^{k_i} \exp[\gamma(t) + z_i(t)'] \right) \right\}$$

$$- \mu_j \exp \left( - \sum_{t=0}^{k_i-1} \exp[\gamma(t) + z_i(t)'] \beta \right)$$

(3.3)

In estimating (3.3) the normalisations $\mu_j = 1$ and $p_j = 1 - \sum_{j \neq i} p_j$ are used. The number of mass points, $J$, is determined from the data. The convention of adding additional masspoints until the log likelihood function value no longer increases is followed. In Section 5 below, estimates from the maximization of the log likelihoods in (3.1), (3.2) and (3.3) are compared.

4. The data

The data analysed in this study were derived from the monthly case records of the social assistance program in N.B. The raw data were the complete case histories of a random 10% sample of all individuals and families who ever received welfare benefits in N.B. during the period January 1986 to December 1993. From the case data the sample of welfare spells was constructed. A spell of welfare was defined as a sequence of consecutive months of benefit receipt.\(^5\) For

\(^5\)To minimise the effects of coding errors or ‘program churning’ on transition rates an exit was defined as two months not in receipt of benefits.
spells beginning after February 1986 it was possible to determine the precise length of the spell unless it was still in progress in November 1993. Spells in progress at the end of the data period are right censored and the econometric methods take this censoring into account. The spells of welfare receipt form the basic unit of the empirical analysis. For each spell there is information on duration and whether it is right-censored, as well as the recipients’ sex and age, whether a spouse was present, the number of dependent children, employability status and the (maximum) potential level of welfare benefits payable.

Potential welfare benefits payable to a family are a function of the number of family members and as a consequence the effect of total potential benefits will in part reflect the number of dependent children and the presence of a partner. To obtain a more pure measure of the effect of welfare benefits on the well-being of families, and hence their program exit behaviour, total benefits are divided by the number of adult equivalents in the family. The adult equivalent scale implicit in Statistics Canada’s Low-Income-Cutoffs for 1992 was adopted. Therefore the benefit variable measures potential real dollars per adult equivalent in the family. The effect of using alternative equivalence scales is examined below.

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9It is not possible to determine the starting date for spells that were in progress in January and February 1986. These ‘left-censored’ spells are dropped from the analysis.
10Individuals who were permanently disabled and individuals who were eligible for Unemployment Insurance but received welfare while their UI claim was processed were excluded from the analysis. The latter ‘UI Pending’ group arose during the late 1980’s as a backlog developed in the processing of UI claims. Typically these individuals were on welfare for a very brief period. The dynamics of welfare participation for these two subgroups of recipients merit separate analyses.
11A person is defined as having low employability if he/she is formally assessed to have severe barriers to obtaining or maintaining gainful employment. The distinction between medium and high employability is determined by the case worker. Individuals classified as having low employability are subject to higher asset exemption levels and are eligible for higher benefit payments relative to other recipients.
12Potential benefits were constructed from the payment tables set out in the relevant NB legislation. Potential benefits payable, rather than actual benefits paid, was used in the analysis in an attempt to obtain a benefit measure exogeneous to individual behaviour. Nominal potential benefits were deflated by the monthly consumer price index for NB, with December 1993 as the base period.
13There are two sources of variation in potential benefits: over time and across family types. The former is due to inflation; however, since payment levels were indexed annually to the CPI this variation is minor. Differences in potential benefits by family type is the primary source of variation identifying the benefit effect.
14The Low-Income-Cutoffs are the relative poverty lines used by Statistics Canada.
15The extreme alternatives are to use total potential benefits, which assumes that benefits are spent on pure public goods within the family, or per-capita potential benefits, which implies that there are no economies of scale in consumption. By using benefits per adult equivalent, an intermediate position is adopted. The adult equivalent scale chosen is particularly appropriate since it explicitly attempts to measure economies of scale in consumption among low-income Canadian families.
The data provide very detailed information on the level of education, and current educational participation, of recipients. This information was refined by constructing a set of 6 dummy variables indicating the level of educational attainment. Specifically, the variables indicate whether an individual has 0–6 years of schooling, 7–9 years of schooling, partial high school, completed high school, partial post-secondary or completed a post-secondary qualification. In addition, a variable indicating whether the recipient was enrolled and attending an educational institution was constructed.

In an attempt to control for business cycle effects the spells are matched with the provincial unemployment rate. The prime age male unemployment rate is used in an attempt to obtain a measure that is exogenous to the welfare program. To control for seasonal variation in labour market conditions a set of dummy variables indicating the quarter of the calendar year over which a given spell progressed where also included in the estimation. Since the data contained multiple welfare spells for a large number of individuals and families, lagged duration was defined as the length of the previous welfare spell experienced by the recipient and included in the estimation.

From the original set of individual records a sample of 11 181 separate welfare spells was generated. The sample was stratified by sex of the recipient to allow the effects of education to differ between men and women. Table 1 presents the descriptive statistics for the samples. The main features of the sample for men are that the average spell length was 10.54 months and the majority of male recipients were single. Male welfare recipients generally had low levels of education, with approximately 50% having 0–9 years of education. A further 24% of the sample had only partial high school, 18% had completed high school and the remaining 8% had some post-secondary education.

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16The administrative data indicated nine possible levels of completed education. This information was aggregated into six categories after preliminary estimation of the Cox partial likelihood model and testing coefficient restrictions across adjacent education categories.

17The impact of the unemployment rate is identified purely from time series variation. There is considerable variation in the series due to the recession of April 1990–October 1992. Although it is possible another time-varying factor may underly any effects found for the unemployment rate, such a factor would need to be highly correlated with the business cycle.

18An issue in the estimation of duration models is the treatment of covariates as either fixed throughout a spell or as time-varying. Gender, age and previous spell duration are treated as fixed and the other variables are treated as time-varying. Although the educational attainment variables are allowed to varying during a spell, there is in fact very little time variation in these variables. Approximately 2% of the spells indicate an increase in educational attainment. The results reported below are robust to the treatment of education as time varying or fixed (at start-of-spell values).

19For the estimation, the continuous explanatory variables are measured as deviations from the mean value.
Table 1
Summary statistics, NB welfare spells 1986–1993

<table>
<thead>
<tr>
<th></th>
<th>Male sample</th>
<th>Female sample</th>
</tr>
</thead>
<tbody>
<tr>
<td>Duration</td>
<td>10.5358</td>
<td>15.5852</td>
</tr>
<tr>
<td>Age</td>
<td>31.6133</td>
<td>29.4045</td>
</tr>
<tr>
<td>Spouse</td>
<td>0.2659</td>
<td>0.1173</td>
</tr>
<tr>
<td>Dep. Children</td>
<td>0.9186</td>
<td>1.0792</td>
</tr>
<tr>
<td>Ed.Grade 0–6</td>
<td>0.1375</td>
<td>0.0692</td>
</tr>
<tr>
<td>Ed.Grade 7–9</td>
<td>0.3621</td>
<td>0.2800</td>
</tr>
<tr>
<td>Ed.PartHS</td>
<td>0.2426</td>
<td>0.2677</td>
</tr>
<tr>
<td>Ed.CompHS</td>
<td>0.1784</td>
<td>0.2708</td>
</tr>
<tr>
<td>Ed.PartPS</td>
<td>0.0433</td>
<td>0.0616</td>
</tr>
<tr>
<td>Ed.CompPS</td>
<td>0.0362</td>
<td>0.0509</td>
</tr>
<tr>
<td>Ed.Curr.Enrolled</td>
<td>0.0488</td>
<td>0.0810</td>
</tr>
<tr>
<td>Empty:Low</td>
<td>0.2729</td>
<td>0.2613</td>
</tr>
<tr>
<td>Empty:Medium</td>
<td>0.0246</td>
<td>0.0604</td>
</tr>
<tr>
<td>Empty:High</td>
<td>0.7024</td>
<td>0.6783</td>
</tr>
<tr>
<td>Welf.Ben.($100/a.e.)</td>
<td>2.8785</td>
<td>3.2895</td>
</tr>
<tr>
<td>Unempl. Rate</td>
<td>0.1199</td>
<td>0.1176</td>
</tr>
<tr>
<td>Previous Dur.</td>
<td>4.4066</td>
<td>6.1704</td>
</tr>
<tr>
<td>Q1</td>
<td>0.2672</td>
<td>0.2188</td>
</tr>
<tr>
<td>Q2</td>
<td>0.2176</td>
<td>0.2339</td>
</tr>
<tr>
<td>Q3</td>
<td>0.2296</td>
<td>0.2635</td>
</tr>
<tr>
<td>Q4</td>
<td>0.2856</td>
<td>0.2838</td>
</tr>
<tr>
<td>right censored</td>
<td>0.1951</td>
<td>0.2669</td>
</tr>
<tr>
<td>observations</td>
<td>6167</td>
<td>5014</td>
</tr>
</tbody>
</table>

In comparison, the sample means for women show that average duration was substantially longer at 15.59 months, a greater proportion did not have a spouse while the average number of children was higher, reflecting the greater incidence of single parents among women. Female recipients generally had higher levels of education. For example, 35% had 0–9 years of schooling, 54% had some secondary schooling and a further 11% had post-secondary education. The average level of benefits was higher for women, as was lagged duration.

5. Empirical Results

The empirical survival probability functions by education category are plotted in Figs. 1 and 2, based on the male and female samples respectively. For presentation purposes only 3 broad education categories are shown. The empirical survival functions show the NB welfare population to be relatively dynamic. The median spell duration for males with elementary schooling is 6 months, for those with
either a high-school or post-secondary education it is 5 months. Males with elementary schooling have longer spells than those with greater education, while the exit patterns for males with either high school or post-secondary education are
very similar. The sample of spells experienced by females has greater variation in spell length by education category. The median spell length for females with elementary schooling is 12 months, with high schooling is 9 months and with a post-secondary education is 6 months. For women, it is apparent that further education is associated with progressively shorter stays on welfare and less reliance on social assistance over a period of time.

The convex shape of the empirical survival functions suggest negative duration dependence; the longer that an individual or family remains on welfare the less likely it is they will exit. However this apparent duration dependence may be a reflection of differences in individual characteristics rather than true state dependence. The results from the duration model estimation, which controls for observable and unobserved characteristics, are presented next.

5.1. Results for the male sample

The duration model estimates for the sample of spells by men are reported in the top panel of Table 2. To aid interpretation of the estimates, the marginal effect of covariates on expected duration are presented in the lower panel of Table 2. Specification (1), which does not control for unobserved heterogeneity, shows that older men have significantly lower exit rates while the presence of a spouse is associated with a substantially higher exit rates. This latter finding is unsurprising since the presence of a spouse means that there are two potential workers in the family who may gain employment thereby enabling the family to leave welfare more quickly than a single adult. More surprisingly, the number of dependent children has an insignificant effect on the exit rate for men.

Turning to the effects of education, males with 0–6 years of schooling had the lowest probability of leaving welfare, with an exit rate almost 18% less than that for an otherwise identical male who completed high school. Similarly, individuals

20Wald tests reject the hypotheses that the survival function for elementary schooling is equal to that for high-schooling (test statistic of 1540.1) or post-secondary education (test statistic equal to 333.5). However, the hypothesis that the survival function for high schooling is equal to that for post-secondary education is not rejected (test statistic of 15.8). The test statistic is distributed chi-squared with 72 degrees of freedom; the critical value at the 1% (10%) level of confidence is 102.8 (87.7).

21The three survival functions are significantly different from each other. The test statistics for the hypothesis of equality for each pairwise comparison is: elementary–high schooling (1352.6), elementary–post secondary (2650.3) and high school–post secondary (848.8). The critical value at the 1% level of significance is 102.8.

22The marginal effect of covariates on expected duration was calculated following the approach detailed in Katz and Meyer (1990: 64–65). The estimate for the last segment of the baseline hazard function (months 61–72) was extrapolated out to 144 months. The calculated marginal effects did not change when the extrapolation was truncated at an earlier month.
Table 2  
Duration model estimates, NB welfare spells  

<table>
<thead>
<tr>
<th>Male sample</th>
<th>Male sample</th>
<th>Female sample</th>
<th>Female sample</th>
<th>Female sample</th>
<th>Female sample</th>
<th>Female sample</th>
</tr>
</thead>
<tbody>
<tr>
<td>Age</td>
<td>0.2189</td>
<td>0.2448</td>
<td>0.2495</td>
<td>0.1889</td>
<td>0.2723</td>
<td>0.2922</td>
</tr>
<tr>
<td>Spouse</td>
<td>0.3477</td>
<td>0.4082</td>
<td>0.4102</td>
<td>0.4164</td>
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Marginal effect of covariates on expected duration (in months)  
Age (10 yr) 0.1209 0.1140 0.1141 0.1082 0.1417 0.1508
Dep.Children 0.1598 0.1030 0.1104 0.4601 0.5117 0.4852
Ed6 1.0033 0.8068 0.7860 1.7272 2.0027 1.8074
Ed79 0.3781 0.3542 0.3426 1.2394 1.3502 1.2108
Ed.ParThr 0.0561 0.0877 0.0779 0.6458 0.8077 0.7272
Ed.ParPS 0.3940 0.4738 0.4914 –1.1870 –1.2795 –1.1655
Ed.CompPS 0.5576 0.7058 0.6982 –1.3562 –1.4225 –1.1308
Ed.Curr.Enrolled 1.3216 1.5572 1.5077 0.8338 1.0305 0.8697
Emp.Low 2.2385 2.3858 2.2323 2.2479 2.4458 2.2321
Emp.Med 1.2018 1.2504 1.2124 0.2615 0.2926 0.3806
Bene/(S20/ae) 0.1620 0.1630 0.1559 0.2816 0.2938 0.2685
Unemp.Rate 1.6251 1.4717 1.4034 1.0111 0.9829 0.9755
Pr.dur. (1 yr) 1.5467 1.6267 1.5514 0.3995 0.4551 0.4306

Asymptotic standard errors are in parentheses. Each model also included three seasonal dummy variables.
with 7–9 years of schooling had an exit rate which was 7% less than that for an otherwise identical high school graduate. However, individuals with only a partial high school education or with post-secondary schooling had an exit rate which was not statistically significantly different from that for high school graduates. In addition, individuals currently participating in a course of education had a significantly lower exit rate which suggests that these individuals maybe reliant on welfare income during this period of schooling.

The coefficient estimates for employability status indicate that, as expected, individuals evaluated as having low or medium employment potential have substantially lower exit rates and hence longer spells. Welfare benefits have a significant negative effect on the exit rate, although the economic impact of an increase in payments (as indicated by the marginal effect on expected duration\(^2\)) is quantitatively small. The point estimate on the unemployment rate is highly significant and implies that an increase in the provincial unemployment rate by 5 percentage points leads to a decrease in the welfare exit rate by 27% and an increase in expected duration by 1.6 months, highlighting the sensitivity of the welfare exit behaviour of men to the general state of the labour market. The coefficient on previous duration is also highly significant and indicates the presence of negative lagged duration dependence. For every year spent on welfare in the previous spell, the current spell has a 27% lower exit rate and 1.55 months longer expected duration.

Although not reported in Table 2, an alternative specification with the set of education dummies variables replaced by a single, continuous variable measuring years of schooling was estimated. The point estimate (and standard error) for this variable was 0.0174 (0.0049). The likelihood ratio test (LRT) statistic of 4.96\(^2\) (\(P\)-value = 0.29) fails to reject the linear restriction implied by this specification; however, this model masks important non-linearities in the effects of education. The significance of the continuous, years of education variable is due to the substantially lower welfare exit rates of males with very low levels of schooling. In comparison to specification (1), the linear years-of-education specification underestimates the magnitude of the reduction in the exit rate for males with only elementary schooling while overestimating the increase in the exit rate for education beyond some high schooling.

The next step in the analysis was to control for unobserved individual characteristics. Specification (2) generalises (1) by allowing for gamma unobserved heterogeneity. Controlling for gamma heterogeneity generally led to an increase in the magnitude of the coefficient estimates, and their asymptotic

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\(^2\)The marginal effect of potential payments on expected duration is calculated for an increase of $20 per adult equivalent per month. This is comparable to the actual changes in benefits that occurred during the data period.

\(^2\)The test statistics has a chi-squared distribution with four degrees of freedom. The critical value at the 1% (10%) level of significance is 13.277 (7.779).
standard error, while the baseline hazard estimates generally decline less with spell duration. Consequently, the marginal impacts of the covariates on expected duration were largely unaffected by incorporating gamma heterogeneity. Moreover, the inferences regarding the impact of educational attainment on the exit rate are unchanged.

The final step was to allow for nonparametric unobserved heterogeneity. As found for the model with gamma heterogeneity, allowing for nonparametric heterogeneity generally led to an increase in the magnitude of the coefficient estimates (and their standard errors). The qualitative inferences regarding the impact of the covariates were unaffected by the assumed form of unobserved heterogeneity. Indeed, the invariance of the coefficient estimates across the specifications shows the robustness of the estimated relationship between welfare exits and educational attainment, demographic characteristics, program benefits and labour market conditions to the treatment of unobserved heterogeneity.

The baseline hazard estimates for the models are illustrated in Fig. 3. The baseline hazards for models (1) and (2) indicate negative duration dependence. Wald tests of the difference in the hazard rates by spell month are highly significant and imply that participating in welfare has a behavioural impact whereby the longer a person is on the program the less likely it is that they will exit. However, this conclusion is not as strongly supported by the baseline hazard estimates for the models illustrated in Fig. 3. The baseline hazards for models (1) and (2) indicate negative duration dependence. Wald tests of the difference in the hazard rates by spell month are highly significant and imply that participating in welfare has a behavioural impact whereby the longer a person is on the program the less likely it is that they will exit. However, this conclusion is not as strongly supported by the baseline hazard estimates for the models illustrated in Fig. 3. The baseline hazards for models (1) and (2) indicate negative duration dependence. Wald tests of the difference in the hazard rates by spell month are highly significant and imply that participating in welfare has a behavioural impact whereby the longer a person is on the program the less likely it is that they will exit. However, this conclusion is not as strongly supported by the baseline hazard estimates for the models illustrated in Fig. 3.

Fig. 3. Baseline hazard functions: Welfare spells by males.

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25This is consistent with the findings of Meyer (1990) and Han and Hausman (1990).
26The data supported the presence of two masspoints. When a third mass point was added to the likelihood function, the point estimate always converged toward that of the second mass point even after starting the optimisation routines from a broad range of initial values.
estimates for model (3) which show a smaller decline in the exit rate over the initial 12–24 months of a spell, and this decline is only marginally statistically significant. Therefore conclusions regarding the degree of duration dependence, and the economic significance of the behavioural effect of participation, depend on the choice of model specification.

An appealing method for choosing the appropriate specification would be to apply the LRT to the unobserved heterogeneity terms. Unfortunately, the standard LRT is inappropriate since the null hypotheses of $\sigma = 0$ (for the gamma distribution) and $p_1 = 1$ (for the multinomial distribution) lie on the boundary of the parameter space. Consequently, the regularity conditions necessary for the asymptotic chi-squared distribution of the LRT statistic are violated.

An alternative is to base the choice of model specification on information criteria. The Akaike Information Criterion (AIC) and Bayesian Information Criterion (BIC) have been used by Gritz (1993) and Deb and Trivedi (1997), respectively, in similar contexts. These criteria have an information theoretic foundation and support the selection of the model which minimises the respective quantities $\text{AIC}(K_M) = -2L_M + 2K_M$ and $\text{BIC}(K_M) = -2L_M + K_M \ln N$, where $L_M$ is the optimised value of the log-likelihood function for model $M$, $K_M$ is the dimension of the parameter vector, and $N$ is the number of observations. BIC places a greater penalty on the inclusion of additional parameters compared to AIC. The AIC and BIC values are reported in Table 2. Both the AIC and BIC support the selection of Model (2) with gamma unobserved heterogeneity, though there is very little difference in the AIC values for Models (2) and (3). The contrast in the baseline hazard estimates between the gamma and multinomial heterogeneity models clearly illustrates that the form of the unobserved heterogeneity may have important consequences for the inferences concerning duration dependence even when a flexible baseline estimator is implemented.

5.2. Results for the female sample

The duration models estimates for the sample of spells experienced by women are also reported in the top panel of Table 2 (and the marginal effect on expected duration are presented in the lower panel). Specification (4) indicates that age does not have a significant effect on the exit pattern for women, contrary to the finding for men. The presence of a spouse has a substantial, and significant, positive impact on the exit rate for women and is larger than that found for men. Dependent children has a highly significant, negative effect on women’s exit behaviour. Although dependent children do not appear to increase mens’ reliance on welfare they have an important effect in increasing the length of time women spend on welfare.

The relationship between education and womens’ welfare exit behaviour is also very different to that found for men. Higher levels of educational attainment are associated with progressively higher exit rates, and hence progressively shorter
stays on welfare. In particular, the exit rate for women with 0–6 years of schooling is 30% less than that for otherwise identical women who graduated high school. Women with 7–9 years of schooling had a 21.7% lower exit rate, and those with an incomplete high school education had an 11% lower exit rate, than comparable women who graduated high school. Post secondary education was associated with even higher exit rates. Therefore educational attainment is clearly important in understanding women’s welfare exit behaviour.

The estimates for specification (4) also show that women currently participating in an educational program had a lower exit rate, as did women classified as having low employment potential. Higher welfare benefits also have a significant negative impact on the exit rate for women, and the impact is quantitatively larger than that found for men. Since the level of potential benefits were greater, on average, for the female sample, combined with the larger impact of benefits on their exit rate, suggest that the disincentive effects of program benefits may be more important for female recipients. Even so, the marginal effect of potential benefits on expected duration is also relatively small for this sample. The coefficient estimate for the unemployment rate implies that a 5 percentage point increase in the unemployment rate leads to an increase in expected spell duration of approximately 1 month for the baseline group. Clearly the welfare exit behaviour of women is sensitive to the state of the labour market; though it is not as sensitive to cyclical factors as the exit pattern of men.

Model (4) includes the measure of previous spell duration. As found for men, previous duration is statistically significant and indicates the presence of negative lagged duration dependence. The estimated coefficient implies that for every year on the program in the previous spell, the exit rate for the current spell is approximately 7% lower (and the expected spell duration is 0.40 months longer). This estimated effect of previous spell length is substantially smaller than that found for men, suggesting the ‘scarring’ impact of welfare participation on women’s labour market careers may be less acute than that for men.

The more restrictive version of model (4) with the education dummy variables replaced by a single variable measuring years of education was also estimated with this sample. The coefficient on years of education was 0.0527 (0.0067), which is more than three times the estimate for males. However this more restrictive model is rejected by the LRT statistic of 10.02 (P-value = 0.040). The linear restriction implied by the years of education variable failed to capture the more than proportionate increase in the exit rate with additional years of education beyond some high schooling. This provides further evidence of significant non-linearities in the relationship between educational attainment and the duration of welfare spells.

The model was generalized by controlling for gamma unobserved heterogeneity. Allowing for gamma heterogeneity typically led to a small increase the magnitude of the point estimates and the corresponding asymptotic standard errors, as found for the spell sample for men. Likewise, the inferences regarding the impact of the
covariates was robust to the inclusion of gamma heterogeneity. The more general model with gamma heterogeneity supported the conclusions drawn from specification (4). The specification incorporating nonparametric unobserved heterogeneity was also estimated.\footnote{This sample also supported the presence of two masspoints.} The choice of distribution for the unobserved heterogeneity component again had very little impact on the qualitative inferences concerning the effect of the covariates. The coefficient estimates (and marginal effect on expected duration) for the schooling and other covariates were entirely robust to form of the unobserved heterogeneity distribution.

The baseline hazard estimates are illustrated in Fig. 4. The distribution of the unobserved heterogeneity distribution again proved critical for drawing inferences concerning the shape of the baseline hazard and the presence of duration dependence. The baseline hazards for models (4) and (5) show a significant decline in the exit rate over the course of a spell, especially in the first 9–12 months of a spell. However, when allowing for nonparametric heterogeneity the decline in the exit rate is no longer evident. Indeed, the baseline hazard function for model (6) is very low and essentially flat.

The two model selection criteria lead to very different conclusions. The AIC supports the selection of model (6) whereas BIC supports model (5). Since no single specification clearly dominates a strong inference regarding duration dependence cannot be drawn. The differences in the baseline hazards for models (6) and (5) further reinforces the finding that the unobserved heterogeneity distribution can be very important in testing for duration dependence even when a flexible baseline estimator is chosen.
5.3. Specification and sensitivity tests

A question of wide policy interest is the effect of program benefits on the exit rate. To examine the robustness of the estimates to the definition of potential benefits, the models were re-estimated with the benefit variable defined as total potential benefits and per-capita potential benefits. For both alternative benefit measures the magnitude of the coefficient estimate was reduced\textsuperscript{28} while the coefficients on the spouse indicator and number of children variables increased.\textsuperscript{29} Total potential benefits are a nonlinear function of family size, and benefits per family member (and per adult equivalent) decline as family size increases. This negative correlation between benefits per family member and family size is reflected in the coefficient estimate on the benefit variable used in Table 2, while the correlation is captured in the coefficient on the dependent children variables when total or per-capita potential benefits are used. The per-adult equivalent benefit measure generated the largest estimate of the disincentive effects of payment levels on spell duration; however, this effect was of relatively minor economic significance.

The sensitivity of the estimated effects of schooling on the welfare exit rate was examined by interacting the educational attainment dummy variables with other covariates. The estimates in Table 2 implied the presence of negative lagged duration dependence. There are several potential sources of this form of state dependence, and if it is due to human capital depreciation then any ‘scarring effect’ of prior welfare use may be greater for those individuals with a larger stock of human capital. The coefficient estimates for the education and education interacted with lagged duration terms are presented in Table 3. Interestingly, the effects of lagged welfare duration appear to be concave in education. The magnitude of the point estimates on previous welfare duration increase with education up to completed high-school and then decline. This finding is consistent for both the male and female samples; however, the hypothesis that the effects of previous duration are the same across all educational categories is rejected only for the female sample.

Cyclical fluctuations in the economy have a differential impact on groups within the labour market. Individuals with higher levels of education tend to be in a stronger position and more insulated from recessions. To test the proposition that the program exit behaviour of individuals with less education is more sensitive to the state of the labour market, the education dummy variables were interacted with the unemployment rate. The coefficient estimates for this model is also reported in

\textsuperscript{28}The coefficient (and standard error) on total potential benefits in Model 1 was \(-0.1196 (0.0336)\) and in Model 4 was \(-0.1933 (0.0228)\). The estimate on per-capita potential benefits was \(-0.1167 (0.0294)\) and \(-0.1621 (0.0177)\) in Models 1 and 4 respectively.

\textsuperscript{29}The estimates for the other covariates were not sensitive to the alternative measures of program benefits.
Table 3
Educational attainment interaction effects

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Asymptotic standard errors are in parentheses. Variables for age, presence of a spouse, number of children, whether currently enrolled in an educational institution, employability status, welfare benefits, the unemployment rate and season were also included.

The LRT statistic is distributed chi-squared, with 5 degrees of freedom, under the null. The critical value at the 10% level of significance is 9.236, at the 5% level is 11.070 and at the 1% level is 15.086.
Table 3. The LRT statistics do not reject the hypothesis that the unemployment rate affects all education groups equally. However, the point estimates for the male sample suggest that individual who did not complete a level of education were more severely affected by the business cycle than those who had completed that level of schooling. The estimates suggest there may be a ‘qualification’ effect in the arrival rate of job offers, with male graduates from a given course of education being in a relatively more secure position.

The final specification checks considered sample composition. The models were estimated with the subsample for single men and women without children. Differences in the exit rate by education group were more pronounced among single, childless women, and the male–female differences in the exit by education, as discussed above, also applied to this subsample. The models were also estimated for the subsample of single mothers. Although the differential in exit rate by educational category were not as large for this group of women, there were significant difference in exit rates by education consistent with the pattern for the full sample. Overall, the qualitative relationship between education and welfare exit rates were consistent across the population subsamples.

6. Conclusion

The empirical analysis indicated that educational attainment had very different effects on the dynamics of welfare participation for men and women. It was found that men with a partial or completed elementary education were more reliant on welfare. Education beyond partial high school was not associated with a significantly higher welfare exit rate. Conversely, additional education was associated with significant and substantially higher welfare exit rates for women. Each step up the educational ladder was associated with a progressively higher welfare exit rate for women. It was also found that demographic characteristics, such as the number of children and the presence of a spouse, were more important in explaining the dynamics of women’s welfare participation than that of men. The exit rate for men was relatively more sensitive to the unemployment rate, and the duration of past welfare spells, than that for women.

The very different relationship between education and the dynamics of welfare participation by gender may reflect the different role of education in accounting for the labour market opportunities of men and women. Cross sectional evidence shows that men are much more likely to be employed in construction and transportation industries while women are more likely to be employed in services. Employment in the male dominated industries is more seasonal in nature, tends to be more sensitive to the business cycle and may not utilise the skills gained with more advanced schooling. In addition, industries with the highest incidence of female employment, such as community and personal services, may utilise the skills augmented through further schooling, such as better numeracy, literacy and communication skills. The different role of education in generating labour market
opportunities for men and women may account for the relationship between education and welfare exit rates uncovered in the analysis.

The results of the analysis are relevant for duration studies more generally. A flexible baseline hazard estimator was implemented while allowing for several forms of unobserved heterogeneity. With the flexible baseline specification, the coefficient estimates (and the marginal effects) for the covariates were found to be robust to the distribution of the unobserved heterogeneity. However, for the samples examined in this study, the shape of the baseline hazard function and hence inferences regarding duration dependence were sensitive to the assumed form of the unobserved heterogeneity distribution. These findings should caution researchers that if the shape of the baseline hazard is of interest it is important to allow for a variety of distributions for the unobserved heterogeneity term before firm conclusions may be drawn.

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References


