The Changing Impact of Women's Educational Attainment and Employment on the Timing of Births in Canada

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Abstract

Analyzing 1995 Canadian General Social Survey data and making use of Cox Proportional Hazards (1972), we consider possible cohort differences in the impact of women's education and employment on birth timing. Consistent with previous studies (De Wit and Rajulton, 1992; De Wit, 1994), our findings demonstrate the continuing importance of education and employment for more recent cohorts and also reflect their increasing impact over time on birth timing. Further analyses of interactions between education and employment show that, for all cohorts, the effect of previous employment increases with level of education, meaning that women with the most education as well as prior work experience have the greatest incentive to initiate childbearing earlier. We conclude that 1) the joint effects of women's education and employment have a more powerful influence on the timing of childbearing than either of their separate effects; and that 2) this combined effect of education and early work experience appears to be increasing over time.

Résumé

D'après une analyse de l'Enquête sociale générale de 1995 et à l'aide du modèle des hasards proportionnels de Cox (1972), nous considérons des différences de cohortes possibles en ce qui touche l'impact de l'éducation et de l'emploi des femmes sur l'échelonnement des naissances. Conformément aux études précédentes (De Wit et Rajulton, 1992; De Wit, 1994), nos résultats démontrent l'importance durable de l'éducation et de l'emploi pour les cohortes plus récentes et illustrent également leur impact grandissant au fil du temps sur l'échelonnement des naissances. D'autres analyses des interactions entre l'éducation et l'emploi montrent que, pour toutes les cohortes, l'effet des antécédents professionnels augmentent avec le niveau d'éducation, c.-à-d. que les femmes les mieux éduquées et bénéficiant d'une expérience professionnelle semblent être plus enclines à procréer plus tôt. Nous concluons que 1) combinés, l'éducation et les antécédents professionnels de la femme exercent une influence plus puissante sur l'échelonnement des naissances que l'un ou l'autre effet pris séparément; et que 2) cet effet combiné de l'éducation et des premières expériences professionnelles semblent augmenter avec le temps.

Key words: *birth timing, higher education, cohort analysis, hazards analysis*

The Research Problem

Research on the socioeconomic determinants of childbearing in economically advanced countries suggests that variables such as education and employment have a significant impact on the timing of births. Higher education is found to be associated with a delay in the timing of first birth (Wineberg, 1988; Marini, 1984; Maxwell, 1987; Murphy, 1987; Rindfuss et al., 1984; De Wit and Rajulton, 1992; De Wit, 1994) and with a postponement of second births indirectly through age at first birth (Wineberg, 1988; Rindfuss et al., 1980). Moreover, women with greater levels of education are more likely to remain childless to the end of the reproductive years (O'Connell and Rogers, 1982; Rao and Balakrishnan, 1988a; Grindstaff et al., 1981; Martinelle, 1993). There is also some evidence that higher education may affect the attainment of parities beyond the second birth (Kravdal, 1992; Wineberg and McCarthy, 1989).

Available evidence on the influence of women's economic activity suggests that employment may, as well, affect fertility outcomes (Ni Bhrolchain, 1986). Previous findings indicate that women at greatest risk of childlessness are those with professional occupational status (Grindstaff et al., 1989). Alternatively, however, women faced with choices between competing activities such as parenthood and market work may be inclined to initiate childbearing earlier in order to minimize career interruptions later on (De Wit and Rajulton, 1992; De

Wit, 1994). This latter set of results parallels findings on the determinants of marriage formation which conclude that women's employment and earnings may have a positive impact on the propensity to marry (Teachman et al., 1987; Lichter et al., 1992). In light of conflicting results, it is, therefore, of interest to attempt to establish the exact influence of paid employment on the timing of a birth.

As argued by De Wit and Rajulton (1992), the growing socioeconomic and cultural importance in recent years, especially for women, of education and employment as investments in human capital may actually work to increase their impact on fertility outcomes. If this is the case, then the timing of births for younger cohorts may depend increasingly on a growing number of alternative and often competing activities such as higher education and paid work. In effect, decisions regarding childbearing among younger women may be less a matter of normative prescriptions and more a question of the conscious balancing of personal achievements with more conventional role expectations. The observed delays and curtailment of fertility for many highly-educated women along with accelerated timing of births for those with early work experience may thus reflect, at least in part, the continuing difficulties faced by women in their attempts to achieve both family and non-family goals.

Using data from the 1984 Canadian Fertility Survey (CFS), De Wit (1994) examined the effects of women's educational attainment and previous employment activity on the timing of first and second births in Canada using the Accelerated Failure Time (AFT) model. This analysis demonstrated significant differences by parity as well as according to birth cohort, with education having a significant delaying impact on first birth timing which appeared to actually increase in magnitude over time; the influence of work status varied similarly over time but with an accelerating effect. Using Cox's (1972) Proportional Hazards model, the present research attempts to build on the earlier work by developing and testing similar models of first and second births with data obtained from the 1995 Canadian General Social Survey (GSS) (Statistics Canada, 1996). This permits not only an opportunity to replicate the earlier findings from the CFS with a comparable national-level sample, but also helps to determine empirically whether the earlier-observed cohort trends in the influences of higher education and labour force participation continue to operate among more recent cohorts. The possible interaction between economic activity and educational qualifications in influencing the decision-making process about childbearing is also investigated.

As in the previous studies, an analysis of birth timing is done separately for each cohort of women and according to parity in order to examine differences in timing and the relative importance of education as well as employment status. This framework for analysis of both parity groups permits examination of a possible interaction between education and employment, birth cohort (i.e., age) and birth timing. Whereas a formal test of this particular effect would require

including a multiplicative term involving education (or employment) and age group in the model, this procedure would tell us little about the structure of significant determinants, including education and employment status, for women from different generations. With the exception of the earlier work by De Wit and Rajulton (1992) and De Wit (1994), explicit treatment of the influence of maternal birth cohort on timing and risk of childbearing has received little attention. Nevertheless, as we demonstrate below, the cohort approach can help to provide insights into not only the time structure of significant demographic and socioeconomic changes, but also the content or nature of such change.

In addition to education and paid employment, the present study further considers a number of other variables shown previously to be correlated with the risk and timing of childbearing, including marital status, contraceptive use, region of residence in Canada, country of birth, and religious affiliation. Marital status at the time of a birth is expected to influence its timing since most births in Canada continue to occur within marriage, despite recent increases in the rate of single parenthood. Analysis of contraceptive use from the CFS reveals that, while almost 70 percent of Canadian women in the reproductive ages have used family planning methods, two-thirds of the sample practiced contraceptive use at some point prior to first birth (Balakrishnan et al., 1985). Evidently, the use of birth control in Canada reflects more the desire to time births than to avoid them altogether. Research on the influences of religious affiliation and religiosity suggests that religious affiliation may affect fertility behavior through value orientations towards certain behaviors such as contraceptive use and non-marital fertility (Balakrishnan and Chen, 1990). More traditional religions such as Catholicism presumably favour the maintenance of the large family ideal, which depends, in part, on the early initiation of childbearing. A greater risk of births should, therefore, be observed for Catholic as opposed to non-Catholic women. Differences according to country of birth are also examined in the context of this study in order to determine whether the historical pattern of early childbearing among Canadian-born women persists. Finally, region of residence in Canada may be important in predicting birth timing given long-standing differences in the fertility patterns of Quebec versus non-Quebec women. For further discussion on the substantive findings for these variables, see De Wit and Rajulton (1992).

In the section that follows, we discuss data characteristics and methodology for the study. We then turn to results of the bivariate (survival) analyses and multivariate hazards analyses, where we consider both main effects of all covariates in the analysis and finally, the interactions between level of education and previous employment.

Data and Methodology

The 1995 GSS is a large national-level probability sample ($n = 10,749$) with a target population of non-institutionalized Canadians, 15 years of age or older, and living in the ten provinces. Data were collected via telephone interviews between January and December, 1995. Sample weights have been applied to the data to adjust for non-uniform interprovincial sampling and non-response rates. The final response rate for the survey was 80.7 percent (for further details, see Statistics Canada, 1996).

The present analyses are based on birth history data of 3,519 women of all marital statuses who ranged in age from 18 to 49 at the time of the survey. The sample is stratified according to age or birth cohorts of the respondents, 18-29, 30-39, and 40-49, and only women exposed to the risk of a birth, either first or second, are included in the analysis. Proportions not at risk of a birth are determined according to whether the respondent reported surgical sterilization or some other medically diagnosed condition preventing a birth and represent 3.5 percent (124 cases) of the first parity group and 6.7 percent (150 cases) of the second parity group. Seventy additional cases were eliminated from the second parity analysis given a lack of information about fecundity status or the date of transition from first to second parity. The first parity group consists of 3,393 women who are at risk of at least one birth. Of these, 66 percent have had at least one birth by the survey date. Among the second parity group, nearly 77 percent of the total of 2,100 exposed to the risk of birth have actually experienced one by the time of the survey.

The Proportional Hazards model used here (Cox, 1972) combines multiple regression with life table techniques and is, therefore, ideal for examining both the structure of relevant predictors as well as the distribution of events according to time. The core assumptions of the Proportional Hazards model are that 1) heterogeneity is measured by the set of characteristics at the time of the event under study (as in ordinary regression); and that 2) the relative risks (i.e., the risk of experiencing the event relative to a pre-established reference group of individuals) are proportional to the baseline hazard over the duration of interest, which, in the present instance, is time elapsed since the starting age of fertility (around age 15) in the case of first birth, and the completion of a first birth, in the case of second birth.

A common problem with event history analysis is that of censoring or data truncation that is caused by the incomplete experience of respondents with respect to the event studied. As the sample includes women between the ages of 18 and 49, the effects of censoring on birth timing can be expected to be greatest for younger women. Whereas women nearing the end of the reproductive span have had considerable time in which to complete births, those in the youngest age groups will be least likely to have experienced a birth due to shorter

exposure time. Although life table techniques like hazards modeling can correct for censoring by combining the experience of those who have undergone the event with those who have not in the calculation of survival probabilities, cohort comparisons should nevertheless be regarded with some caution as no adjustment can be made for differences in the extent of censoring among the various birth cohorts. Since the analysis is based on cohorts with differing levels of censoring, being heaviest among the youngest cohorts of first and second parity women, results should be interpreted carefully. The problem of selectivity arises from the fact that those experiencing the behaviour of interest at an early age will be overrepresented in the sample. Thus, while women in the older cohorts have had time to complete births at all ages, the younger cohort can only have given birth at younger ages. The cohort framework for analysis adopted in this study helps to control for differences in exposure time among various age groups of women. Selectivity can also be due to the influence of various socioeconomic and demographic characteristics (i.e., population heterogeneity) which impact on the timing of a birth. As with multiple regression, however, the Proportional Hazards procedure can incorporate and simultaneously control for a number of explanatory variables.

To identify trends in birth timing, proportions experiencing the events of interest classified by the given demographic and socioeconomic characteristics were first estimated. The Proportional Hazards modeling technique (Cox, 1972) was then applied to the data in order to determine the factors which influence the rate at which women experience a birth. The application of this procedure yields estimates of the magnitude and direction of effects on timing of a given level of the independent variable as well as the relative risk of undergoing the transition of interest associated with a particular demographic or socioeconomic characteristic. Different levels of risk, higher or lower, imply differences in timing, either earlier or later, compared to that for an established reference group.

For k time-constant variables, Cox's Proportional Hazards model may be written as:

$$\text{Log } h(t) = a(t) + b_1x_1 + b_2x_2 + \dots b_kx_k$$

where $a(t)$ can be any function of time. Because this function does not have to be specified, the model is often described as partially parametric or semiparametric. The semiparametric nature of the Cox regression model can be seen as advantageous over the fully-parametric Accelerated Failure Time model since, with the Cox model, it is not necessary to make assumptions about the underlying time distribution of the event(s). While explicit examination of the underlying distribution of events can increase our understanding of the process of interest, empirical information on the shape of this function is, in practice, often lacking, making it difficult to choose appropriately from among the various mathematical functions available with the AFT procedure. The Cox

model is referred to as the proportional hazards model because for any two individuals at any point in time, the ratio of their hazards (h) is a constant. Formally, for any time t , $h_i(t)/h_j(t) = c$ where i and j refer to distinct individuals and c may depend on explanatory variables but not on time.

As noted above, parameter estimates for the covariates reflect negative or positive effects on the timing of the event under study. In unstandardized form, the coefficients mean that a unit change in the level of the covariate involves a given change (either positive or negative) in the hazard, controlling for other variables in the equation. Exponentiating the coefficients (i.e., raising the number e (approximately 2.718) to the power of the parameter estimate value) in order to fix the baseline or reference category to one, coefficients less than one work to decelerate the timing of the event by a given factor, while coefficients greater than one are interpreted as having an accelerating effect on timing. These exponentiated parameter estimates, therefore, yield numerical multipliers of risk per unit difference in the predictor (Singer and Willett, 1991). For dummy variables in particular, the exponentiated coefficient gives the relative hazard for the groups corresponding to values of the dummy variable, controlling for other predictors.

Given that the unstandardized coefficients depend on the metric of the variable, it is also instructive to examine the t -statistics for the null hypothesis that each coefficient is zero. In moderate to large samples (such as those employed here), the t -statistics can be interpreted like those in an ordinary multiple regression and represent the ratios of the estimates to their standard errors (Allison, 1984). These give some indication of the magnitude of the effects in the model. As in ordinary regression analysis, standardized coefficients can also be computed for purposes of comparing the relative magnitude of the variables' effects on the hazard rate within the same population. However, the standardization of the parameter estimates in the present context is not very useful since the distribution of responses on categories of the independent variables are restricted between zero and one by dummy coding, whereas the dependent variable, the hazard, is analyzed in continuous form. As a result of this categorization, the standard deviations of the independent variables needed to calculate the standardized coefficients are affected. For this reason, the t -statistics associated with each of the parameter estimates are particularly useful for making within-model comparisons of magnitudes of effects. The unstandardized coefficients are also helpful, but mostly in comparing the magnitude of a given variable's effects across the study samples given that, in each case, the scale of the variable will be constant. Overall, it is most convenient to interpret the exponentiated coefficients in order to look at the influence of a given covariate on survival time rather than on the natural logarithm of survival time, as given by the unstandardized coefficients.

Due to data unavailability, many studies do not make use of characteristics measured at the appropriate time point, relying instead on the survey-date

measures. For example, Rao and Balakrishnan (1988b) in their analysis of CFS data utilize level of education at the survey date as a covariate of first birth timing. In our study, with the exception of religion and region of residence (for which no alternative measures were available), and country of birth, all relevant covariates are estimated for the time of first or second birth. Religiosity, as measured via frequency of church attendance, was initially considered for inclusion in the analyses. However, given the relatively large percentage of missing cases on this variable (approximately 17 percent) and the fact that this measure was given as of the survey date (and not at the time of a birth) even though religious commitment is well-known to vary over the life course, the decision was made to exclude it from the final models. The result of doing so for the remaining parameter estimates was negligible. In the case of educational attainment, prior employment, marital status and contraceptive use, estimated values at the time of first and second births were determined for specific categories of women according to birth order and cohort (see Appendix for details on the estimation procedures for these variables). Along with educational attainment and employment status, the control variables for the analysis are all measured at the categorical level. Dummy variables are especially useful in helping to detect any non-linearity in the relationship between education and birth timing (see Vuchinich et al. (1991)).

As with the earlier paper by De Wit (1994), education refers to years of formal education completed by the time of first or second birth, and is measured according to less than 9 years (the reference category), 9 to 11 years, 12 to 13 years, and 14 or more. This categorization scheme matches that used in the earlier comparable studies of CFS data and is intended to correspond with the meaningful distinctions between incomplete versus complete secondary school attendance, as well as between completion of secondary school versus post-secondary attendance; the adopted coding scheme also aims to ensure adequate numbers within categories while at the same time preserving as much as possible of the underlying continuous metric of the variable. Employment status as of the time of a birth refers to whether a woman had ever initiated full-time regular employment at some point prior to a birth. Women not commencing full-time work are considered as the reference category. Contraceptive use refers to whether or not family planning practices had ever been initiated prior to having a birth. Those not initiating use are treated as the reference group. Marital status at the time of a birth is measured according to the categories 'cohabiting', 'married', 'previously married', and 'single and never married', the reference category. Religious affiliation includes the categories 'Protestant', 'Catholic', 'Other or None' with Catholics as the reference group. Place of birth consists of the categories 'Europe or North America outside Canada', 'Canadian' and 'other foreign country', with Canadian-born women as the reference group. Region of residence is measured via 'West', 'Ontario', 'East' and 'Quebec', with Quebec as the reference.

The dependent variable in the first birth analysis is measured in years and months and may be interpreted as survival time from age 15 in the childless state before experiencing a first birth. Throughout this interval, individuals may either survive the event up to the survey time, in which case they are said to be censored, or experience a birth. In the analysis of second birth, survival time is measured in years and months from the date of first birth.

Results and Discussion

Table 1 presents numbers and percentages of women experiencing first and second births, classified by age group, education and employment status, as well as the other covariates, at the time of a birth. Table 2 gives the relative risks of first and second births according to cohort associated with the various predictor variables. Although shown in Table 2, discussion about covariates other than education and employment status is not reported here. In Table 3 we present the results of analyses of the interaction between education and prior employment. Final sample sizes in the multivariate analyses are based on list-wise deletion of missing cases.

Considering the bivariate relationships between education, employment status and birth timing in Table 1, the results can be summarized as follows: Concerning first the overall relationship between education and childbearing, regardless of age group, there appears to be a strong negative linear association between level of attainment and proportions having a birth, whether first or second, with higher education associated with lower proportions. Thus, for example, among the youngest cohort (18 to 29) of first-parity women, approximately 67 percent in the lowest-education category (less than 9 years) report a birth, compared with about 58 percent in the 9 to 11 years category, 29 percent in the 12 to 13 years group, and only 19 percent among the most highly-educated group with 14 or more years. Comparing the two parity groups on the proportion having a birth in the highest education category, we also see that a large proportion of women in the sample who have one birth also go on to have two, even among the youngest group with the shortest exposure time. Thus, only 19 percent of women in the youngest age group with the highest level of education have initiated childbearing, as compared to almost 34 percent of this same group who have gone on to have a second birth. In general, a higher proportion of women at any level of education having one birth go on to have two as compared to proportions initiating childbearing. Proportions of women having a second birth appear to be less strongly influenced by educational attainment than proportions having a first, as suggested by the comparatively larger numbers having a second birth, regardless of cohort or level of education.

Second, as is evident for both parity groups, the proportions having a birth with a given level of higher education are generally diminishing over time or across

Table 1
Number and Percentages of Women Experiencing a Birth by Age,
Parity, Educational Attainment, Employment Status and
Selected Demographic and Socioeconomic Characteristics, Canada: 1995

Cohort	First Birth						Second Birth					
	18-29		30-39		40-49		18-29		30-39		40-49	
	#	%	#	%	#	%	#	%	#	%	#	%
Education												
9-11 years	101	57.8	182	91.7	155	93.4	49	54.1	145	88.2	141	95.1
12-13 years	123	28.9	315	79.6	188	89.9	63	47.5	246	81.8	143	94.1
≥14 years	99	19.1	360	65.0	221	79.0	47	33.8	258	66.9	182	84.6
<9 years	62	67.2	179	97.0	234	95.9	26	82.7	111	93.8	176	96.2
Work Status												
Working	256	81.5	814	96.2	607	96.2	124	62.0	615	92.4	471	96.1
Not Working	129	13.3	224	43.0	205	71.2	61	35.6	147	45.9	178	81.3
Contraceptive Use												
Use	61	13.4	80	38.2	13	42.6	28	20.4	66	43.6	15	67.4
No Use	324	43.2	958	85.0	799	90.3	157	63.8	696	85.3	634	92.8
Marital Status												
Cohabiting	108	38.2	132	62.6	76	80.9	44	48.4	80	65.6	55	91.4
Married	135	55.3	544	85.5	466	93.5	88	55.4	465	79.5	406	93.2
Prev. Married	43	68.6	214	93.0	209	95.2	17	49.0	145	83.3	155	90.8
Single/N.M.	99	14.7	148	51.1	61	54.7	36	34.7	72	71.0	33	77.6
Religion												
Protestant	117	33.9	346	78.8	279	88.9	60	52.8	258	81.5	229	94.3
Other/None	71	27.5	166	70.9	120	84.3	29	37.8	117	74.3	95	91.4
Catholic	188	30.4	497	78.3	390	91.0	91	49.7	371	77.9	309	90.7
Place of Birth												
Europe/NA	22	41.9	62	82.2	69	85.9	11	52.3	47	73.8	51	82.7
Other Foreign	36	34.9	96	77.8	62	86.4	20	55.0	63	72.2	50	90.1
Canada	327	29.6	880	76.3	681	89.3	154	47.0	652	79.1	548	93.5
Region of Residence												
West	119	29.5	338	77.7	287	88.9	63	49.7	257	81.8	229	90.3
Ontario	76	31.7	202	74.8	152	87.4	34	46.4	150	76.1	132	95.0
East	83	37.3	197	80.1	142	93.5	32	37.3	137	73.9	116	91.7
Quebec	107	29.0	301	78.7	231	88.2	56	55.5	218	77.0	172	88.8
# of Births	388		1043		819		187		766		654	
Sample Size	1061		1380		952		376		992		732	
% Births	37		76		86		50 (18)		77(59)		89(77)*	

*Unbracketed percentages are for those women at risk of either a first or second birth; thus, unbracketed percentages for second-parity women represent the total proportion of women within a given age cohort who, having one birth, go on to have two. Bracketed percentages, conversely, refer to the total proportion of women within each age group who, by the survey date, have given birth to a second child.

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cohorts: In particular, although the proportions having a first birth at the lowest level of education are fairly similar for the three cohorts of women (from youngest to oldest: 67.2, 97.0 and 95.9 percent) the negative effect of education appears most pronounced for the younger groups at the highest level (from youngest to oldest: 19.1, 65.0 and 79.0 percents). A similar pattern is also observed for the second parity sample where, among those with the least education, proportions having a second birth from youngest to oldest are 82.7, 93.9 and 96.2, respectively, whereas lower proportions of the most highly-educated women go on to have a subsequent birth, with the proportions being smallest for the youngest cohort (33.8, 66.9 and 84.6 percents, from youngest to oldest, respectively). These trends appear to reflect a growing impact of higher education on childbearing decisions.

The bivariate analysis of the effect of work status on timing of a birth suggests uniformly that full-time regular employment is associated with an increase in the proportion of women having a birth, whether first or second. For the first parity group, the differences in proportions seem most pronounced for the youngest cohort and least so for the oldest one (i.e., differences between working and not working categories of 68.2, 53.2 and 25 percents from youngest to oldest cohorts, respectively). However, a similar trend is not apparent with the second parity group.

From the multivariate results in Table 2, it is apparent that a number of variables may affect the rate at which women from various age cohorts have a first or second birth. Among the most notable effects are those of employment status and educational attainment, which appear to influence timing of childbearing quite dramatically and in a fairly consistent manner. Employment status has an impressive and systematic influence on birth timing, insofar as employment prior to a birth appears to increase the risk of first and second childbearing for all age groups except the oldest group of second parity women. In the case of first birth, paid employment increases the likelihood by a factor of 4.64 among the youngest group, by 2.50 among the middle age group and by 1.30 times among the oldest women; with a second birth, the corresponding figures are a 113 percent increase in the risk for the most recent cohort, a 138 percent increase for the middle age group, and a nonsignificant change in risk for the oldest cohort. Overall, the results found with the CFS data are similar to those obtained here, with paid employment prior to a birth leading to an acceleration of the process (see De Wit and Rajulton, 1992); and as posited previously, the strong positive effect on the risk of births among younger cohorts in particular may be due to women's growing propensity to minimize career disruptions associated with childbearing.

As was also found in the earlier cohort analyses of the CFS data (see De Wit and Rajulton, 1992; De Wit, 1994), the results in Table 2 indicate that, for women of all age groups, education at first birth exerts a strong and significant negative impact on the risk of childbearing, net of the other explanatory variables. A

very similar pattern of significant effects is observed for the model of second birth timing. Indeed, for all of the age and parity groups considered, education at the time of a birth has a powerful impact on the risk of childbearing and is, in every case, judging by the size of the *t*-values, comparable in the magnitude of risk reduction to the effects of contraceptive use. This is a remarkable finding when considering that education is often found to be associated strongly with other indicators of women's status such as contraceptive use and paid employment; that its net effect, therefore, in the present study is still significant and substantial is quite noteworthy. In contrast to a factor like contraceptive use, educational achievement is more readily conceptualized as a non-proximate factor which affects fertility behaviors primarily through the set of direct biological determinants (see Bongaarts, 1978).

Table 2 indicates further some notable cohort differences in the risk of first and second births among those with higher levels of educational attainment. Quite clearly, the greatest influence is found among the youngest group, compared to the two older cohorts. Thus, the most highly-educated group of first parity women in the youngest cohort are at 92 percent lower risk than their least-educated counterparts, whereas the corresponding risk reductions among the middle and oldest age groups are 71 and 54 percents, respectively. Among second parity women, the youngest cohort with the highest educational attainment experiences a 91 percent lower risk; the middle age group, a 79 percent reduction; and the oldest group, a 56 percent reduction in risk. The results for both parity groups essentially reflect an interaction between education and birth cohort, even though no formal test of this particular effect (i.e., between education and age on birth timing) is performed in the context of this study. Comparing the overall results on the timing of first and second births, the initial transition to parenthood among the cohorts under study does not appear to be any more strongly affected by educational achievement than is the occurrence of a second birth. This observed pattern of results by parity is consistent with previous research suggesting that attained education may also exert a significant delaying effect on the timing of higher order births through age at initiation of childbearing (see Wineberg, 1988). The results do contrast, however, with those of De Wit (1994) which indicated a stronger relative impact of educational achievement on the timing of the first birth. The apparent emergence in the present study of a significant influence of higher education on the timing of second births could mean that obstacles to fulfilling birth timing intentions among Canadian women actually have longer-term effects than may have been the case for earlier cohorts of highly-educated women (i.e., those surveyed 11 years earlier).

The finding for the increasing significance of higher levels of educational qualifications among younger women, regardless of parity attainment, points to the changing relationship over time between women's education and their fertility behaviour and is consistent with the growing necessity of higher education as a basis for occupational entry and career progression. In

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Table 2
Relative Risks of a Birth by Age, Parity,
Educational Attainment, Employment Status and Selected
Demographic and Socioeconomic Characteristics, Canada: 1995

Cohort	First Birth			Second Birth		
	18-29	30-39	40-49	18-29	30-39	40-49
Education						
9-11 years	0.56 **	0.82	0.82	0.34 ***	0.57 ***	0.70 **
12-13 years	0.17 ***	0.47 ***	0.74 **	0.16 ***	0.34 ***	0.74 **
14 years	0.08 ***	0.29 ***	0.46 ***	0.09 ***	0.21 ***	0.44 ***
<9 years ^a	1.00	1.00	1.00	1.00	1.00	1.00
Work Status						
Working	4.64 ***	2.50 ***	1.30 **	2.13 ***	2.38 ***	0.98
Not Working ^a	1.00	1.00	1.00	1.00	1.00	1.00
Contraceptive Use						
Use	0.47 ***	0.35 ***	0.26 ***	0.31 ***	0.34 ***	0.38 ***
No Use ^a	1.00	1.00	1.00	1.00	1.00	1.00
Marital Status						
Cohabiting	1.10	1.35 *	2.19 ***	1.35	1.18	1.36
Married	1.47 **	1.78 ***	2.75 ***	1.51	1.40 *	1.35
Prev. Married	2.00 **	1.80 ***	2.40 ***	1.96	1.37	1.23
Single/N.M. ^a	1.00	1.00	1.00	1.00	1.00	1.00
Religion						
Protestant	1.04	0.98	0.81 *	1.05	1.05	1.08
Other/None	0.97	0.92	0.71 **	0.87	0.82	0.79
Catholic ^a	1.00	1.00	1.00	1.00	1.00	1.00
Place of Birth						
Europe/NA	0.76	1.09	0.82	0.69	1.03	0.59 ***
Other Foreign	0.98	0.78 *	0.92	0.96	0.90	0.87
Canada ^a	1.00	1.00	1.00	1.00	1.00	1.00
Region of Residence						
West	1.44	1.24	1.42 *	1.24	1.25	1.29
Ontario	1.12	1.16	1.47 ***	1.12	1.27 *	1.32 *
East	1.35	1.25 *	1.38 **	1.58	1.26 *	1.18
Quebec ^a	1.00	1.00	1.00	1.00	1.00	1.00
Sample Size	1031	1335	899	365	960	700
% Censored	69	23	11	52	22	8
Global X²	684.58 ***	610.80 ***	219.13 ***	162.70 ***	382.60 ***	139.77 ***
	(df=15)	(df=15)	(df=15)	(df=15)	(df=15)	(df=15)

^a Reference category for the variable.

* level of significance .05
 ** level of significance .01
 *** level of significance .001

addition to the increased importance of higher levels of educational attainment over the past few decades, qualitative changes in the content and social significance of education may also help to explain substantial cohort variations in birth timing. These changes in particular are evident in comparing the risks of a birth for women from different age groups, holding constant the level of education. For example, among all women with 14 or more years of education, those from the youngest cohort in the first birth analysis are at 38 percent lower risk of childbearing than the oldest cohort; similarly, the youngest cohort of women in the second birth analysis is at 35 percent lower risk than the oldest group. Only with the use of the cohort framework in our study have we been able to quantify empirical variations in risk of a birth for women of different age groups according to characteristics like education and work status.

Considering finally the interaction between education and employment status (shown in Table 3), this effect as well appears to be more greatly accentuated for younger cohorts of both parity groups, suggesting that the impact of the joint influence of women's labour force involvement and education is also increasing over time, along with the main effects. Note, however, that in models containing significant interaction terms, the meaning of the associated main effects are unclear so that it is not useful to interpret these associated main effects, regardless of their magnitude or statistical significance. Also note that, while included in the analyses, main effects other than education and work status are not presented in Table 3. Given that the reference category for the education variable is women with fewer than 9 years, and for employment status, women with no prior employment, in the analyses involving interaction effects the appropriate comparisons are between women at a given level of education with previous employment, compared to those without prior work history. Thus, among the youngest cohort of first-parity women, the combined impact of previous employment and 12 to 13 years of formal education elevates the risk of a birth by 8.6 times, compared to women not employed prior to a birth with the same level of education; the corresponding risk increase for those with 14 or more years of education is 27 times. For the 30 to 39 age group, the results are somewhat less striking, with significant increases by almost 3 times and 9 times among those with 12 to 13 years and 14 or more years of education and no prior employment. Among the 40 to 49 year old cohort, prior full-time employment combined with 12 to 13 years of education increases the risk of a birth by 54 percent, whereas among those with 14 or more years of education and previous employment, the risk increases by 3.8 times.

In the case of second-parity women, having 12 to 13 years of education by the time of birth combined with prior full-time work experience increases the risk by 2.7 times whereas an increase in risk by almost 19 times is experienced by women having attained 14 or more years of schooling along with prior work. In a similar manner, women in the middle age cohort with 12 to 13 years of education along with previous employment are about 2.5 times more likely to have a birth while women with the highest possible educational attainment and

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Table 3

**Relative Risks of a Birth by Age, Parity, and the Interaction between
Educational Attainment and Employment Status, Holding Constant Selected
Demographic and Socioeconomic Characteristics, Canada: 1995**

Cohort	First Birth			Second Birth		
	18-29	30-39	40-49	18-29	30-39	40-49
1. Main Effects						
Education						
9-11 years	0.73	0.60 **	0.74	0.43 *	0.63	0.67 *
12-13 years	0.06 ***	0.16 ***	0.44 ***	0.09 ***	0.14 ***	0.65
≥14 years	0.01 ***	0.04 ***	0.13 ***	0.00 ***	0.04 ***	0.23 ***
<9 years ^a	1.00	1.00	1.00	1.00	1.00	1.00
Work Status						
Working	1.25	0.70	0.79	0.71	0.82	0.76
Not Working ^a	1.00	1.00	1.00	1.00	1.00	1.00
2. Education x Employment						
9-11 years	1.38	1.05	0.89	0.82	0.76	0.81
12-13 years	8.55 ***	2.79 ***	1.54 **	2.66 *	2.47 ***	0.94
≥14 years	26.78 ***	9.10 ***	3.81 ***	18.84 ***	7.34 ***	1.70 **
<9 years ^a	1.25	0.70	0.79	0.71	0.82	0.76

* level of significance ≤ .05

** level of significance ≤ .01

*** level of significance ≤ .001

previous full-time work experience exhibit over 7 times the risk of their non-working counterparts. Among the oldest cohort of second-parity women, the risk is increased significantly by 70 percent among those with 14 or more years of education and previous full-time employment.

Summary and Conclusions

This analysis found that higher education and previous labour force activity exert a powerful influence on birth timing for women of all age and both parity groups. As with the earlier results using CFS data (see De Wit and Rajulton, 1992; De Wit, 1994), cohort variations are also evident, with the overall pattern involving a relative increase across cohorts in the magnitude of effects of both higher educational attainment and prior employment. A major contribution of

the present paper has been to replicate the original findings and support the earlier-positied increase in effects over time.

In the case of educational achievement, observed cohort trends may be interpreted as resulting from changes in the content and socioeconomic significance of formal schooling over the past few decades which have accentuated the delaying impact of educational achievement on women's timing of childbearing. Moreover, unlike the previous work (using CFS data) which also included an examination of second births (De Wit, 1994), the present research finds very similar patterns of effect for both parity groups. The findings obtained here on the impact of women's paid employment are generally quite similar to those observed with the CFS data; taken together, patterns of variation in the influence of education and paid employment suggest that the earlier-observed trends persist among younger cohorts of Canadian women.

One possible interpretation of the general patterns of cohort differences documented above is that, although parenthood continues to be normative, the occurrence of births for younger cohorts of Canadian women is increasingly contingent on personal decisions and achievements such as higher education and paid work. Owing to more restrictive gender role definitions and fewer non-family opportunities for earlier generations of women, it is likely that higher education and market activity in the past interfered less with women's childbearing careers than is apparent for more recent cohorts. This factor is reflected in the results of the present study which finds a greater decelerating effect on birth timing of educational achievement for younger cohorts. At the same time, the increasing accelerating effect of full-time employment suggests similarly that contemporary cohorts of childbearing age are more influenced in their fertility behaviors by market opportunities than were previous cohorts.

A final point for consideration is the apparent interaction between education and employment. Analyses of the present data suggest that, net of other covariates in the model, prior work experience combined with higher educational achievement significantly and dramatically increases the likelihood of a birth. At higher levels of education, women with prior work experience and thus presumably greater commitment to paid employment appear to be more highly motivated to reduce later career interruptions than their less-educated and possibly less career-minded working counterparts. More highly educated women presumably also enjoy higher incomes, a factor which could accelerate birth timing due to their greater ability to afford formal childcare. Future analyses should consider further possible interpretations of the observed associations, their interplay and changing nature.

Appendix

Estimation Procedures for Variables Measured Indirectly as of the Time of First and Second Births¹

Estimation of Level of Education at the Time of First and Second Births:

The GSS data contained information on the number of years of education as of the survey date, measured in single years up to grade 13, and grouped, thereafter, according to categories of highest completed attainment (specifically, MA or PhD, BA or Teaching Certificate, Community College Diploma, some university, some college, and some trade, technical, vocational or business schooling). On the basis of this information, likely durations of schooling beyond high school were computed. Thus, for example, a graduate degree was estimated to require a further five years beyond high school completion; a college diploma was assumed to involve two additional years; some college or other vocational school, one year, etc. A measure of education in single years at the time of the survey was constructed on the basis of combined information about elementary and secondary attainment and post-secondary attainment. The resulting measure was then employed as an estimate of the respondent's survey-date years of education. Information on the age at completion of studies was also included in the data set. The data on timing of first and second births, timing of completion of studies, and (estimated) education at the time of the survey were examined and likely estimates of the number of years of schooling at the time of first and second births for specific categories of women according to parity were determined as follows:

1. For women without a birth by the survey date (censored cases), education was set to be equal to the survey-date measure of education; this category consists of 1,143 cases in the first birth analysis and 493 cases in the second birth analysis;
2. For women having a first (2,250 cases) or second birth (1,607 cases) by the survey date (uncensored cases):
 - a) For those completing their education by the survey date (2,097 in the case of first birth and 1,516 in the case of a second birth):
 1. Among those having a birth before completing their education (403 first births and 245 second births), education at the time of a birth was estimated to be equal to their survey-date level of education minus the difference between the respondent's age at completion of education and age at the time of childbearing; and,

- ii. Among those having a birth after completing their studies (1,678 first births and 1,260 second births), education at the time of birth was set to be equal to education completed at the survey date.
1. For those not completing their education by the survey date (153 cases in the first birth analysis and 91 cases in the second birth analysis), education at birth was estimated as the survey-date education minus the difference between the respondent's age at the survey and the age at birth².

Estimation of Initiation of Full-time Regular Work Prior to the Time of First and Second Births:

The GSS data also contained information on the age at which first regular full-time employment began. The date at initiation of first regular work was compared to the age at which a first or second birth occurred in order to determine a woman's employment status prior to a birth. The various categories were assigned as follows:

1. Women not giving birth to a child, whether first (115 cases) or second (143 cases), by the survey date, but indicating that they had ever worked at a job or business on a regular basis, were considered to have initiated full-time regular employment before the time of a birth; those without a birth and without previous employment by the survey date were considered as not working at the time of birth (1,029 in the case of first births and 337 in the case of second birth);
2. Among those having a first or second birth by the survey date, women whose date of first regular employment preceded the date of birth, were categorized as having initiated full-time regular work before the birth of a child (1,673 first births and 1,209 second births). Conversely, women who had yet to begin working by the survey date or who initiated regular work after the birth were classified as not employed full-time at any point prior to a birth (494 first births and 340 second births).

Estimation of Initiation of Contraceptive Use Prior to the Time of First and Second Births:

Information in the GSS data on contraceptive use included questions about weeks of prior use by the respondent or her spouse/partner. From this variable, it was possible to estimate the women's age at first contraceptive use by first

converting the number of weeks of use into years and then subtracting years of use from the survey-date age. In order to determine whether an individual had initiated contraceptive use at some point prior to a first or second birth, women were classified as follows:

1. Women not giving birth to either a first or second child by the survey were considered to be contracepting prior to a birth if they indicated use by themselves or their partner at any time prior to the survey (564 first birth cases and 223 second birth cases); those not initiating family planning and not giving birth to a first child or a second child comprised 522 and 246 women, respectively;
2. Those completing a first or second birth were considered to have initiated use of family planning methods prior to a birth if their estimated age at first use was less than or equal to their age at the birth of the child (157 first births and 109 second births); those whose age at a birth was less than the age at which first contraceptive use took place were considered as not contracepting prior to a birth (394 first births and 212 second births); those completing a birth but reporting no previous contraceptive use at any time comprised 1,635 first parity women and 1,252 second parity women.

Estimation of Marital Status at the Time of First and Second Births:

In order to estimate the marital status of the woman at the time of birth, whether first or second, information on respondents' ages at union formation and dissolution were compared with the age at birth of a child, resulting in the following categories:

1. Among women giving birth prior to the survey date:
 - a) Married women were those who had entered into a legal union of any order before the birth of a child and had not experienced marriage dissolution (due to either separation, divorce, or death) prior to the child's birth (1,145 first births and 959 second births);
 - b) Women living common-law at the time of birth were determined to be those who had entered into a common-law union of any order prior to the birth of a child and had not experienced dissolution due to either separation or death before having a birth (317 first births and 179 second births);
 - c) Previously married women were those who had experience marital dissolution due to separation, divorce or death and had not remarried or entered into a first or subsequent common-law union

by the time of childbearing (471 first births and 325 second births); and,

- d) Women who were single and never-married at the time of a birth were those who had never been married and were not living in a common-law union at the time of a birth (317 first births and 270 second births).
2. Women not having a birth, whether first or second, by the time of the survey, were assigned their current marital status. This applied to 1,143 women in the first birth analysis and 493 women in the case of second birth.

Endnotes:

1. Small discrepancies between category totals and estimated figures are due to estimation errors since missing information precludes determination of the exact status of some respondents with respect to the variables measured.
2. For all those having a birth before completion of studies, it is assumed that attendance is continuous and full-time, except in cases where the time elapsed between the date of a birth and the date of the survey is greater than or equal to total years of education attained by the survey date. Thus, in cases where the difference is equal to or greater than total years of education, education at the survey date is reduced by one-half of the interval to arrive at estimated levels of education at the time of birth, or by only one-quarter where the difference is equal to at least twice the total years of education at the survey date. Although this approximation is based on untestable assumptions, it is also the case that such assumptions apply mostly to older women with a prolonged interval between childbearing and school completion; any potential distortions in estimation, therefore, affect only a relatively small number of cases (i.e., 153 first birth cases and 91 second birth cases).

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