NEW EVIDENCE ON THE HUMAN CAPITAL EXPLANATION OF THE FAMILY GAP

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Abstract

This paper contributes to the literature on the family gap, that is, the income difference between women with children and women without children, in Canada. Using the most recent microdata -- the 2001 Statistics Canada General Social Survey (GSS 2001) -- I find new evidence that a family gap does exist in Canada. The child penalty decreases when I control for actual work experience and the duration of work interruptions. This suggests that when women spend a long time out of the labour market, they not only gain less human capital but also depreciate their existing human capital. However, human capital losses and depreciation explain only part of the family gap, not all. When I distinguish maternity leave and child care interruptions from demand-related interruptions, health-related interruptions and other interruptions, I find that human capital depreciation occurs during the work interruptions, but there is no significant difference in the effects of different types of interruption. Thus decomposing the total duration more accurately does little to explain the family gap.

I. Introduction

A number of papers have focused on the family gap, that is, the income difference between women with children and women without children, in the United States and the United Kingdom, but much less research has been concentrated on the family gap in Canada. There exists a family gap between mothers and non-mothers in Canada, although it has become smaller over time. For instance, in 2001 mothers who were 25 to 45 years old and worked full-time for a full reference year earned 93.1 per cent of the income of non-mothers, while in 1995 mothers received 87.3 per cent of the income of non-mothers. The situation of fathers has changed too, though there is still a premium for men associated with having children. Fathers receive 16.4 per cent more income than non-fathers (aged 25 to 54, worked full-time for a full year) in 2001, compared with 33.6 per cent more income in 1995.[1]

In an attempt to explain this family gap, researchers have used econometric methods to see if the gap persists after controlling for such variables as actual experience, education, and region of residence, etc. Phipps, Burton, and Lethbridge (2001) examined the family gap between mothers and non-mothers using data from the 1995 Statistics Canada General Social Survey (GSS 1995). I follow their approach and focus on finding new evidence of a family gap in Canada from the view of human capital attainment to see if anything has changed since 1995. In this paper, I use the 2001 Statistics Canada General Social Survey (GSS 2001) to address this question. As did Waldfogel (1998a), Budig and England (2001), and England and Budig (1999) for the United States and the United Kingdom and Phipps, Burton, and Lethbridge (2001) for Canada, I find that considering time out of the labour market only accounts for part of the family

[1] To derive these comparisons, I simply computed mean incomes for men and women with and without children from the GSS 2001. The data for 1995 are based on raw estimates from the GSS 1995 (Phipps, Burton, and Lethbridge 2001). In 2001, women earn 69.3 per cent as much as men, and women with children earn 67.6 per cent as much as men; women without children earn 72.6 per cent as much as men.

gap, not all. Furthermore, I add occupation as an additional explanatory variable to the OLS regression model used in Phipps, Burton, and Lethbridge (2001) to make it more reasonable to explain the child penalty on the income of mothers. In most specifications, the coefficients of the occupational dummies are statistically significant. I find that a family gap does exist in Canada even after controlling for actual work experience, marital status, education, province of residence, and occupation.

Five possible explanations for the persistence of the family gap have been proposed in the literature (see Waldfogel 1997a; England and Budig 1999; Phipps, Burton, and Lethbridge 2001): (1) women with children are more likely to spend time out of the labour market, caring for their children, than other women, and thus acquire less human capital and depreciate their existing human capital; (2) since traditionally women always play the roles of housekeeper and caregiver to elders and children, they may put too much effort into household production, which means less energy can be spared for their paid work; (3) the institutional structure of the labour market, including some family policies such as maternity leave and childcare, may also have an effect on the family gap; (4) employers may discriminate against women with children; and (5) the wage penalty of motherhood maybe due to the existence of unobservable heterogeneity and endogenous work interruptions or fertility.

Because of data limitations, I cannot examine all the above explanations of the family gap in Canada. This paper focuses on re-examining the part that has received the most attention -- human capital theory -- as an explanation for the earning penalty of motherhood in Canada.[2] I find that women who take a long time out of the labour market not only gain less human capital

[2] There is a family gap of 0.10 in the mean log of income in Canada. The results in specification 4 in Table 10 show that the coefficient of child penalty is 0.045 after controlling actual experience, education level, province of residence, occupation, duration of work interruption, and marital status. So the human capital theory helps explain 55 per cent of the family gap.

but also depreciate their existing human capital. I also find that distinguishing maternity leave and child care interruptions from demand-related interruptions, health-related interruptions and other interruptions does little to account for the family gap between mothers and non-mothers. Most of results are consistent with the findings of Phipps, Burton, and Lethbridge (2001).

This paper proceeds as follows. In section II I review the related research and focus on three aspects of the family gap: explanations, econometric models, and magnitude. In section III I describe the data from the 2001 Statistics Canada General Social Survey (GSS 2001) and the method I use to estimate the family gap. In section IV I present the estimation results and discuss three related questions. First, is there any family gap in Canada? Second, does time out help to explain the family gap? Third, do child-related work interruptions have an effect on the family gap? Section V summarizes my conclusions.

II. Literature Review

Many researchers have focused their efforts on the "gender gap" -- that is, the pay gap between men and women -- while relatively little research has focused on the "family gap," which is the pay gap between women with children and women without children. Generally, women take more responsibility for rearing children and other household tasks than men, which implies that women with children spend more energy and hours on unpaid housework and caring for children than women without children. So it is not surprising that on average mothers earn less than other women, and far less than men. Thus a good understanding of the "family gap" may help explain the "gender gap." In this section I will first summarize some explanations for the family gap, then present econometric models, and finally give the magnitude of the family gap in different countries and different years.

a. Explanations for the family gap

A number of papers have focused on the family gap in the United States and the United Kingdom (e.g., Korenman and Neumark 1992; Neumark and Korenman 1994; Waldfogel 1997a,b, 1998a,b; Lundberg and Rose 1999; England and Budig 1999; Joshi and Newell 1989; Joshi, Pace, and Waldfogel 1999; and Harkness and Waldfogel 1999), but much less research has been concentrated on the family gap in Canada. Phipps, Burton, and Lethbridge (2001) are the first authors to step into the field. Besides in the United States, the United Kingdom, and Canada, a family gap also exists in Australia (Baxter 1992), Germany, Finland and Sweden (Harkness and Waldfogel 1999). Five possible explanations for the persistence of the family gap have been proposed (e.g., Waldfogel 1997a; England and Budig 1999; Phipps, Burton, and Lethbridge 2001).

First, the human capital explanation, which was introduced by Becker (1985) and has received the most attention, is an important potential explanation for the family gap. Women with children are more likely to spend time out of the labour market, to care for their children, than women without children. The more time out women take, the less human capital -- such as education and work experience -- they acquire, and the less pay they receive. An early study in the United States by Hill (1979) concluded that all the negative wage effects of motherhood disappeared after controlling for actual work experience,[3] but some subsequent studies have found that only part of the negative effect of children on mothers' pay is explained by human

[3] Hill (1979) used as independent variables formal education, whether living in the South, city size, experience (=age - education – 6), work hours, health, marital status, and number of children. After Hill refined her model by controlling for potential work experience, training, and measures of labour force attachment such as absences from work, the number of children effect on the wages of white women with children became insignificant.

capital, not all (Korenman and Neumark 1992; Waldfogel 1995, 1997a,b, 1998a; England and Budig 1999; Jacobsen and Levin 1995; Baxter 1992; Joshi and Newell 1989).

Using Canadian data, Phipps, Burton, and Lethbridge (2001) examined the effect of time out of the labour market on mothers' current incomes and concluded that controlling for time out only reduces the family gap, but does not eliminate it. They also found that absence from work led to not only human capital loss but also human capital depreciation as found in the United States and the United Kingdom (Jacobsen and Levin 1995; Mincer and Ofek 1982; Stewart and Greenhalgh 1984; and Stratton 1995).[4] It is important to note that Phipps, Burton, and Lethbridge (2001) studied other aspects of work interruptions besides their duration. They found that:

> When we go on to distinguish interruptions followed by a return to the same job from interruptions followed by a change of job, we find that there is no income penalty associated with time out without a job change, but a significant negative penalty associated with time out followed by a job change. (p. 428)

Furthermore, they found that a "job change" does not mean changing to a "part-time job" because women who had changed jobs but worked full time both before and after a job change still had lower current incomes.[5]

Second, since traditionally women play the roles of housekeeper and caregiver to elders and children, they may put much effort into household production, which means less energy can be spared for their paid work. This explanation of the family gap was discussed in Waldfogel (1995, 1998b) and England and Budig (1999). They thought of it as a potential explanation for a residual child penalty that remained after controlling for human capital and heterogeneity, but they did not measure it. Phipps, Burton, and Lethbridge (2001) examined the hypothesis that "traditional gender role assignments means that mothers often have much heavier domestic

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[4] But Corcoran (1977) found that work interruptions had relatively little effect on wages using data from the 1976 panel of income dynamics of the United States.

[5] They also distinguish women who have had interruptions but have only ever worked full-time from women who have ever returned to work part time.

responsibilities, which may limit the energy they can devote to their paid employment" and found that:

> Current weekly hours of unpaid work are negatively associated with women's current incomes and that controlling for unpaid work hours as well as career interruptions substantially lowers, though still does not generally eliminate, the family gap in Canada (p. 428).

The authors used data from the 1995 Statistics Canada General Social Survey, which provided information about current hours of unpaid work (housework, childcare, elder care). So they were able to do this analysis and examined "the alternative hypothesis that, because of more onerous unpaid work responsibilities, mothers are less productive in their paid work than non-mothers" (Phipps, Burton, and Lethbridge 2001, p. 414).

Third, the institutional structure of the labour market, including some family policies such as maternity leave and child care, may also have an effect on the family gap. Waldfogel (1997b) "provided some evidence to suggest that job-protected maternity leave could be an important remedy for the pay penalties associated with motherhood" (p. 122), and positive effects of job-protected maternity leave on the wages of mothers have been found in both the United States and Britain (Waldfogel 1998a). Zveglich and Rodgers (2003) found that the mandated maternity benefits did not have a significant reduction effect on women's wages in Taiwan. Because of the shortness of the relevant longitudinal data on job-protected maternity leave together with women's education, occupation, work history, and family status, thus far there have been no corresponding studies in Canada, while only Phipps (2000) has examined how fertility and labour-supply behaviour have been influenced by maternity and parental benefits.[6]

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[6] Phipps used the 1988/89/90 Labour Market Activity Survey and found "first, that fertility behaviour was not significantly influenced by the availability of benefits, and, second, that there is no evidence that women adjusted their labour-supply behaviour in order to gain access to benefits" (Phipps 2000, p. 415).

The fourth explanation of the family is gap that employers may discriminate against women with children. This hypothesis has been put forward in several papers to explain a residual child penalty (Waldfogel 1995, 1998a, 1998b; England and Budig 1999; Phipps, Burton, and Lethbridge 2001). There is no empirical evidence to support this explanation.

Finally, the existence of a wage penalty of motherhood is also attributed to the existence of unobservable heterogeneity, for example, mothers may prefer "mother-friendly" jobs to higher pay, and endogenous work interruptions or fertility. Fixed-effects modeling has been used in dealing with unobserved heterogeneity. Korenman and Neumark (1992) found that "short first-differences (estimated over a two-year period) indicate no negative effects of motherhood on wages" (p. 253-254). In contrast, all subsequent researchers found evidence of a significant wage penalty after controlling for unobserved differences. Even the same authors, in Neumark and Korenman (1994), found a negative association between children and women's wages after using sibling models to correct for unobserved heterogeneity by first- difference estimates. The advantages of using sibling data include: first, sibling data remove the heterogeneity bias due to unmeasured attributes such as parental investment in their daughters' human capital; and second, using sisters' fixed family background and attitudes as instruments allows then to account for both heterogeneity bias and endogeneity bias.

Waldfogel (1997a) compared a standard short first-difference model (with one to two years elapse) and a range of longer difference models (in two to nine years elapse) and found that "the longer the difference, the greater the estimated penalties associated with having children" and "using short first-differences might underestimate this penalty" (p. 213). She argued that short intervals may not be sufficient to capture the child penalty on wages if new mothers do not work at all and/or if the penalty is cumulative over time. This may explain the results of Korenman and

Neumark (1992). Comparing pooled OLS models and fixed-effects models, she concluded that unobserved heterogeneity does not bias the estimated effects of children on women's wages or bias the effects toward zero, which suggested unobserved heterogeneity is not significantly important in estimating the child-penalty (Waldfogel 1998a). Using the same techniques, England and Budig (1999) also draw the conclusion that a clear wage penalty for motherhood exists after controlling for unobservable heterogeneity.

Alternatively, it has been argued that work interruptions may be endogenous to income; that is, if absence from work is a choice, women with low income may be more likely to take longer interruptions (Stratton 1995; Phipps, Burton, and Lethbridge 2001). Regarding endogenous fertility, Lundberg and Rose (1999) argued that "since fertility is a choice and the market wage is one component of the cost of children, we would expect women with low (current or expected) market productivity to be more likely to become mothers" (p.6).[7]

Unfortunately, there do not currently exist for Canada any data sets that would allow me to examine all the above explanations of the family gap. However, I can re-examine what is generally considered to be the most important explanation -- the human capital theory explanation. I follow the lead of Phipps, Burton, and Lethbridge (2001), but use a more recent and larger sample of data -- the 2001 Statistics Canada General Social Survey (GSS 2001) -- to estimate the child effect on women's income, and then compare the results with theirs.

b. Econometric models of the family gap

Since Becker (1985) first used human capital theory to explain why married women earn less than married men, all researchers have used a human capital earnings equation to study the family gap between mothers and non-mothers. In recent years, Jane Waldfogel has written many

[7] I obtained this article from the Internet. No page numbers are shown.

articles on the family gap in the United States and the United Kingdom (Waldfogel 1995, 1997a,b, 1998a,b; Harkness and Waldfogel 1999; Joshi, Pace, and Waldfogel 1999). The econometric model she used is representative of those used in studying the family gap. Since Phipps, Burton, and Lethbridge (2001) are the first authors to study the family gap in Canada, the model they present is more practical for Canada. I use Waldfogel's model and Phipps, Burton, and Lethbridge's model to illustrate the econometric methods used to analyze the family gap.

In Waldfogel (1997a), first she used 1988 data from the National Longitudinal Survey of Young Women of the United States (NLS-YW), and estimated the following equation:

$$\begin{aligned} \ln W_{it} = {} & \beta_1 \cdot Exp_{it} + \beta_2 \cdot Exp_{it}^{\,2} + \beta_3 \cdot Age_{it} + \beta_4 \cdot Age_{it}^{\,2} + \beta_5 Educ_{it} \\ & + \beta_6 \cdot Mar_{it} + \beta_7 \cdot Sep_{it} + \beta_8 \cdot Div_{it} + \beta_9 \cdot Wid_{it} + \beta_{10} \cdot Onechild_{it} \\ & + \beta_{11} \cdot Children_{it} + \beta_{12} \cdot Black_{it} + \beta_{13} \cdot Hispanic_{it} + \mu_{it}, \end{aligned} \tag{1}$$

where *i* indexes individuals; *t* indexes time; *lnW* is the natural log of the hourly wage in 1988 dollars; *Exp* is years of actual work experience; *Age* is the individual's age in years; *Educ* is years of schooling; *Mar*, *Sep*, *Div*, and *Wid* are dummy variables for marital status equal to 1 if the individual is married (separated, divorced, or widowed) and 0 otherwise respectively (single is the omitted category); *Onechild* and *Children* are dummy variables for parental status equal to 1 if the individual has one child (has two or more children) and 0 otherwise respectively (no children is the omitted category); *Black* and *Hispanic* are dummy variables for race equal to 1 if the individual is black (Hispanic) and 0 otherwise respectively (other race is the omitted category); and μ_{it} is a disturbance term.[8]

Then she pooled the responses of individuals over the 1968-1988 period,[9] so that she could use a much larger sample and test for unobserved heterogeneity using a first difference

[8] In this equation, Waldfogel (1997a) did not include constant term and did not mention one in the article either.
[9] The NLS-YW interviewed the women in the sample 15 times between 1968 and 1988.

specification and a fixed-effects specification and found that unobserved heterogeneity did not bias the child penalty on women's wages. The first difference specification is:

$$\begin{aligned}\Delta \ln W_i = {} & \beta_1 \cdot \Delta Exp_i + \beta_2 \cdot \Delta Exp_i^{\,2} + \beta_3 \cdot \Delta Age_i + \beta_4 \cdot \Delta Age_i^{\,2} + \beta_5 \cdot \Delta Educ_i \\ & + \beta_6 \cdot \Delta Mar_i + \beta_7 \cdot \Delta Sep_i + \beta_8 \cdot \Delta Div_i + \beta_9 \cdot \Delta Wid_i + \beta_{10} \cdot \Delta Onechild_i \\ & + \beta_{11} \cdot \Delta Children_i + \Delta\alpha_i + \Delta\mu_i,\end{aligned} \quad (2)$$

where ΔlnW_i equals $(lnW_{i,t+1} - lnW_{i,t})$, ΔExp_i equals $(Exp_{i,t+1} - Exp_{i,t})$, and so on, and where α_i is an individual fixed effect and μ_i is a disturbance term. The fixed-effects specification is:

$$\begin{aligned}(\ln W_{it} - \overline{\ln W_i}) = {} & \beta_1 \cdot (Exp_{it} - \overline{Exp_i}) + \beta_2 \cdot (Exp_{it}^{\,2} - \overline{Exp_i^{\,2}}) + \beta_3 \cdot \\ & (Age_{it} - \overline{Age_i}) + \beta_4 \cdot (Age_{it}^{\,2} - \overline{Age_i^{\,2}}) + \beta_5 \cdot (Educ_{it} - \overline{Educ_i}) \\ & + \beta_6 \cdot (Mar_{it} - \overline{Mar_i}) + \beta_7 \cdot (Sep_{it} - \overline{Sep_i}) + \beta_8 \cdot (Div_{it} - \overline{Div_i}) \\ & + \beta_9 \cdot (Wid_{it} - \overline{Wid_i}) + \beta_{10} \cdot (Onechild_{it} - \overline{Onechild_i}) + \beta_{11} \cdot \\ & (Children_{it} - \overline{Children_i}) + (\alpha_{it} - \overline{\alpha_i}) + (\mu_{it} - \overline{\mu_i}),\end{aligned} \quad (3)$$

where lnW_{it} equals the natural log of the hourly wage for individual i at time t; $\overline{\ln W_i}$ equals the mean natural log of the hourly wage for individual i; and so on. She did not include region of residence or occupation variables in the model. However, because Harkness and Waldfogel (1999) study the family gap in pay in seven industrialized countries, they added a set of dummy variables for region and whether the woman lives in an urban area, changed education in years to dummies for level of education, and omitted the dummy variables *Black* and *Hispanic*.

Like Waldfogel, Phipps, Burton, and Lethbridge (2001) also estimate a typical human capital earnings equation, although they define the dependent variable differently and use slightly different explanatory variables. In their basic specification, the dependent variable is the natural log of personal income, and the explanatory variables include level of education, region of residence, and potential or actual experience. Regarding the reason why they chose the natural

log of personal annual income to be the dependent variable instead of the natural log of the hourly wage, they argued that "in addition to market earnings, motherhood could affect subsequent access to social welfare benefits associated with labour market attachment (e.g., EI versus social assistance, CPP versus GIS)," and because they could not obtain wages and salaries from the GSS 1995, their choice to study income was a pragmatic measure (p. 415). The same data deficiencies exist in the GSS 2001, so I choose the natural log of personal income to be the dependent variable as well. They did not include age or occupational dummies in their model. Although England and Budig (1999) included various job characteristics, such as "measures of on-the-job training, physical strength, and physical hazards demanded in one's job", they found that occupational characteristics did not help explain the family gap (p.5).[10] But Miller (1987) found occupation and industry made the magnitude of the child effect on wages smaller.

c. The magnitude of the family gap

It is well-known that women with children earn less than women without children in many countries. Most researchers measure the family gap in mean hourly wages or some transformation thereof between women with children and women without children, shown in Table 1.

Waldfogel (1995) estimated the family gap in the United Kingdom in the mean wage female/male pay ratio and found a larger family gap for women aged 33 in 1991 (20 per cent) and than for women aged 23 in 1981 (14 per cent). Harkness and Waldfogel (1999) measured the family gap in the mean wage female/male pay ratio from seven industrialised countries for women aged 24-44 who worked full-time. They found that Germany had the largest family gap (9.0 per cent) and Australia had the smallest (1.4 percent). However, Phipps, Burton, and

[10] I obtained this article from the Internet. No page numbers are shown.

Lethbridge (2001) found a family gap in mean income in 1995 such that mothers aged 25-54 earned 87.3 per cent of non-mothers' income.

Even after controlling for some characteristics, such as actual work experience, education, age, and marital status, researchers still find a child penalty for mothers. Most researchers used the *log of hourly wages* as the dependent variable and measured the remaining family gap as the coefficient of a variable related to the number of children (with two dummy variables: *Having one child* and *Having two or more children,* or sometimes they separate *Having two or more children* into *Having two children* and *Having three or more children*), while Phipps, Burton, and Lethbridge (2001) used the dependent variable *log of personal annual income* and included the dummy variable *Ever had children.* See Table 2 for details.

In the United States, Korenman and Neumark (1992) used panel data from the 1982 wave of the National Longitudinal Survey of Young women (aged 28-38) and found a one child penalty of 4 per cent and two or more children penalty of 7 per cent for white women with children as compared to white women without children. When they used short first difference modeling (1980-1982) to correct for the unobserved heterogeneity, the child penalty on women's wages disappeared. In order to study sources of bias in women's wage equations, using sibling data from the same survey as they used in 1992, Neumark and Korenman (1994) estimated the effect of the number of children on women's wages and found that for white women the magnitude was 7.4 per cent in an OLS equation and 3.9 per cent in a fixed-effects equation; 0.8 per cent and 3.3 per cent respectively for black women. Using panel data from the 1968-1988 National Longitudinal Survey of Young Women (NLS-YW) Waldfogel (1997a) indicated that for women aged 34-44 in 1988 the magnitude of the one child penalty was 4.1 per cent and the penalty for having two or more children was 9.6 per cent. After pooling fifteen years of data over

the 1968-1988 period and estimating fixed-effects specifications, the magnitude of the child penalty did not change significantly.

Waldfogel (1998b) also used panel data from the National Longitudinal Survey of Youth (NLSY) and found that for women with an average age of 30 the one child penalty doubled from 1980 (4.5 per cent) to 1991 (9.8 per cent), but the penalty for having two or more children increased only slightly (from 9.2 per cent in 1980 to10.9 per cent in 1991). In Canada, Phipps, Burton, and Lethbridge (2001) used microdata from the GSS 1995 and found that for women aged 25-54 who worked full time the magnitude of the coefficient of the dummy variable *Ever had children* was 15.7 per cent.

Harkness and Waldfogel (1999) studied the family gap in seven industrialised countries using microdata from the Luxembourg Income Study. They included three dummy variables related to the number of children: *Having one child, Having two children,* and *Having three or more children*. For women aged 24-44 who worked full-time, the United Kingdom had the largest two or more children penalty (the magnitudes of the two children penalty and the three or more children penalty were 17.6 per cent and 29.5 per cent respectively) and Finland had the smallest (the magnitudes of the two children penalty and the three or more children penalty were 3.4 per cent and 8.2 per cent respectively). Australia had the largest magnitude of the one child penalty (12.2 per cent). There was no one child penalty in the United Kingdom, Germany, or Sweden.

III. Data and Model

The data used in this paper are from the fifteenth cycle of the General Social Survey (GSS 2001), conducted from February through December 2001.[11] It is the most recent data set that collects information about people's work histories in Canada. The target population for the survey was all persons 15 years of age and older in Canada, excluding residents of the Yukon, Northwest Territories, and Nunavut and full-time residents of institutions. The original size of the sample for GSS 2001 was 24,310, two and half times larger than that of the GSS 1995 (10,749 observations). For most of this study I use a subsample of women aged 25 to 54 who worked full-time full-year during the reference year and who reported all relevant information, such as education, income, duration of work interruptions, reasons for work interruptions, occupation, and family status (2,740 observations).[12] This sample is more than twice as large as the sample used by Phipps, Burton, and Lethbridge (2001) (1,296 observations).

In a slight departure from the GSS 1995, in the GSS 2001 respondents were asked the reason for any work absence longer than three months, not of "at least six months" duration. Seventeen different reasons are available, which makes it possible to separate time out for maternity or childcare reasons from time out for other reasons.[13]

In contrast to the GSS 1995, the GSS 2001 dropped the questions about whether the respondent returned to the same or a different job after the interruption and hours of unpaid work, so I cannot re-examine the analyses related to these. In addition, the GSS 1995 and 2001 do not

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[11] The GSS 2001 was conducted during 2001 and the income requested was for the last 12 months before the respondent was asked. I use 2001 as the reference year.

[12] People aged 25 to 54 are considered "the standard 'prime-aged worker' group" (Phipps, Burton, and Lethbridge 2001, p. 415).

[13] The reasons include: Seasonal worker; Temporary layoff/end of contract; Lack of work; Business or company closure; Moved or transferred; Changed jobs; Own illness/disability/accident; Immigration/emigration; Returned to school; Retired; Maternity/paternity leave; Child care; Elder care; Marriage; Personal or family responsibilities; Other reason; Personal choice.

include information about hourly wages, which most previous researchers have used, so I use personal income and focus the analysis on people who are currently working full-time full-year as did Phipps, Burton, and Lethbridge (2001). Since individuals who work full-time full-year are more likely to be working approximately the same number of hours, this restriction of the sample is likely to reduce the problems associated with the joint determination of hours and wages.[14]

Basically I use Phipps, Burton, and Lethbridge's model because I am studying the same country as they did, and use the same survey (although a different cycle). My analysis is based on the following human capital earnings equation:

$$\ln Income_i = \beta_1 + \beta_2 \cdot Exp_i + \beta_3 \cdot Exp_i^2 + \beta_4 \cdot Child_i + \beta_5 \cdot Marry_i + x_i' \cdot \alpha + \mu_i, \quad (4)$$

where $i = 1,\cdots,N$ indexes individuals; *lnIncome* is the natural log of personal income; *Exp* is potential or actual experience in years; *Child* is a dummy variable equal to 1 if the individual ever had children, and 0 otherwise; *Marry* is a dummy variable equal to 1 if the individual ever married, and 0 otherwise; β_1 is a constant term; μ is an error term; and x_i is a vector that contains all the other control variables, including *Edu* dummies, *Prov* dummies, and *Ocup* dummies. *Edu* is a set of dummy variables for the individual's highest level of education; *Prov* is a set of dummy variables for the individual's province of residence; and *Ocup* is a set of dummy variables for the individual's occupation.

The GSS 2001 data do not provide personal income as a continuous value, but in twelve categories, while there are twenty categories in GSS 1995. I use the mid-point of each category as the income value as did Phipps, Burton, and Lethbridge (2001). For individuals in the "more than $100,000" category, I use the mean income for persons in the Survey of Household

[14] The Survey of Consumer Finances definition of a full-year full-time worker is a person who works 30 hours or more per week for 49-52 weeks in the reference year (Drolet 1999).

Spending 2001 (SHS) with reported incomes above $100,000, which is $154,137.[15] Although Stewart (1983) showed that in general, applying OLS after setting the dependent variable equal to the midpoint of each income range will result in inconsistent estimates, Phipps, Burton, and Lethbridge (2001) noted that the qualitative results were the same when interval regressions were estimated. I leave this discussion to future work.

Because most datasets do not include a measure of actual experience, economists usually compute potential experience as age less years of education less six. But the GSS 2001 provides the respondent's age at the beginning of the first work period, so I can use the number of years since the first work period to measure potential experience as do Phipps, Burton, and Lethbridge (2001). Furthermore, they argue that the dummy variable *Child* should be set equal to 1 if the individual "ever had" rather than "currently has" a child because a dummy variable for "currently having a child" would exclude older women from the class of "mothers."

To the basic equation of Phipps, Burton, and Lethbridge (2001), I add dummy variables for occupation to the set of independent variables, since the large sample provided by GSS 2001 permits one to do so. I do not include age in my model because the correlation between age and potential work experience (0.88) and the correlation between age and actual work experience (0.83) are too high.[16] In addition, I use dummy variables for province of residence instead of region of residence because I want to be able to distinguish Alberta from the other Prairie provinces. Wages in Alberta are consistently above the Canadian average, while in the other Prairie provinces they tend to lie below the Canadian average. Thus there are likely to be differences between the wage structure in Alberta and that in the other two provinces. As will be

[15] Phipps, Burton, and Lethbridge (2001) used the mean income for individuals in the 1995 Survey of Consumer Finance (SCF) with reported incomes above $100,000. Because there is no SCF for 2001, I use the SHS 2001 instead.

[16] I also estimated all the specifications for women with *Age* included. The magnitude of the *child* dummy increased a little bit. Although the magnitudes of all the coefficients were a little different, the conclusions did not change.

seen later, there is indeed a statistically significant difference between incomes in Alberta and the other Prairie provinces.[17]

There are three groups of dummy variables included in my model. The first group reflects the respondents' highest level of education. It is in six categories: doctorate or Master's or Bachelor's degree; diploma or certificate from community college or trade or technician; some university or community college; high school diploma; some secondary or elementary or no schooling (reference level); and other beyond elementary or high school. The second group identifies the respondents' geographic area of residence. They are classified as ten provinces: Newfoundland and Labrador; Prince Edward Island; Nova Scotia; New Brunswick (reference province); Quebec; Ontario; Manitoba; Saskatchewan; Alberta; and British Columbia. The third group indicates the respondents' occupation. Following the Standard Occupational Classification (1991), the occupations of respondents were divided into eight categories: management occupations; professional occupations; technologists, technicians and technical occupations; clerical occupations; sales and services occupations; trades, transport and equipment operators and related; occupations unique to primary industries (reference occupation); and occupations unique to processing, manufacturing and utilities. Means and standard errors for all variables for women aged 25 to 54 who were full-time full-year workers are shown in Table 3, for mothers in Table 4, and for non-mothers in Table 5.

From Table 3, it can be seen that 33.6 per cent of working women have a diploma or certificate from a community college or trade or technician, and 30.0 per cent of them have a doctorate or Master's or Bachelor's degree. Most women worked in the provinces of Ontario

[17] The region of residence variable in the GSS 2001 combines the provinces of Manitoba, Saskatchewan, and Alberta together into one region. The results show that Manitoba and Saskatchewan have no significant effect on women's current income and the coefficient of Alberta is positive and significant compared to New Brunswick (see Tables 4, 5, and 8).

(29.7 per cent) and Quebec (20.0 per cent) and most were engaged in clerical occupations (27.0 per cent), professional occupations (25.7 per cent), and sales and services occupations (23.5 per cent). 75.3 per cent of women had ever married and 63.6 per cent of women had a child. Comparing mothers and non-mothers (Table 4 and Table 5), one can see that non-mothers are more likely to achieve a higher level of education (42.4 per cent of non-mothers have a doctorate or Master's or Bachelor's degree as compared to 22.9 per cent of mothers); non-mothers are more likely to be engaged in professional occupations (32.2 per cent) than mothers (21.9 per cent) while mothers are more likely to be engaged in clerical occupations (29.0 per cent) than non-mothers (23.6 per cent); and mothers take a longer length of time out of the labour market with a mean total duration of 2.30 years, mostly related to child care (1.39 years), than non-mothers with a mean total duration of work interruptions of 0.43 years.[18]

IV. Results and Discussion

I build on the study of Phipps, Burton, and Lethbridge (2001) and use more recent data from the GSS 2001 to see if any of their conclusions change. Basically I estimate the same equations as theirs, aside from adding occupational dummies, and structure my discussion along the same lines as theirs. The starting point for estimation is equation (4), in which the dependent variable is the log of personal annual income. Additional variables were added to this equation to explore different possible explanations of the family gap. In total five different specifications of the basic model were estimated for women aged 25-54 who worked full-time, full year, while two specifications were also estimated for men.

[18] Here, women aged 25 to 54 who worked full-time full-year without any work interruptions are included.

Phipps, Burton, and Lethbridge (2001) used OLS to estimate all the equations discussed in their paper. I too use OLS estimation, but also carried out some diagnostic tests of the underlying assumptions of the linear regression model for the equations reported in Table 10 (specification 5). These tests indicated, first of all, that the errors of the regression model were not normal; and second, that the errors are heteroskedastic.[19] I think that the main reason for the nonnormality is likely that the dependent variable -- the natural log of personal income -- in the original data source is not continuous. Instead it is set equal to the midpoint of each range and is truncated below at zero. However, as was noted earlier it is beyond the scope of this paper to deal with this problem.

To deal with the heteroskedasticity problem, I examined standard errors computed using White's heteroskedasticity-consistent covariance matrix estimator. While this affected the p-value of the estimated coefficients, in most cases the conclusions did not change. For consistency with Phipps, Burton, and Lethbridge (2001), the OLS standard errors are reported in the tables, and any changes to the conclusions resulting from using White's heteroskedasticity-consistent covariance matrix estimator are discussed in the Appendix.

a. Is there any family gap in Canada?

The first question to be addressed is whether or not there was still a family gap in Canada in 2001. For women aged 25 to 54 who were full-time full-year workers, the family gap between mothers and non-mothers is 0.10 in the mean natural log of income and 4.8 per cent in the mean female/male income ratio (see column 1 in Table 4 and column 1 in Table 5),[20] which is smaller than in 1995. It is also smaller than that of the United States and Germany in 1994 and larger

[19] See the Appendix for the results of the diagnostic tests.
[20] In 2001, for men aged 25 to 54 who were full-time full-year workers, the mean income is $57,679.

than that of the United Kingdom in 1995, Australia in 1994, and Finland and Sweden in 1991 (Harkness and Waldfogel 1999).

To see whether there still exists a family gap after controlling for work experience, education, residence, and occupation, first I estimate two specifications of equation (4) for men and women aged 25 to 54 who were full-time full-year workers, both of which include potential work experience. As Phipps, Burton, and Lethbridge (2001) show the effects of children and marriage on the incomes of women and men were different in 1995. To see if this is still the case in 2001, I estimate the two specifications for men as well as women. The first specification does not include a dummy variable that equals one if the individual has ever been married, while the second one does.[21] The results shown in Table 6 indicate that the coefficient of the "ever had a child" dummy is negative and significant in both specifications for women, but positive and significant in both specifications for men. In specification 2, the coefficient of the "ever married" dummy implies a significant earnings premium associated with marriage for both men and women. When the two specifications are compared, one can see that when marital status is added, the positive coefficient of "ever had a child" for men drops from 0.130 to 0.067, while the negative coefficient of "ever had a child" for women increases in magnitude from -0.049 to -0.073. Thus the "marriage premium" constitutes part of the "child premium" for men observed in specification 1. Similarly, it reduces the "child penalty" for women estimated in specification 1. For women, the "marriage premium" is smaller than the "child penalty" (the coefficient of "ever had a child" is -0.073 and that of "ever married" is 0.051) and cannot completely offset it. Therefore, there is a "child penalty" for women, a "child premium" for men, and a "marriage premium" for both, though the latter is larger for men (0.121) than for women (0.051). The

[21] Correlations between the "ever had a child" and "ever married" dummy variables are low in specification 2 (0.53 for women and 0.58 for men).

possible explanation may be that men who are married or have children are considered to be more stable and to put more energy into their job (e.g., they are less likely to change their job and their wives take on a greater share of the housework). However, for women with children the results tell a different story – it would seem that from the employer's perspective they are considered to be less stable and less productive (e.g., they may be more likely to change their job and spend more energy on caring for children).

Hence, in Canada the family gap between mothers and non-mothers does exist even after one controls for marital status, which is consistent with the results from the GSS 1995 (see Phipps, Burton, and Lethbridge 2001). The magnitudes I obtain for the child penalty for women are smaller than theirs. Before marital status is added, I estimate the magnitude to be -0.049, while their estimate is -0.152. After controlling for marital status, my estimate of the child penalty increases in magnitude to -0.073, while theirs increases to -0.172. When I performed a t-test of the null hypothesis that the coefficient of the child dummy is equal to that in Phipps, Burton, and Lethbridge (2001) for specifications 1 and 2 in Table 6, the results show that my estimates of the family gap are statistically different from theirs.[22] Thus the magnitude of the child penalty appears to be smaller in 2001 than in 1995.

Another difference between my results and those of Phipps, Burton, and Lethbridge (2001) is that they reported that when marital history is added, the "ever had a child" dummy becomes insignificant for men, and the "ever married" dummy becomes insignificant for women.[23] Thus, they concluded that "the child penalty is not actually a marriage penalty for women, though the child premium may be a marriage premium for men" (p. 417). They used the GSS 1995 data set

[22] For specification 1, the t statistic is 5.42 and for specification 2, the t statistic is 4.55. In both cases, they are statistically different at the 1% significant level.

[23] I also tried the same OLS regression specifications as in Phipps, Burton, and Lethbridge's (2001) Table 1, that is, without occupational dummies and with region dummies instead of province dummies, but using the GSS 2001. The results are not very different from those in Table 4.

with a relatively small sample (1802 observations for men and 1296 observations for women), while I use the GSS 2001 data set with 3649 observations for men and 3015 observations for women, which may be responsible for the difference in results.

In addition, due to the six-year time interval between the two surveys, those people aged 49-54 in 1995 are excluded from my sample and those aged 19-24 in 1995 are included. Since the GSS does not follow the same people, it is impossible to restrict the sample to include the same individuals in both years. To test for cohort effects, I re-estimated specifications 1 and 2 for men and women aged 31-54 in the GSS 2001 (3022 observations for men and 2481 observations for women). However, the results are the same as those in Table 6, which suggests that there are no cohort effects.

In Table 6, all the dummy variables for education for both men and women show positive and significant results relative to the reference education level of some secondary or elementary or no schooling, which suggests that a higher education level is associated with a higher income. All the coefficients of the education dummies for women are larger than those for men, so education is more important to women's income than to men's. As far as the residence dummies are concerned, only the provinces of Ontario, Alberta, and British Columbia are significantly different from New Brunswick for both men and women. There is a greater residence premium for women than men in Ontario, but a lower one in Newfoundland and Labrador, Manitoba, Saskatchewan, Alberta, and British Columbia.

Turning to the occupational dummies, I chose "Occupations unique to primary industries" to be the reference category. The coefficients of all the occupational dummies are statistically significant except that for "Technologists, technicians and technical occupations". Table 6 indicates that the coefficients for some occupations, including "Clerical occupations," "Sales and

services occupations," "Trades, transport and equipment operators and related occupations," and "Occupations unique to processing, manufacturing and utilities," are negative and significant for men, but positive and significant for women, suggesting that men earn more in primary industry occupations than in those four fields, while women earn more in those four fields than in primary industry occupations. "Management occupations" are most lucrative for men and "professional occupations" are most lucrative for women. For men the coefficients of all the occupational dummies range from -0.170 to 0.145, and for women they range from 0.178 to 0.645, which shows that the choice of occupation has a greater effect on women's income than on men's.

The results for specifications 1 and 2 when the occupational dummies are excluded are shown in Table 7. When I do not control for occupation, for women, the child penalty becomes larger. When the "ever married" dummy is included, it changes from -0.073 (Table 6) to -0.077 (Table 7); without the "ever married" dummy it changes from -0.049 (Table 6) to -0.054 (Table 7). This means that controlling for the occupations chosen by women reduces the family gap, though it is still statistically significant. This result is consistent with those of Miller (1987). He used data from the 1981 census of Canada and controlled for potential work experience, education, marital status, province of residence, ethnicity, and language for women aged 25-64. When he added occupation and industry dummy variables to the specification, the coefficients of the child effect on women's wages became smaller in magnitude, although his results showed that the coefficient of the number of children living at home less than six years old was positive and the coefficients of number of children living at home between six and fourteen and between fifteen and twenty five years old were negative. Because the coefficients of all but one of the occupational dummies are individually significant in the OLS regressions and all the joint F statistics for the test of the null hypothesis that all the coefficients of the occupational dummies

are jointly zero show that they are statistically different from zero at the significant level of 1%,[24] I control for occupations in subsequent estimation.

In other words, for women, when I control for potential experience, age, education, province of residence, and occupation, with or without the variable of "ever married," all the results show that the coefficient of the "ever had a child" variable is negative and statistically significant. Therefore, in Canada a family gap does exist between women with children and women without children.

b. Does time out help to explain the family gap?

Before addressing this question, in Table 8 I compare the percentage who take an interruption of at least six months, the mean total duration of the interruptions, and the reasons for the interruptions between men and women who are employed and men and women who are full-year full-time workers in the reference year of 2001. In order to see if anything changed between 2001 and 1995, I adopt the same design as Phipps, Burton, and Lethbridge (2001) did in their Table 2 (p. 418).[25] My Table 8 shows that the distribution of responses is almost the same for the full-time only group and the full-time and part-time group. 12.7 per cent of currently employed men have had at least one longer interruption, which is less than the percentage of women (37.4 per cent of women). The mean total duration of interruptions for men is 1.70 years, which is half of that for women (3.53 years for women). Interestingly, Phipps, Burton, and Lethbridge (2001) using the GSS 1995 found that the mean total duration of interruptions for women was much longer (6.15 years). Otherwise, the results are the same as those of Phipps,

[24] In Table 6, for men in specification 1 the F statistic is 21.19, while in specification 2 the F statistic is 20.32; for women in specification 1 the F statistic is 55.84, while in specification 2 the F statistic is 55.99. All the p-values are less than 0.000005.

[25] Any interruptions of less than six months are excluded here. In addition, interruptions for the reasons education, changed jobs, immigration, personal choice, retired or other are excluded.

Burton, and Lethbridge (2001), who found that "not only are women more likely to experience interruptions, but their interruptions tend to be of significantly longer duration" (p. 417). I also computed the same statistics for all work interruptions of more than 3 months (the values in parentheses shown in Table 8 and Table 9); again, the pattern is much the same.

In addition, I notice that only 10.5 per cent of the 12.7 per cent of men with any interruptions reported family reasons for the interruption, but 84.6 per cent of the 37.4 per cent of women with any interruptions did. Most men had an interruption because of demand-related reasons or an inability to work (63.7 per cent of men with any interruptions gave "seasonal worker", "temporary layoff/end of contract", "lack of work", or "business or company closure" as a reason; 30.0 per cent gave "own illness or disability or accident" as a reason). Not surprisingly, women took interruptions mainly for family reasons, not demand-related reasons (14.5 per cent of women with any interruptions) or health problems (9.6 per cent). Furthermore, 90.7 per cent (79.4 per cent for maternity leave and 11.3 per cent for child care) of women's family-related work interruptions are associated with children rather than other family reasons (18.0 per cent).[26] Clearly women do spend more time out of the labour market on child care than men.

I also compare the same items for mothers who worked full-time and for non-mothers who worked full-time (see Table 9). The results show that women without children have similar responses to men, although more non-mothers have interruptions for "family" reasons than men (23.5 per cent for non-mothers, Table 9, and 10.8 per cent for men, Table 8) and fewer for "demand" reasons (53.8 per cent for non-mothers, Table 9, and 62.9 per cent for men, Table 8). This confirms that women do spend more time on family work than men. For mothers who had

[26] The percentages may add to more than 100 because a respondent can have more than one interruption and/or more than one reason for any particular interruption.

an interruption, "family" reasons were the most important (91.7 per cent), and 93.8 per cent of them are related to child care. Furthermore, the mean total duration of interruptions was only 1.59 years for non-mothers (similar to men with 1.64 years), but mothers had a mean total duration of 3.55 years. It is obvious that women with children do spend much more time out of the labour force than women without children.

To test for the effect of time out of labour market on the family gap between mothers and non-mothers, first I control for actual work experience, defined as potential experience less the total duration of any interruptions of more than six months (see specification 3 in Table 10).[27] When I replace potential work experience with actual work experience, the "child penalty" falls from -0.073 (Table 6, column 4) to -0.055 (Table 10, column 1). That is, 18 per cent of the family gap disappears after controlling for actual work experience and 55 per cent remains (the family gap in the natural log of income is 0.10). This result supports the human capital explanation of the family gap; that is, the more time out of the labour market women take, the less human capital they accumulate. It is also consistent with the findings of Phipps, Burton, and Lethbridge (2001) and other studies (Korenman and Neumark 1992; Waldfogel 1995, 1997a,b, 1998a; England and Budig 1999). Comparing the magnitude of the two children penalty estimated in Harkness and Waldfogel (1999), the child penalty in Canada in 2001 is smaller than in the United Kingdom in 1995, Australia and Germany in 1994, and Sweden in1991 and larger than the United States in 1994 and Finland in 1991. At the same time, the positive coefficient of experience increases from 0.024 (Table 6, column 4) to 0.029 (Table 10, column 1), while the absolute value of the coefficient of experience squared increases much less in magnitude -- from -0.0004 (Table 6, column 4) to -0.0005 (Table 10, column 1) -- so the returns to experience

[27] I also estimate the OLS regression with any interruptions of more than three months, but there are no big differences in the results.

become larger. This is possible since potential work experience over-estimates actual work experience for many women, but surprisingly Phipps, Burton, and Lethbridge (2001) did not find any noticeable change in these coefficients.

Next, I add the total duration of work interruptions, measured in years, to specification 3 (see specification 4 in Table 10). The negative coefficient of "ever had a child" becomes much smaller in magnitude, changing from -0.055 (Table 10, column 1) to -0.045 (Table 10, column 2). The duration of work interruptions has a significantly negative effect on income (the coefficient is -0.006), which supports the view that women's existing human capital depreciates during their time out of the labour market (Corcoran 1977; Mincer and Ofek 1982; Phipps, Burton, and Lethbridge 2001). I find that the magnitude of human capital depreciation cannot be neglected. The negative effect of one year out of the labour force is equal to 20.3 per cent of the positive effect of one year of work, as compared to 37 per cent in Phipps, Burton, and Lethbridge (2001).

In other words, taking into account the total duration of time out of labour market, there are two effects on women's current income: one is human capital depreciation during the time out, the other is reduced human capital accumulation during the time out. These effects reduce the child penalty from -0.073 (Table 6, column 4) to -0.045 (Table 10, column 2). So controlling for time out of the labour market really reduces the family gap between mothers and non-mothers, although the child penalty is still statistically significant.

c. Do child-related work interruptions have an effect on the family gap?

To see if child-related work interruptions have any consequences for women's current income, I separate the total duration into four parts: the first is the duration of maternity and child care interruptions; the second is the duration of demand-related interruptions; the third is the

duration of health-related interruptions; and the fourth is the duration of other interruptions.[28,29] The results for specification 5 reported in Table 10 indicate that only demand-related and other interruptions have a negative and significant effect on women's current income, while child-related and health-related interruptions are insignificant. The negative effect of the duration of demand-related interruptions is that the loss from one year out of work is equal to 67.8 per cent of the gain from one year of work, and the negative effect of the duration of other interruptions is that the loss from one year out of work is equal to 33.9 per cent of the gain from one year of work. This finding suggests that human capital depreciation does occur during long absences from work due to demand-related and other reasons, but child-related and health-related interruptions do not have a negative effect on women's current income. A possible explanation for this result may be on-the-job training. For instance, when women leave work due to demand-related and other reasons, they will never benefit from on-the-job training during previous job, but when women are absent from work due to child-related or health-related reasons, they may still benefit from on-the-job training if they do not change jobs after the work interruptions. When Phipps, Burton, and Lethbridge (2001) distinguished the duration of child-related work interruptions followed by a return to the same job from those followed by a job change, they found that "there is no extra penalty associated with time out followed by a return to the same job" (p. 423), which supports my explanation.

My results with respect to different types of job interruptions are slightly different from those of Phipps, Burton, and Lethbridge (2001).They found that the duration of child-related

[28] The demand-related interruptions include seasonal worker, temporary layoff/end of contract, lack of work, business or company closure. The other interruptions include moved or transferred, elder care, marriage, personal or family responsibilities and those who gave education, immigration, personal choice, retired or other as a reason for any interruption.

[29] Some respondents to the GSS 2001 gave more than one reason for an interruption. In such cases, I divide the duration equally between causes. But this does not greatly affect the results because only 1.4 per cent of women gave more than one reason for an interruption.

interruptions had a negative and significant effect on women's current income: the loss from one year out of work was equal to 43 per cent of the gain from one year of work. Demand-related interruptions had a larger negative effect on women's current income, and the loss from one year out was more than the gain from one year in. In their equation, the coefficients of the duration of health-related and other interruptions were insignificant. They found that time out of the labour market does not have negative effects on current incomes purely as a result of human capital depreciation because different types of work interruptions have different effects on women's current incomes.

But interestingly, when I carry out a Wald test of the joint null hypothesis that the coefficients of the four types of durations are the same, using White's heteroskedasticity-consistent covariance matrix estimator, the value of the test statistic is 4.761 with a p-value of 0.190, which indicates that the null hypothesis cannot be rejected. That is, the effects of the durations of different types of work interruption on women's current income are not significantly different. This finding suggests that human capital depreciation does occur during long absences from work no matter what the cause of the absences.

The coefficients of the other explanatory variables do not change much when I decompose the total duration of work interruptions into child-related, demand-related, health-related, and other interruptions. With respect to the effect of the duration of child-related interruptions on the family gap, the coefficient of the child dummy does not change much. Thus I draw the same conclusion as Phipps, Burton, and Lethbridge (2001): that "simply constructing a more accurate measure of work history does little to account for the family gap" (p. 427).

V. Conclusion

Using a more recent and larger sample of data from the 2001 Statistics Canada General Social Survey, I find new evidence on the family gap between mothers and non-mothers in Canada. There is a child penalty associated with being a mother, though a child premium is associated with being a father. I follow Phipps, Burton, and Lethbridge (2001) but add occupational dummies to their OLS regression model. The results indicate that occupational choice has a bigger impact on women's incomes than on men's, and that women receive the highest returns in "professional occupations," while men get the highest pay in "management occupations." In addition, controlling for occupation has an effect on the child penalty. In all the specifications I estimate, the coefficient of the "ever had a child" dummy variable is negative and significant, implying that having had children has a negative effect on women's current income. This suggests the family gap between women with children and other women persists in Canada.

I also examine the human capital explanation of the family gap. Descriptive statistics show that on average, women with children take more time out and their work interruptions are of longer duration than those of women without children and men. When I control for actual work experience by subtracting from potential work experience the total duration of work interruptions, the returns to work experience increase from 0.024 (Table 6, column 4) to 0.029 (Table 10, column 1) and the child penalty falls from -0.073 (Table 6, column 4) to -0.055 (Table 10, column 1). This means that the more time out of the labour market women take, the less human capital they acquire. The child penalty remains statistically significant, so controlling for actual work experience does not eliminate the family gap. This result is consistent with the finding of Phipps, Burton, and Lethbridge (2001) and other studies (e.g., Korenman and Neumark 1992; Waldfogel 1995, 1997a,b, 1998a; England and Budig 1999). Then, when I take account of the

effect of the total duration of work interruptions, the results show that women also suffer from human capital depreciation during work interruptions (the human capital loss from one year out of work is equal to 20.3 per cent of the human capital gain from one year of work). In conclusion, women who take a long break from the labour market not only gain less human capital, but also suffer a depreciation of their existing human capital.

When I distinguish maternity leave and child care interruptions from demand-related interruptions, health-related interruptions and other interruptions, I find that the four kinds of interruptions have different coefficients but the differences among them are not statistically significant. This suggests that human capital depreciation occurs during all work interruptions regardless of the reasons. Although my results differ in this respect from those of Phipps, Burton, and Lethbridge (2001), we draw the same conclusion on the effect of child-related interruptions on the family gap: controlling more accurately for work interruptions does not help explain the family gap between mothers and non-mothers.

Because of data limitations, I can only examine the human capital explanation for the earning penalty of motherhood in Canada. Other explanations remain to be examined in future research. For example, regarding the effect of the institutional structure of the labour market, employment insurance policies related to parental leave benefits changed a great deal on December 31, 2000. Beginning in 1971, mothers who had worked for more than 20 insurable weeks could claim up to 15 weeks of maternity benefits. As of 1990, women were entitled to 15 weeks of maternity leave benefits and 10 weeks of parental leave benefits. The 10 weeks of parental leave could be split between parents. As of December 31, 2000, the threshold for women's qualification was reduced from 700 hours of work in the past 52 weeks to 600 hours, and parental leave benefits increased from 10 weeks to 35 weeks. But the rate of benefit did not

change: it remained equal to 55 per cent of prior weekly insurable earnings with a maximum of $413 per week (Marshall 2003).[30] As of March 3, 2002, the 50 weeks of benefits for parents is not reduced by weeks of sickness benefits which mothers take during pregnancy (Pérusse 2003). However, the GSS 2001 was conducted during 2001 and I include in my sample only people who were full-time full-year workers in 2000, so I am unable to examine the effect of the new policy on the family gap. This question should be addressed in future research.

[30] "The Employment Insurance Coverage Survey is a supplement to the Labour Force Survey since 1997", and from 2000 "a special maternity supplement was added to help minitor the effect of the extended parental benefit program" (Marshall 2003, p 14). But I cannot access to it through the library of the University of Ottawa.

References

Baxter, Janeen, "Domestic labour and income inequality," *Work, Employment and Society*, 6 (1992): 229-249

Becker, Gary, "Human capital, effort, and the sexual division of labour," *Journal of Labour Economics*, 3 (1985):s33-s58

Budig, Michelle J. and Paula England "The wage penalty for motherhood," *American Sociological Review*, 66 (2) (2001):204-255

Corcoran, Mary, "Work experience, labour force withdrawals, and women's wages: empirical results using the 1976 panel of income dynamics," *Women in the Labour market*, ed. Cynthia Lloyd, Emily Andrews, Curtis Gilroy (New York: Columbia University Press 1977)

Drolet, Marie, "The persistent gap: New Evidence on the Canadian Gender Wage Gap," *Statistics Canada*, Income Statistics Division, Catalogue no: 75F0002MIE, (December 1999)

England, Paula and Michelle J.Budig, "The effects of motherhood on wages in recent cohorts: findings from the National Longitudinal Survey of Youth," paper presented at the annual meetings of the American Economic Association, 4 (January 1999), New York, Internet address: http://www.olin.wustl.edu/macarthur/working%20papers/wp-england3.html

Harkness, Susan and Jane Waldfogel, "The family gap in pay: evidence from seven industrialised countries," Luxemnourg Income Study, working paper No.219 (1999)

Hill, Martha S., "The wage effects of marital status and children," *Journal of Human Resources,* 14(4) (1979): 579-594.

Jacobsen, Joyce and Laurence Levin, "Effects of intermittent labour force attachment on women's earnings," *Monthly Labour Review*, 118 (9) (1995): 14-19

Joshi, Heather and Marie-Louise Newell, "Pay differentials and parenthood: analysis of men and women born in 1946," Coventry, England: University of Warwick Institute for Employment Research (1989).

Joshi, Heather, Pierella Pace, and Jane Waldfogel, "The Wages of motherhood: better or worse?" *Cambridge Journal of Economics*, 23(5) (1999):543-564

Korenman, Sanders and David Neumark "Marriage, motherhood, and wages," *The Journal of Human Resources*, Vol. 27, No. 2, (1992): 233-255.

Lundberg, Shelly and Elaina Rose, "Parenthood and the earnings of married men and women," Presented at the 1999 annual meetings of the American Economic Association, New York (1999), Internet address: http://csde.washington.edu/downloads/98-9.pdf

Marshall, Katherine, "Parental leave: more time off for baby," *Canadian Social Trends*, (Winter 2003):13-18

Miller, Paul W., "Gender differences in observed and offered wages in Canada, 1980," *The Canadian Journal of Economics*, 20(2) (1987):225-244

Mincer, Jacob, and Haim Ofek, "Interrupted work careers: depreciation and restoration of human capital," *Journal of Human Resources,* 17 (1982):3-23

Neumark, David and Sanders Korenman, "Sources of bias in women's wage equations: results using sibling data," *Journal of Human Resources*, 29 (1994):379-405

Pérusse, Dominique, "New maternity and parental benefits," *Perspectives*, Statistics Canada, (March 2003)

Phipps, S., "Maternity and parental benefits in Canada: are there behavioural implications?," *Canadian Public Policy*, No.4 (2000): 415-436

Phipps, S., P.Burton, and L.Lethbridge, "In and out of the labour market: long-term income consequences of child-related interruptions to women's paid work," *Canadian Journal of Economics*, vol.34, No.2, (2001):411-429

Stewart, Mark, and Christine Greenhalgh, "Work history patterns and the occupational attainment of women," *Economic Journal* 94 (1984):493-519

Stewart, Mark, "On least squares estimation when the dependent variable is grouped," *The Review of Economics Studies*, 50(4) (1983):737-753

Stratton, Leslie, "The effect interruptions in work experience have on wages," *Southern Economic Journal* 61 (1995):955-970

Waldfogel, Jane, "The price of motherhood: family status and women's pay in a young British cohort," *Oxford Economic Papers* 47 (1995): 584-610

Waldfogel, Jane, "The effect of children on women's wages," *American Sociological Review*, 62 (1997a):209-217

Waldfogel, Jane, "Working mothers then and now: a cross-cohort analysis of the effects of maternity leave on women's pay," in *Gender and Family Issues in the Workplace*, ed. Francine D.Blau and Ronald G. Ehrenberg (New York: Russell Sage Foundation) (1997b)

Waldfogel, Jane, "The family gap for young women in the United States and Britain: can maternity leave make a difference?" *Journal of Labour Economics*, 16 (1998a):505-545

Waldfogel, Jane, "Understanding the 'family gap' in pay for women with children," *Journal of Economic Perspectives*, 12 (1998b):137-156

Zveglich, Joseph E., Jr. and Yana van der Meulen Rodgers, "The Impact of Protective Measures for Female Workers," *Journal of Labor Economics*, vol. 21, no. 3 (2003)

Appendix

The diagnostic tests were carried out for specification 5 of Table 10. First, I implemented the Jarque-Bera test and the Chi-squared Goodness of Fit test for normality. The null hypothesis for both tests is that the errors are normally distributed. The results shown in Table A1 indicate that the null hypothesis can be rejected at the 1% level of significance in both cases. That is, the errors of the equation are not normal.

Second, when I used White's heteroskedasticity-consistent covariance matrix to estimate the standard errors of the OLS estimates in specification 5 of Table 10, the p-values of some coefficient estimates changed considerably (see Table A1, which contains just the standard errors, t-ratios, and p-values). The coefficients of some dummy variables become insignificant: for the level of education dummy for other beyond elementary or high school, the p-value changes from 0.029 to 0.252; for the occupational dummy for technologists, technicians and technical occupations, the p-value changes from0.002 to 0.091; for the occupational dummy for clerical occupations, the p-value changes from 0.005 to 0.141; and for the occupational dummy for trades, transport and equipment operators and related occupations, the p-value changes from 0.018 to 0.172. The coefficient of the residence dummy for the province of Saskatchewan becomes significant (its p-value changes from 0.110 to 0.078). The coefficient of duration of demand-related work interruptions becomes more significant (its p-value changes a lot from 0.080 to 0.017). Although the p-values of the coefficients of the dummy variables for "sales and services occupations" and "occupations unique to processing, manufacturing and utilities" increase from 0.473 to 0.704 and from 0.507 to 0.715 respectively, they are still insignificant.

Because of the nonnormality of the errors, the F statistic for the test of overall significance rises will no longer be valid, so I do a Wald test instead to test the null hypothesis that the

coefficients of all the explanatory variables are equal to zero. With using White's heteroskedasticity-consistent covariance matrix, the Wald statistic rises from 1383.0 to 1417.7, but the results of the Wald test do not change as the Wald statistic previously was significant at 1% level.

Tables

TABLE 1

Magnitude of the family gap in different countries and different years

Country	Source	Estimate of family gap	Year	Sample	Magnitude of family gap
The United Kingdom	Waldfogel (1995)	Mean wage femal/male pay ratio	1981	Women aged 23	14%
			1991	Women aged 33	20%
The United States	Waldfogel (1997a)	Mean hourly wage	1988	Women aged 34-44	0.37
	Waldfogel (1998b)	Mean log of hourly wage	1980	Women aged 30	0.17
			1991		0.25
The United States	Harkness and Waldfogel (1999)	Mean wage female/male pay ratio	1994	Women aged 24-44	8.7%
The United Kingdom			1995		4.7%
Canada			1994		5.4%
Australia			1994		1.4%
Germany			1994		9.0%
Finland			1991		2.4%
Sweden			1991		4.6%
Canada	Phipps, Burton, and Lethbridge (2001)	Mean income mothers/non-mothers pay ratio	1995	Women aged 25-54	87.3%

TABLE 2

Magnitude of family gap in different countries and different years after controlling for some characteristics such as experience, education, age, and marital status

Source	Country	Data source	Dependent variable	Estimate of family gap	Years	Sample	Magnitude of family gap
Korenman and Neumark (1992)	The United States	National Longitudinal Survey of Young Women (NLS-YW), panel data	Log of hourly earnings	One child penalty	1982	White women aged 28-38	4%
				Two or more children penalty			7%
Neumark and Korenman (1994)	The United States	Sibling data of the NLS-YW, sibling panel data	Log of hourly earnings	Number of children	1982	White women aged 28-38	7.4%
							3.9% (in fixed-effects equation)
						Black women aged 28-38	0.8%
							3.3% (in fixed-effects equation)
Waldfogel (1997a)	The United States	NLS-YW, panel data	Log of hourly wages	One child penalty	1988	Women aged 34-44	4.1%
				Two or more children penalty			9.6%
				One child penalty	1968-1988	Women aged 34-44	4.1%
				Two or more children penalty			9.6%
Waldfogel (1998b)	The United States	National Longitudinal Survey of Youth (NLSY), panel data	Log of hourly wages	One child penalty	1980	Women with an average age of 30	4.5%
				Two or more children penalty			9.2%
				One child penalty	1991		9.8%
				Two or more children penalty			10.9%

Waldfogel (1995, 1998a)	The United Kingdom	National Child Development Study, panel data	Log of hourly wages	One child penalty	1981-1991	Women aged 23-33	9.6%
							9.1% (in fixed-effects model)
				Two or more children penalty			19.5%
							16.1% (in fixed-effects model)
Harkness and Waldfogel (1999)	The United States	Luxembourg Income Study, microdata	Log of hourly wages	One child penalty	1994	Women aged 24-44	2.5%
				Two children penalty			4.8%
				Three or more children penalty			10.2%
	The United Kingdom			One child penalty	1995		0%
				Two children penalty			17.6%
				Three or more children penalty			29.5%
	Canada			One child penalty	1994		5.8%
				Two children penalty			6.5%
				Three or more children penalty			20.3%
	Australia			One child penalty	1994		12.2%
				Two children penalty			16.7%
				Three or more children penalty			24.4%
	Germany			One child penalty	1994		0%
				Two children penalty			8%
				Three or more children penalty			13.4%

	Finland			One child penalty	1991		5.4%
				Two children penalty			3.4%
				Three or more children penalty			8.2%
	Sweden			One child penalty	1991		0%
				Two children penalty			5.6%
				Three or more children penalty			10.2%
Phipps, Burton, and Lethbridge (2001)	Canada	General Social Survey, microdata	Log of personal annual income	Ever had children penalty	1995	Women aged 25-54	15.7%

TABLE 3

Means and standard errors for all variables for women aged 25-54 with full-year full-time job (dependent variable = natural log of income)

Variable	Mean	Standard error	Minimum value	Maximum value
Income	40,488	24,652	2,500	154,140
Natural log of income	10.46	0.554	7.824	11.946
Potential experience (years)	18.35	9.36	0	42
Potential experience square	424.43	362.1	0	1,764
Actual experience (years)	16.73	8.69	0	40.67
Actual experience square	355.6	316.1	0	1653.8
Dummy = 1 if with Doctorate/ Master's/ Bachelor's degree	0.300	0.458	0	1
Dummy = 1 if with Diploma/ certificate from community college or trade/ technician	0.336	0.472	0	1
Dummy = 1 if with Some university/community college	0.126	0.332	0	1
Dummy = 1 if with High school diploma	0.166	0.372	0	1
Dummy = 1 if with Other beyond elementary/ high school	0.002	0.043	0	1
Dummy = 1 if with Some secondary/elementary/no schooling (reference)	0.070	0.256	0	1
Dummy = 1 if residence in Newfoundland and Labrador	0.051	0.219	0	1
Dummy = 1 if residence in Prince Edward Island	0.024	0.153	0	1
Dummy = 1 if residence in Nova Scotia	0.055	0.228	0	1
Dummy = 1 if residence in Quebec	0.200	0.400	0	1
Dummy = 1 if residence in Ontario	0.297	0.457	0	1
Dummy = 1 if residence in Manitoba	0.048	0.214	0	1
Dummy = 1 if residence in Saskatchewan	0.045	0.208	0	1

Dummy = 1 if residence in Alberta	0.093	0.291	0	1
Dummy = 1 if residence in British Columbia	0.131	0.337	0	1
Dummy = 1 if residence in New Brunswick (reference)	0.055	0.229	0	1
Dummy = 1 if Management occupations	0.091	0.288	0	1
Dummy = 1 if Professional occupations	0.257	0.437	0	1
Dummy = 1 if Technologists, technicians and technical occupations	0.082	0.274	0	1
Dummy = 1 if Clerical occupations	0.270	0.444	0	1
Dummy = 1 if Sales and services occupations	0.235	0.424	0	1
Dummy = 1 if Trades, transport and equipment operators and related occupations	0.018	0.134	0	1
Dummy = 1 if Occupations unique to processing, manufacturing and utilities	0.039	0.195	0	1
Dummy = 1 if occupations unique to primary industries (reference)	0.007	0.083	0	1
Total duration of job interruptions (years)	1.6193	3.274	0	28
Total duration of maternity/ child care interruptions (years)	0.887	2.453	0	25
Total duration of demand-related interruptions (years)	0.120	0.776	0	15
Total duration of health-related interruptions (years)	0.053	0.428	0	10
Total duration of interruptions for other reasons (years)	0.402	1.791	0	28
Dummy = 1 if ever had a child	0.636	0.481	0	1
Dummy = 1 if ever married or common law	0.753	0.432	0	1
Number of observations	2740			

TABLE 4

Means and standard errors for all variables for mothers aged 25-54 with full-year full-time job (dependent variable = natural log of income)

Variable	Mean	Standard error	Minimum value	Maximum value
Income	39,492	25,297	2,500	154,140
Natural log of income	10.42	0.569	7.824	11.946
Potential experience (years)	20.79	8.766	0	42
Potential experience square	508.88	366.2	0	1,764
Actual experience (years)	18.49	8.23	0	40.67
Actual experience square	409.39	318.3	0	1,653.8
Dummy = 1 if with Doctorate/ Master's/ Bachelor's degree	0.229	0.420	0	1
Dummy = 1 if with Diploma/ certificate from community college or trade/ technician	0.355	0.479	0	1
Dummy = 1 if with Some university/community college	0.123	0.329	0	1
Dummy = 1 if with High school diploma	0.195	0.396	0	1
Dummy = 1 if with Other beyond elementary/ high school	0.002	0.048	0	1
Dummy = 1 if with Some secondary/elementary/no schooling (reference)	0.095	0.294	0	1
Dummy = 1 if residence in Newfoundland and Labrador	0.050	0.218	0	1
Dummy = 1 if residence in Prince Edward Island	0.026	0.159	0	1
Dummy = 1 if residence in Nova Scotia	0.057	0.232	0	1
Dummy = 1 if residence in Quebec	0.219	0.413	0	1
Dummy = 1 if residence in Ontario	0.279	0.449	0	1
Dummy = 1 if residence in Manitoba	0.054	0.226	0	1

Dummy = 1 if residence in Saskatchewan	0.050	0.219	0	1
Dummy = 1 if residence in Alberta	0.090	0.286	0	1
Dummy = 1 if residence in British Columbia	0.115	0.319	0	1
Dummy = 1 if residence in New Brunswick (reference)	0.061	0.239	0	1
Dummy = 1 if Management occupations	0.096	0.294	0	1
Dummy = 1 if Professional occupations	0.219	0.414	0	1
Dummy = 1 if Technologists, technicians and technical occupations	0.079	0.269	0	1
Dummy = 1 if Clerical occupations	0.290	0.454	0	1
Dummy = 1 if Sales and services occupations	0.239	0.427	0	1
Dummy = 1 if Trades, transport and equipment operators and related occupations	0.018	0.134	0	1
Dummy = 1 if Occupations unique to processing, manufacturing and utilities	0.050	0.218	0	1
Dummy = 1 if occupations unique to primary industries (reference)	0.009	0.092	0	1
Total duration of job interruptions (years)	2.301	3.842	0	28
Total duration of maternity/ child care interruptions (years)	1.392	2.959	0	25
Total duration of demand-related interruptions (years)	0.127	0.792	0	15
Total duration of health-related interruptions (years)	0.064	0.414	0	7
Total duration of interruptions for other reasons (years)	0.503	2.152	0	28
Dummy = 1 if ever had a child	1	0	0	1
Dummy = 1 if ever married or common law	0.915	0.279	0	1
Number of observations	1743			

TABLE 5

Means and standard errors for all variables for non-mothers aged 25-54 with full-year full-time job (dependent variable = natural log of income)

Variable	Mean	Standard error	Minimum value	Maximum value
Income	42,230	23,391	2,500	154,140
Natural log of income	10.52	0.520	7.824	11.946
Potential experience (years)	14.10	8.836	0	37
Potential experience square	276.8	302.9	0	1,369
Actual experience (years)	13.67	8.65	0	36
Actual experience square	261.7	289.3	0	1,296
Dummy = 1 if with Doctorate/ Master's/ Bachelor's degree	0.424	0.494	0	1
Dummy = 1 if with Diploma/ certificate from community college or trade/ technician	0.302	0.459	0	1
Dummy = 1 if with Some university/community college	0.131	0.338	0	1
Dummy = 1 if with High school diploma	0.114	0.318	0	1
Dummy = 1 if with Other beyond elementary/ high school	0.001	0.032	0	1
Dummy = 1 if with Some secondary/elementary/no schooling (reference)	0.070	0.256	0	1
Dummy = 1 if residence in Newfoundland and Labrador	0.052	0.222	0	1
Dummy = 1 if residence in Prince Edward Island	0.021	0.144	0	1
Dummy = 1 if residence in Nova Scotia	0.052	0.222	0	1
Dummy = 1 if residence in Quebec	0.168	0.374	0	1
Dummy = 1 if residence in Ontario	0.327	0.469	0	1
Dummy = 1 if residence in Manitoba	0.038	0.192	0	1

Dummy = 1 if residence in Saskatchewan	0.036	0.187	0	1
Dummy = 1 if residence in Alberta	0.100	0.301	0	1
Dummy = 1 if residence in British Columbia	0.159	0.366	0	1
Dummy = 1 if residence in New Brunswick (reference)	0.055	0.229	0	1
Dummy = 1 if Management occupations	0.083	0.276	0	1
Dummy = 1 if Professional occupations	0.322	0.467	0	1
Dummy = 1 if Technologists, technicians and technical occupations	0.087	0.282	0	1
Dummy = 1 if Clerical occupations	0.236	0.425	0	1
Dummy = 1 if Sales and services occupations	0.229	0.420	0	1
Dummy = 1 if Trades, transport and equipment operators and related occupations	0.018	0.133	0	1
Dummy = 1 if Occupations unique to processing, manufacturing and utilities	0.021	0.144	0	1
Dummy = 1 if occupations unique to primary industries (reference)	0.007	0.083	0	1
Total duration of job interruptions (years)	0.427	1.192	0	14.75
Total duration of maternity/ child care interruptions (years)	0.003	0.095	0	3
Total duration of demand-related interruptions (years)	0.108	0.747	0	14.75
Total duration of health-related interruptions (years)	0.061	0.451	0	10
Total duration of interruptions for other reasons (years)	0.225	0.819	0	8
Dummy = 1 if ever had a child	0	0	0	1
Dummy = 1 if ever married or common law	0.468	0.499	0	1
Number of observations	997			

TABLE 6

OLS regression results for age 25-54, full-year full-time workers (dependent variable = natural log of income) [a)]

	Specification 1		Specification 2	
	Men	Women	Men	Women
Potential experience	0.017*	0.024*	0.016*	0.024*
	(0.004)	(0.004)	(0.004)	(0.004)
Potential experience squared	-0.0002**	-0.0004*	-0.0002**	-0.0004*
	(0.0001)	(0.00009)	(0.0001)	(0.00009)
Dummy = 1 if ever had a child	0.130*	-0.049**	0.067*	-0.073*
	(0.020)	(0.019)	(0.023)	(0.022)
Dummy = 1 if ever married			0.121*	0.051**
			(0.024)	(0.022)
Dummy = 1 if with Doctorate/Master's/Bachelor's degree	0.377*	0.559*	0.366*	0.556*
	(0.036)	(0.039)	(0.036)	(0.039)
Dummy = 1 if with Diploma/certificate from community college or trade/technician	0.178*	0.274*	0.169*	0.273*
	(0.031)	(0.036)	(0.031)	(0.036)
Dummy = 1 if with Some university/community college	0.180*	0.339*	0.177*	0.339*
	(0.037)	(0.041)	(0.037)	(0.040)
Dummy = 1 if with High school diploma	0.128*	0.187*	0.122*	0.187*
	(0.033)	(0.038)	(0.033)	(0.038)
Dummy = 1 if with Other beyond elementary/high school	0.406*	0.433**	0.399*	0.440**
	(0.133)	(0.191)	(0.133)	(0.191)
Dummy = 1 if residence in Newfoundland and Labrador	0.185*	-0.035	0.175*	-0.037
	(0.061)	(0.052)	(0.060)	(0.052)
Dummy = 1 if residence in Prince Edward Island	-0.092	0.057	-0.103	0.058
	(0.083)	(0.064)	(0.083)	(0.064)
Dummy = 1 if residence in Nova Scotia	0.063	0.037	0.055	0.037
	(0.056)	(0.051)	(0.056)	(0.051)

Dummy = 1 if residence in Quebec	0.058	0.028	0.076***	0.037
	(0.044)	(0.040)	(0.044)	(0.041)
Dummy = 1 if residence in Ontario	0.242*	0.258*	0.236*	0.258*
	(0.043)	(0.039)	(0.043)	(0.039)
Dummy = 1 if residence in Manitoba	0.104***	0.045	0.100***	0.046
	(0.055)	(0.052)	(0.055)	(0.052)
Dummy = 1 if residence in Saskatchewan	0.186*	0.049	0.184*	0.051
	(0.054)	(0.053)	(0.054)	(0.053)
Dummy = 1 if residence in Alberta	0.196*	0.088**	0.196*	0.089**
	(0.048)	(0.045)	(0.048)	(0.045)
Dummy = 1 if residence in British Columbia	0.190*	0.137*	0.189*	0.140*
	(0.047)	(0.043)	(0.046)	(0.043)
Dummy = 1 if Management occupations	0.145*	0.638*	0.134*	0.637*
	(0.050)	(0.107)	(0.050)	(0.107)
Dummy = 1 if Professional occupations	0.107**	0.643*	0.100**	0.645*
	(0.049)	(0.105)	(0.049)	(0.105)
Dummy = 1 if Technologists, technicians and technical occupations	0.013	0.452*	0.009	0.455*
	(0.053)	(0.107)	(0.053)	(0.107)
Dummy = 1 if Clerical occupations	-0.167*	0.416*	-0.170*	0.415*
	(0.053)	(0.104)	(0.053)	(0.104)
Dummy = 1 if Sales and services occupations	-0.162*	0.178***	-0.163*	0.179***
	(0.048)	(0.105)	(0.048)	(0.104)
Dummy = 1 if Trades, transport and equipment operators and related occupations	-0.088***	0.384*	-0.092**	0.394*
	(0.046)	(0.119)	(0.046)	(0.119)
Dummy = 1 if Occupations unique to processing, manufacturing and utilities	-0.123**	0.192***	-0.126**	0.190***
	(0.051)	(0.112)	(0.051)	(0.111)
Constant term	10.165*	9.311*	10.155*	9.299*
	(0.072)	(0.120)	(0.072)	(0.120)
Adjusted R^2	0.155	0.322	0.161	0.323
F-statistic	28.89*	60.54*	28.94*	58.42*

Number of observations	3649	3015	3649	3015

Standard errors are in parentheses.
* significant with 99 per cent confidence
** significant with 95 per cent confidence
*** significant with 90 per cent confidence

a) Some secondary/elementary/no schooling is the reference level of education; New Brunswick is the reference province; Occupations unique to primary industries is the reference occupation.

TABLE 7

OLS regression results for women aged 25-54, full-year full-time workers (dependent variable = natural log of income)

	Specification 1	Specification 2
Potential experience	0.027*	0.026*
	(0.004)	(0.004)
Potential experience squared	-0.0004*	-0.0004*
	(0.0001)	(0.0001)
Dummy = 1 if ever had a child	-0.054*	-0.077*
	(0.020)	(0.023)
Dummy = 1 if ever married		0.049**
		(0.023)
Dummy = 1 if with Doctorate/Master's/Bachelor's degree	0.832*	0.829*
	(0.038)	(0.038)
Dummy = 1 if with Diploma/certificate from community college or trade/technician	0.410*	0.408*
	(0.036)	(0.036)
Dummy = 1 if with Some university/community college	0.474*	0.474*
	(0.042)	(0.042)
Dummy = 1 if with High school diploma	0.265*	0.264*
	(0.039)	(0.039)
Dummy = 1 if with Other beyond elementary/high school	0.569*	0.576*
	(0.202)	(0.202)
Dummy = 1 if residence in Newfoundland and Labrador	0.027	0.025
	(0.055)	(0.055)
Dummy = 1 if residence in Prince Edward Island	0.076	0.077
	(0.068)	(0.068)
Dummy = 1 if residence in Nova Scotia	0.042	0.041
	(0.054)	(0.054)
Dummy = 1 if residence in Quebec	0.050	0.058
	(0.043)	(0.043)

Dummy = 1 if residence in Ontario	0.271*	0.271*
	(0.041)	(0.041)
Dummy = 1 if residence in Manitoba	0.044	0.045
	(0.055)	(0.055)
Dummy = 1 if residence in Saskatchewan	0.051	0.052
	(0.057)	(0.057)
Dummy = 1 if residence in Alberta	0.114**	0.115**
	(0.048)	(0.048)
Dummy = 1 if residence in British Columbia	0.141*	0.143*
	(0.045)	(0.045)
Constant term	9.533*	9.523*
	(0.061)	(0.062)
Adjusted R^2	0.235	0.236
F-statistic	55.38*	52.61*
Number of observations	3015	3015

Standard errors are in parentheses.
* significant with 99 per cent confidence
** significant with 95 per cent confidence
*** significant with 90 per cent confidence

TABLE 8 [a)]

Percentage with an interruption of more than 6 months (with an interruption of more than 3 months in parentheses), mean duration and reason for interruptions, [b) c)] men and women aged 25-54

variable	Full time and part time		Full time only	
	Men	Women	Men	Women
Any job interruptions	12.7% (16.2%)	37.4% (42.9%)	12.1% (15.6%)	35.7% (41.5%)
Mean total duration, all with an interruption (years)	1.70 (1.43)	3.53 (3.17)	1.64 (1.37)	3.35 (2.97)
Any interruption for "demand" reasons [d)]	63.7% (64.8%)	14.5% (15.5%)	62.9% (63.7%)	14.0% (14.9%)
Any interruption for health reasons	30.0% (31.2%)	9.6% (13.1%)	30.2% (31.1%)	8.9% (9.3%)
Any interruption for family reasons	10.5% (10.4%)	84.6% (84.0%)	10.8% (10.9%)	84.6% (84.3%)
Any family interruption for maternity/paternity leave	10.3% (15.3%)	79.4% (82.1%)	9.1% (15.3%)	80.1% (83.4%)
Any family interruption for child care	4.4% (3.5%))	11.3% (10.2%)	1.8% (1.4%)	11.2% (9.9%)
Any family interruption for other family reasons [e)]	85.3% (81.2%)	18.0% (17.3%)	89.1% (83.3%)	16.8% (15.7%)
Number of observations	5081	5049	4225	3555

a) The design of this table is basically the same as Table 2 in Phipps, Burton, and Lethbridge (2001).

b) Those who gave education, immigration, personal choice, changed jobs, retired or other as a reason for any interruption are excluded. This table only includes demand-related, health-related, and family-related work interruptions, and maternity/paternity leave, child care and other family reasons are the subset reasons of family-related work interruptions.

c) The percentages may add to more than 100 because a respondent can have more than one interruption and/or more than one reason for any particular interruption.

d) Includes seasonal worker, temporary layoff/end of contract, lack of work, business or company closure.

e) Includes moved or transferred, elder care, marriage, personal or family responsibilities.

TABLE 9 [a)]

Percentage with an interruption of more than 6 months (with an interruption of more than 3 months in parentheses), mean duration and reason for interruptions, women without child and women with child aged 25-54 full-time workers

variable	Full time only	
	Non-mothers	Mothers
Any job interruptions	10.6%	49.2%
	(14.7%)	(56.0%)
Mean total duration, all with an interruption (years)	1.59	3.55
	(1.27)	(3.22)
Any interruption for "demand" reasons	53.8%	9.4%
	(51.4%)	(9.7%)
Any interruption for health reasons	27.3%	6.8%
	(35.5%)	(9.1%)
Any interruption for family reasons	23.5%	91.7%
	(19.7%)	(93.4%)
Any family interruption for maternity/paternity leave	3.2%	82.3%
	(2.8%)	(85.8%)
Any family interruption for child care	0% [b)]	11.5%
	(0%)	(10.2%)
Any family interruption for other family reasons	96.8	14.4%
	(97.2%)%	(13.3%)
Number of observations	1244	2311

a) The definitions of variables are the same as Table 8.

b) This is for women without child, so the percentage of any child care reason is 0.

TABLE 10

OLS regression results for women aged 25-54, full-year full-time workers (dependent variable = natural log of income)

	Specification 3	Specification 4	Specification 5
Actual experience	0.029*	0.030*	0.030*
	(0.004)	(0.004)	(0.004)
Actual experience squared	-0.0005*	-0.0005*	-0.0005*
	(0.0001)	(0.0001)	(0.0001)
Dummy = 1 if ever had a child	-0.055**	-0.045**	-0.048**
	(0.022)	(0.022)	(0.022)
Dummy = 1 if ever married or common law	0.036	0.038	0.036
	(0.023)	(0.023)	(0.023)
Dummy = 1 if with Doctorate/Master's/Bachelor's degree	0.580*	0.576*	0.574*
	(0.041)	(0.041)	(0.041)
Dummy = 1 if with Diploma/certificate from community college or trade/technician	0.299*	0.294*	0.293*
	(0.038)	(0.038)	(0.038)
Dummy = 1 if with Some university/community college	0.366*	0.361*	0.359*
	(0.043)	(0.043)	(0.043)
Dummy = 1 if with High school diploma	0.217*	0.214*	0.213*
	(0.040)	(0.038)	(0.040)
Dummy = 1 if with Other beyond elementary/high school	0.454**	0.455**	0.450**
	(0.206)	(0.206)	(0.206)
Dummy = 1 if residence in Newfoundland and Labrador	-0.002	-0.0006	-0.002
	(0.053)	(0.053)	(0.053)
Dummy = 1 if residence in Prince Edward Island	0.087	0.089	0.089
	(0.067)	(0.067)	(0.067)
Dummy = 1 if residence in Nova Scotia	0.037	0.039	0.036
	(0.052)	(0.052)	(0.052)
Dummy = 1 if residence in Quebec	0.050	0.054	0.052
	(0.042)	(0.042)	(0.042)

Dummy = 1 if residence in Ontario	0.266*	0.267*	0.264*
	(0.040)	(0.040)	(0.040)
Dummy = 1 if residence in Manitoba	0.045	0.048	0.044
	(0.054)	(0.054)	(0.054)
Dummy = 1 if residence in Saskatchewan	0.087	0.092***	0.088
	(0.055)	(0.055)	(0.055)
Dummy = 1 if residence in Alberta	0.082***	0.086***	0.082***
	(0.047)	(0.047)	(0.047)
Dummy = 1 if residence in British Columbia	0.153*	0.157*	0.155*
	(0.044)	(0.044)	(0.044)
Dummy = 1 if Management occupations	0.524*	0.520*	0.516*
	(0.108)	(0.108)	(0.108)
Dummy = 1 if Professional occupations	0.534*	0.530*	0.528*
	(0.106)	(0.106)	(0.106)
Dummy = 1 if Technologists, technicians and technical occupations	0.344*	0.341*	0.339*
	(0.109)	(0.109)	(0.109)
Dummy = 1 if Clerical occupations	0.298*	0.296*	0.294*
	(0.106)	(0.105)	(0.105)
Dummy = 1 if Sales and services occupations	0.081	0.078	0.076
	(0.106)	(0.106)	(0.106)
Dummy = 1 if Trades, transport and equipment operators and related occupations	0.288**	0.286**	0.289**
	(0.123)	(0.123)	(0.123)
Dummy = 1 if Occupations unique to processing, manufacturing and utilities	0.077	0.072	0.075
	(0.114)	(0.114)	(0.114)
Total duration of job interruptions		-0.006**	
		(0.003)	
Total duration of maternity/child care interruptions			-0.003
			(0.004)
Total duration of demand-related interruptions			-0.020***
			(0.011)
Total duration of health-related interruptions			-0.028

			(0.020)
Total duration of interruptions for other reasons			-0.010**
			(0.005)
Constant term	9.341*	9.343*	9.353*
	(0.122)	(0.122)	(0.122)
Adjusted R^2	0.329	0.330	0.331
F-statistic	54.81*	52.97*	47.69*
Number of observations	2740	2740	2740

Standard errors are in parentheses.
* significant with 99 per cent confidence
** significant with 95 per cent confidence
*** significant with 90 per cent confidence

TABLE A1

OLS regression results for women aged 25-54, full-year full-time workers (dependent variable = natural log of income) and the results using White's heteroskedasticity-consistent covariance matrix in parentheses, the same equation as specification 5 in Table 10 [a)]

Variable	Standard error	T-ratio	P-value
Actual experience	0.004	7.494	0.000
	(0.004)	(7.336)	(0.000)
Actual experience squared	0.0001	-4.499	0.000
	(0.0001)	(-4.319)	(0.000)
Dummy = 1 if ever had a child	0.022	-2.173	0.030
	(0.021)	(-2.262)	(0.024)
Dummy = 1 if ever married or common law	0.023	1.548	0.122
	(0.022)	(1.629)	(0.103)
Dummy = 1 if with Doctorate/Master's/Bachelor's degree	0.041	14.01	0.000
	(0.041)	(14.01)	(0.000)
Dummy = 1 if with Diploma/certificate from community college or trade/technician	0.038	7.683	0.000
	(0.038)	(7.746)	(0.000)
Dummy = 1 if with Some university/community college	0.043	8.384	0.000
	(0.043)	(8.443)	(0.000)
Dummy = 1 if with High school diploma	0.040	5.289	0.000
	(0.039)	(5.430)	(0.000)
Dummy = 1 if with Other beyond elementary/high school	0.206	2.184	0.029
	(0.393)	(1.145)	(0.252)
Dummy = 1 if residence in Newfoundland and Labrador	0.053	-0.043	0.996
	(0.051)	(-0.045)	(0.964)
Dummy = 1 if residence in Prince Edward Island	0.067	1.321	0.187
	(0.060)	(1.470)	(0.142)
Dummy = 1 if residence in Nova Scotia	0.052	0.683	0.495
	(0.044)	(0.804)	(0.421)

Dummy = 1 if residence in Quebec	0.042	1.255	0.209
	(0.039)	(1.351)	(0.177)
Dummy = 1 if residence in Ontario	0.040	6.565	0.000
	(0.038)	(6.968)	(0.000)
Dummy = 1 if residence in Manitoba	0.054	0.818	0.413
	(0.051)	(0.875)	(0.382)
Dummy = 1 if residence in Saskatchewan	0.055	1.599	0.110
	(0.050)	(1.763)	(0.078)
Dummy = 1 if residence in Alberta	0.047	1.759	0.079
	(0.044)	(1.864)	(0.062)
Dummy = 1 if residence in British Columbia	0.044	3.510	0.000
	(0.043)	(3.629)	(0.000)
Dummy = 1 if Management occupations	0.108	4.777	0.000
	(0.202)	(2.551)	(0.011)
Dummy = 1 if Professional occupations	0.106	4.980	0.000
	(0.200)	(2.638)	(0.008)
Dummy = 1 if Technologists, technicians and technical occupations	0.109	3.127	0.002
	(0.201)	(1.690)	(0.091)
Dummy = 1 if Clerical occupations	0.105	2.784	0.005
	(0.199)	(1.472)	(0.141)
Dummy = 1 if Sales and services occupations	0.106	0.718	0.473
	(0.200)	(0.380)	(0.704)
Dummy = 1 if Trades, transport and equipment operators and related occupations	0.123	2.358	0.018
	(0.212)	(1.368)	(0.172)
Dummy = 1 if Occupations unique to processing, manufacturing and utilities	0.114	0.663	0.507
	(0.206)	(0.366)	(0.715)
Total duration of maternity/child care interruptions	0.004	-0.928	0.354
	(0.004)	(-0.970)	(0.332)
Total duration of demand-related interruptions	0.011	-1.752	0.080
	(0.008)	(-2.379)	(0.017)
Total duration of health-related interruptions	0.020	-1.381	0.167