Sons, Daughters, and Family Structure

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1. INTRODUCTION

There is a substantial body of literature documenting the effect of parenthood on individuals' decisions to marry or divorce. For example, studies show that couples with children are less likely to divorce than those without (e.g. Becker et al, 1977), and that this likelihood is inversely related to a couple's family size. However, a causal effect of child-rearing on marital stability cannot be determined. A couple's decision to have children – and when to have them – may be correlated with their expectations of marital stability. That is, a couple may decide to postpone child-bearing or to have less children if they expect their marriage to dissolve. Moreover, parents who decide to have (more) children may differ than those who do not, in terms of additional unobservable characteristics which are, in turn, correlated with their likelihood to end their marriage.

One way to study the effect of children on their parents' marital outcomes is to focus on child gender. More specifically, one can estimate the effect of the gender of a first-born child on a couple's probability of marrying or divorcing. Because the gender of a first-born child is randomly assigned to parents, it is not likely to be correlated with any of the parents' unobserved characteristics which could influence their marital decisions. Therefore, a causal effect of a child's gender on their parents' divorce outcomes can be determined. On the other hand, the gender of a child born after the couple's first-born child may not be considered exogenous to their likelihood of marrying or divorcing. In other words, parents who give birth to a first-born daughter, and then decide to have a second child, may have different characteristics than parents who give birth to a first-born son before deciding to have a second child – these characteristics may be correlated with parents' marital outcomes.

Using data from developed countries, researchers have that found parents of first-born daughters are more likely to divorce than those of first-born sons (e.g. Spanier and Glick, 1981; Morgan, Lye and Condran, 1988; Dahl and Moretti, 2008), although this effect may only exist for earlier marriage cohorts (Morgan and Pollard, 2003). Sons have also been shown to reduce both a mother's likelihood of anticipating divorce (Katzev, Warner and Acock, 1994), and a father's probability of leaving a household (Mott, 1994; Lundberg, McLanahan and Rose 2007), and to increase a divorced father's likelihood of obtaining custody of his children (Dahl and Moretti, 2008). Studies have also shown that having a daughter rather than a son negatively affects a single mother's likelihood of ever marrying (Lundberg and Rose, 2003; Dahl and Moretti, 2008). There is only one such study (to my knowledge) that uses Canadian data. In a cross-country study, using the OECD Fertility and Family Survey (FFS), Diekmann and Schmidheiny (2004) find that the divorce rate among Canadian couples with an only son is 35 percent lower than among those with an only daughter. However, out of eighteen different OECD countries, only four samples provided an estimate with the expected sign, and the Canadian sample was the only sample to provide a statistically significant estimate. Moreover, the authors found no statistically significant effect of child gender on divorce for couples with more than one child in Canada or in any other country.

The goal of this paper is to replicate certain methods of Dahl and Moretti (2008) and Lundberg and Rose (2003) in order to estimate the effect of child gender on family structure using Canadian data. That is, this paper aims to test the hypothesis that parents of first-born daughters are more likely to divorce, and are less likely to marry following a non-marital birth, compared to parents of first-born sons.

Why might child gender affect parents' marriage and divorce probabilities? Although not all authors explore the mechanisms behind this causal relationship, two main explanations can be found in the literature (e.g. Morgan, Lye and Condran, 1988; Lundberg and Rose, 2003; Dahl and Moretti, 2008). The first explanation is that fathers have a preference for sons over daughters, so that they derive more utility from living in a household in which a son is present, compared to a daughter. For example, this may be because fathers feel they can better relate to sons than daughters and can engage in more activities with them. A second explanation is that fathers do not have a preference for sons over daughters, but that instead, one or both of the parents believe that paternal absenteeism affects the well-being of sons and daughters differently. Parents may believe that fathers are more important to the development of sons than daughters, perhaps because fathers are (or are at least believed to be) more important role models for sons than for daughters. In this case, sons would suffer more than daughters from losing their father from their household, in terms of the quality or amount of parenting they receive. If a parent's utility is a function of their children's well-being, they incur a bigger loss in utility from divorce if they have a son.

Using the 2001 General Social Survey (GSS), this paper finds that first-born girls are 1.6 percent less likely to live with their biological father than first-born boys, and that parents of first-born girls are 4.5 percent more likely to be lone parents than those of first-born boys; although these estimates are not statistically significant, they are robust with the inclusion of control variables accounting for parental and child characteristics. This paper also finds that parents of first-born girls are 6.8 percent more likely than those of first-born boys to be separated or divorced at the time of the survey; this estimate is robust although statistically insignificant. When the sample is restricted to fathers, a robust and statistically significant estimate implies that fathers are 18.5 percent more likely to be separated or divorced if their first-born child is a girl; however, there is no effect of child gender on a mother's probability of being separated or divorced at the time of the survey. Among parents who had their first-born child during their first marriage, parents of first-born girls are 9.5 percent more likely to have their first marriage end in divorce than parents of first-born boys; this estimate is robust and statistically significant at the 10 percent level. This study also finds that the gender of a first-born child does not have a statistically significant effect on a parent's likelihood of experiencing a shotgun marriage, marrying following a non-marital birth, or remarrying following a divorce.

This paper is divided as follows: Section 2 is a review of literature examining the effect of child gender on family structure; Section 3 is an overview of the data used; Section 4 presents the econometric model and an analysis of the main findings; and Section 5 concludes.

2. LITERATURE REVIEW

A causal effect of child-rearing on marital stability cannot be determined, because the decision to have children (and when to have them) may not be independent of a couple's expectations of marital stability – in other words, a couple may decide to postpone child-bearing or to have less children if they expect their marriage to dissolve. The endogeneity problem associated with estimating the impact of family size on marital stability led researchers to study the effect of child gender on marital stability. Because the gender of a first-born child is virtually randomly assigned to parents, it is not likely to be correlated with any of the parents' unobserved characteristics which could influence their marital decisions. Therefore, a causal effect of a child's gender on their parents' divorce outcomes can be determined. Although the studies which will be reviewed do not focus solely on first-born children or on marital stability, they all examine the link between child gender and family structure.

The first study to document a direct link between child gender and marital stability is Spanier and Glick (1981). Using 1970 Census data, they look at a sample of mothers who first married between 1950 and 1970 under the age of 30. Among women with two children, the ratio of separated and divorced women to women in intact first marriages is 0.28 for women with two daughters, compared to 0.24 for those with two sons and 0.23 for those with one child of each gender. In other words, mothers who have at least one son are more likely to remain in their first marriage than mothers with no sons. This differential is larger among mothers with less than a high school education: the ratio of separated or divorced women to women in intact first marriages is 0.72 for women with two daughters, compared to 0.52 for women with two boys, and 0.43 for those with one child of each gender. For college educated women, the ratio differences are smaller, and in some cases reversed. Using data from the 1975 Current Population Survey

(CPS), they also find that boys under 18 years of age are 5 percent more likely to live in a two-parent household than their female counterparts. In light of these results, Spanier and Glick (1981) hypothesize that fathers may strongly desire a son – perhaps to carry on their family name or their family business – and that failure to fulfill their desire leads to divorce. They also speculate that mothers may find it a more daunting task to raise a son without a husband, leading them to avoid divorce. Because the result is much less evident among a sample of less-educated women, the authors suggest that the correlation between child gender and marital status will disappear in future years, as the average American woman becomes more educated.

Morgan, Lye and Condran (1988) use the June 1980 Current Population Survey to estimate the effect of child gender on parents' risk of divorce. They use a sample of white women who were married before the age of forty, and after January 1960. They find that, while all children promote marital stability, couples with only one daughter are 9 percent more likely to divorce than couples with only one son. For two-child families, having both a daughter and a son makes couples 9 percent more likely to divorce than families with two sons, and having two daughters makes them 18 percent more likely to divorce.

According to Morgan, Lye and Condran (1988), the effect of daughters on divorce risk may be due to the belief that fathers are more important to the social development of sons than daughters. Because fathers do not usually get sole custody of children, divorce would likely lead to a decrease in the child's contact with their father. Therefore, parents believe that divorce would adversely affect the social development of a son to a larger degree than that of a daughter, leading parents of sons to be less likely to divorce. Moreover, because fathers are believed to be more crucial to the development of sons than daughters, fathers of sons are more actively involved in child-rearing than fathers of daughters. The higher involvement of fathers of sons in family life, in turn, promotes marital stability because parents are more emotionally connected. Another explanation, based on the theory of Becker, Landes and Michael (1977), is that fathers of sons make more marriage-specific investments than fathers of daughters, as a result of their higher degree of involvement in family life. Marriage-specific investments (i.e. investments which are valuable within a marriage, but have no value outside of a marriage) are inversely related to divorce rates, which makes fathers of sons less likely to divorce than fathers of daughters. Using the National Longitudinal Survey of Youth (NLSY), Morgan, Lye and Condran (1988) find that mothers are more likely to perceive their sons, rather than their daughters, as "close" to their fathers. Sons are themselves more likely to report themselves "close" to their fathers than are daughters. They also find that mothers spend more time with daughters than sons, and that fathers spend more time with sons than daughters. Lastly, the reports of children and mothers indicate that fathers are more involved in the disciplining and rule-setting of sons rather than daughters. These findings support the theory that the higher probability of marital stability observed in couples with sons is due to fathers' greater parental involvement with sons.

Morgan and Pollard (2003) replicate the analysis of Morgan, Lye and Condran (1988) using more recent American data. Using retrospective marriage histories from the 1980, 1985, 1980 and 1995 June Current Population Surveys, they estimate the effect of the number of daughters on the probability of a first marriage ending in divorce. For the 1960-1979 marriage cohorts, the probability of a first marriage ending in divorce increases by 8 percent for each daughter, relative to each son. This finding is consistent with Morgan, Lye and Condran (1988), and the effect still holds when data from the 1980 wave are omitted (the wave used by Morgan, Lye and Condran, 1988). They also compare the effect of child gender composition on divorce probability using different variables to account for child gender composition. For the 1960-1979 period, they find that the risk of a first marriage ending in divorce is higher for a couple with a first-born daughter, and for a couple with all daughters; the risk is lower for a couple with all boys. The authors find that the number of daughters in the family has the greatest explanatory power when estimating the effect of child gender on marital stability. However, for the 1980-1994 marriage cohorts, Morgan and Pollard (2003) find no statistically significant effect of the number of daughters on marital stability. The authors theorize that the attenuation of the child gender effect may be due to the fact that gender roles have become more egalitarian. This societal change has led to an attenuation of the belief that fathers are relatively more important role in the parenting of sons versus daughters. The change has also made parents view boys and girls as perfect substitutes, so that they are indifferent between having a son or a daughter. In fact, Morgan and Pollard (2002) study American fertility data and find that couples have become more indifferent between the birth of a son and that of a daughter after 1985. That is, from the 1960s and prior to 1985, couples with two children of the same gender were more likely to have a third child, implying a preference for one child of each gender. They find that this effect has weakened since 1985.

Diekmann and Schmidheiny (2004) provide a cross-national comparison of the effect of child gender on the probability of divorce. Their goal is to replicate the results of Morgan, Lye and Condran (1988). They also aim to test the theory advanced by these authors that sons lead to increased paternal involvement, which increases solidarity and companionship within a marriage, and thus leads to a smaller risk of divorce. They estimate the effect of child gender on a married couple's risk of divorce using the cross-national Fertility and Family Surveys (FFS) of 18 different industrialized countries from the early nineties (including Canada, the US and 16 European countries). They use 19 samples which consist of married and previously-married women, and include marriage cohorts from the 1970s up to the early 1990s.

Among one-child families, those with a son are less likely to divorce in 11 of the 18 samples. In 4 of the samples, there is no effect of child gender on divorce risk, while the remaining 4 samples show a negative effect of daughters on divorce risk. Although most of the samples provide the expected sign for the child gender effect, the estimates are not statistically significant, with the exception of Canada. For Canada, they find that the divorce rate among couples with a son is 35 percent lower than among couples with a daughter, and this is significant at the 5 percent level. The authors find no statistically significant effects of gender composition on divorce for samples of two-child families. Moreover, the signs of the estimates vary between the countries. The expected pattern is only found for 11 countries; in these samples, mothers with two boys are at less risk of divorce than those with one child of each gender, while mothers with two daughters have the highest risk.

Diekmann and Schmidheiny (2004) test the paternal involvement hypothesis explored by Morgan, Lye and Condran (1988). For eight countries, they use samples of married female respondents who were raising one child under the age of 15 at the time of the survey. Respondents were asked to report their husbands' degree of involvement in caring for each of the child's following needs: being fed, being dressed, being cared for during illness, being played with, and being helped with homework. A statistically significant child gender effect in paternal involvement was only found for one of these dimensions in Spain (48 percent of Spanish fathers cared for a son during illness, while 37 percent cared for a daughter, according to their wives' reports). Because there is no systematic effect of child gender on paternal involvement across the samples, the paternal involvement hypothesis is not supported.

Following Morgan and Pollard (2003), Diekmann and Schmidheiny (2004) test whether there is a significant child gender effect for earlier marriage cohorts; they compare the effect of child gender on divorce for marriages contracted before 1979 to those contracted after 1979. The attenuation hypothesis implies that the effect of sons on divorce would be negative and significant in the pre-1979 cohorts, and then disappear in the post-1979 cohorts. This is the case for one sample only (Switzerland) – the cross-country comparison therefore does not support the attenuation hypothesis of Morgan and Pollard (2003).

Mott (1994) examines the relation between the gender of children and their fathers' probabilities of leaving the household, and how this relation varies between white and black fathers. Using American data from the National Longitudinal Survey of Youth (NLSY), the author finds evidence of a slight preference of white fathers for sons over daughters. For white families, boys are more likely to have a father in their household than are girls. Interestingly, this effect is only statistically significant when controls are included to account for family and maternal characteristics. The author explains that maternal characteristics such as a mother's likelihood of participating in the labour force may be related to both child gender and the likelihood of living without the child's father. He also finds that white fathers who do not live with their children maintain contact with girls just as much as they do with boys. He finds no evidence of paternal gender bias in black families – that is, black fathers are equally likely to reside or maintain regular contact with sons as they are with daughters.

Dahl and Moretti (2008) study the effect of gender on a child's probability of living without a father, and identify three channels leading to this effect. Using data from the 1960 to 2000 U.S. Censuses, they show that a first-born daughter is 3.1 percent less likely to be living with her father compared to a first-born son. In their model, the likelihood of a child living without a father is a function of three factors: the child's mother's probability of ever marrying, her probability of divorcing (conditional on marriage), and her probability of awarding custody to the child's father (conditional on divorcing).

Firstly, Dahl and Moretti (2008) show that mothers of first-born daughters are 1.4 percent more likely to have never been married than those of first-born sons. The authors also find, using California Birth Certificate data, that the gender of an unborn child affects the probability of marriage at delivery. That is, among women who have taken an ultrasound test during pregnancy (and thus presumably know the gender of their unborn child), those who give birth to a girl are approximately 5 percent less likely to be married at delivery than those who give birth to a boy. This effect of child gender on marital status at delivery is only observed for women who have taken an ultrasound – there is no effect when they use the entire sample of first-time mothers.

Secondly, conditional on having ever married, parents who have first-born girls are 1.3 percent more likely to be divorced at the time of the survey. For larger families, the child gender effects increase; however, the estimates can no longer be interpreted as causal because the genders of the second and third child cannot be assumed independent of other characteristics affecting parents' divorce probability. Parents whose two oldest children are girls are 1.9 percent more likely to be divorced than those whose two oldest children are boys. Those who had three daughters first are 3.2 percent more likely to be divorced than those who had three sons first. Because the gender of a child could theoretically affect a divorced parent's probability of remarriage, they also look at the effect of child gender on the probability of a parent's *first* marriage ending in divorce. They find that the first-born daughter effect is slightly higher for the probability of a first marriage ending in divorce than for the probability of being divorced at the time of the survey.

Thirdly, divorced fathers are less likely to obtain custody of their daughters than their sons. Fathers have a 14.7 percent probability of obtaining custody of their first-born son, compared to a 12.2 percent probability of obtaining custody of their first-born daughter. Dahl and Moretti (2008) state that these empirical findings, taken together, suggest that American parents may prefer boys over girls. They find some evidence of a child gender preference; on average, families with first-born girls are more likely to have more children than those with first-born boys. The total number of children is 0.3 percent higher in families with a first-born daughter. But the theory and the empirical findings do not indicate whether the preference for sons is exhibited by the mother, the father, or both. They find evidence from 2000 and 2003 Gallup polls that fathers largely prefer sons, and mothers marginally prefer daughters; fathers' gender bias is found to be relatively stronger than that of mothers. Thus, the authors conclude that the effect of child gender on parents' marriage and fertility decisions may be due to fathers' gender preferences.

Lundberg and Rose (2003) look at the effect of child gender on mothers' probability of marriage following a non-marital birth, and on their probability of remarriage following a divorce. They argue that, because having a son rather than a daughter leads parents to place a higher value on marriage relative to divorce, unmarried couples who give birth to a son should be more likely to wed. Firstly, if sons lead to more successful marriages, then they may lead to more successful non-marital long-term partnerships, which may lead to a greater probability of marriage or remarriage (assuming marriage is an indicator of success in a relationship). Moreover, never-married mothers of sons may have a greater demand for, or face a greater supply of, husbands (other than the biological father). They may have a greater demand for a husband than mothers of daughters because they perceive a larger cost to the absence of a father-figure for their child – that is, if men are more effective parents of sons, or if fathers are (or are at least believed to be) more crucial to the development of a son than to that of a daughter. The supply of husbands to a never-married mother of a son may be greater if potential husbands prefer to assume parenthood of sons over daughters. The same theory may be applied for divorced (or widowed) mothers of sons. They may have a greater demand for husbands than divorced mothers of daughters, if mothers of sons view the absence of a stepfather as more costly than do mothers of daughters. They may also have a greater supply of potential husbands, if men prefer step-sons over step-daughters. The authors argue that these supply and demand factors should increase the probability of marriage following a non-marital birth, and that of remarriage following a divorce.

To estimate the effect of child gender on a mother's transition to a first marriage after premarital birth, Lundberg and Rose (2003) use data from the Panel Study of Income Dynamics (PSID). They find that after a premarital birth, mothers are more likely to marry if they have a son, and that the transition into marriage is a lot faster for mothers of boys; mothers of sons are 35-40 percent more likely to marry at every age of the child. An interesting pattern emerges when marriages are separated into two categories, namely marriages to biological fathers and marriages to men other than biological fathers. Mothers of boys are 60 percent more likely to marry their child's biological father than mothers of girls, while there is only a small effect of child gender on mothers' probability of marrying someone other than the biological father. The authors find no significant effect of child gender on the probability of remarriage, or on the duration from the end of one marriage to the start of a new one.

Lundberg and Rose (2003) explain why the gender of the first-born child affects the likelihood of marriage after premarital birth, but not the probability of remarriage for a divorced/widowed mother. Their data show that the effect of child gender on marriage outcomes is strongly associated with biological fatherhood – that is, the effect of child gender on marriage probability following a non-marital birth is largest when the outcome is marriage to the child's biological father. On the supply side, a man's preference for a biological son (rather than daughter) seems more important than his preference for stepson (rather than stepdaughter). On the demand side, a mother's desire to give her son (rather than daughter) a biological father may be stronger than her desire to give her son (rather than daughter) a stepfather. A child's biological father is likely to be part of the "pool" of potential husbands after a premarital birth, and unlikely to be part of it once the mother is divorced or widowed.

Lundberg, McLanahan and Rose (2007) examine the effect of child gender on paternal involvement and parents' living arrangements. While previous studies have shown that fathers are more involved with sons than daughters, this paper looks at how this differential varies by the marital status of the father at the time of birth of their child. The authors explain that a comparison between married and unmarried fathers can shed light on the reasons for fathers' increased involvement with sons. Firstly, if fathers are more involved with sons because they prefer sons over daughters, then there should be no difference in the effect of child gender on paternal involvement between fathers who were married and those who were unmarried at the time of the child's birth. However, fathers could be more involved with sons because they are (or are expected to be) more productive parents of sons than of daughters, perhaps because of gender-specific parenting skills, or because of gender differences in children's developmental requirements. If this is so, then parents view time spent with a father as more valuable (in terms of child well-being) in the long-run to a son rather than to a daughter. According to the authors, this means that the relative value of a father's involvement with a son, compared to a daughter, increases with the number of years he lives in the child's household, and with the degree of control he has over child-rearing.

Therefore, this second explanation implies the child gender bias should be greater for fathers who were married (i.e. initially more committed to their child's mother) at the time of birth, compared to those who were unmarried.

Lundberg, McLanahan and Rose (2007) use data from the American Fragile Families and Child Wellbeing (FFCW) survey. The survey was conducted for a nationally-representative sample of births from 1998-2000. Mothers were interviewed at the time of their child's birth, and again when the child was 1, 3, and 5 years old. They find that sons born to unmarried parents are more likely than their female counterparts to receive their father's last name, and that this effect is larger when the analysis is restricted to first-born children. Again for children born to unmarried parents, there seems to be no effect of child gender on parents' living arrangements or on the time and financial investments of fathers, one year after the birth of the child. However, fathers who were married at the time of their child's birth are more likely to be living with a son rather than a daughter one year after birth.

Katzev, Warner and Acock (1994) study the interaction between child gender composition, paternal involvement, and mothers' satisfaction with married life. Following studies which have documented an effect of child gender on the probability that a divorce has already occurred, this paper estimates the effect of child gender on married mothers' perception of the likelihood that divorce will occur in the future. They hypothesize that child gender affects mothers' perception of marital stability because it affects fathers' involvement in family life which, in turn, affects mothers' perceived advantage or disadvantage within the marriage.

Katzev, Warner and Acock (1994) use data for a sample of mothers from the National Survey of Families and Households. Mothers were asked to rate the likelihood of their marriage ending in divorce using a 5-point scale. The authors constructed a measure of marital satisfaction using the mothers' reports of perceived disadvantage regarding household chores, paid work, money spending, and child care. A variable for paternal interaction time was constructed using respondents' husbands' own reports of their time spent with their children in leisure activities, having talks, or helping with homework. Their results indicate that having at least one boy, compared to having all girls, decreases a mother's estimation of her marriage ending in divorce. Having at least one boy, compared to all girls, increases the father's time spent interacting with children. Fathers' time spent with children is negatively related to the mothers' perceived disadvantage in the marriage, and the mothers' perceived disadvantage is positively related to her estimation of the marriage ending in divorce. They conclude that mothers without sons are indeed more likely to perceive their marriage as unstable because fathers spend less time with daughters than with sons, leading mothers of daughters to perceive their marriage as less equitable, and thus less satisfying and less stable. The authors also discuss the possibility that sons lead to marital stability because parents take into account the fact that marital disruption is more detrimental to the well-being of sons compared to daughter. However, the authors note that the relative importance of each explanation remains unknown.

The studies reviewed share a common goal of providing insight into the link between child gender and family structure, although they achieve this in different ways. Morgan, Lye and Condran (1988) find that couples are more likely to divorce if they have a daughter rather than a son. Morgan and Pollard (2003) find that a couple's risk of divorce is positively related to their number of daughters, but that this effect only exists for earlier marriage cohorts. Diekmann and Schmidheiny (2004) compare the effect of child gender on the probability of divorce across 18 OECD samples, and conclude that there is no systematic effect of child gender on parents' divorce probability. Mott (1994) finds that, among white families, sons are more likely to have their father in their household than daughters. Dahl and Moretti (2008) find that first-born girls are less likely than first-born boys to live with their father, and identify three channels leading to this differential – mothers are less likely to ever marry, more likely to divorce, and more likely to receive sole custody if their first-born is a girl. Lundberg and Rose (2003) find that mothers of first-born daughters are less likely to marry following a non-marital birth, and that the child gender effect is largest when the outcome is marriage to the biological father. Katzev, Warner and Acock (1994) find that mothers without sons are more likely to perceive their marriage as unstable.

Authors also explain the effect of child gender on family structure in different ways. Morgan, Lye and Condran (1988) explain that sons decrease a couple's risk of divorce because fathers are more involved with sons than with daughters. However, in their cross-country study, Diekmann and Schmidheiny (2004) find no evidence to support the father involvement hypothesis of Morgan, Lye and Condran (1988). Dahl and Moretti (2008) argue that fathers have a preference for sons over daughters, and support this hypothesis by showing that first-born daughters lead parents to have more children. Lundberg and Rose (2003) explain that child gender affects the probability of marriage following a non-marital birth because fathers may prefer to assume parenthood of a son over a daughter, or because parents may view a father's absence as more costly to a son than to a daughter. Katzev, Warner and Acock (1994) explain that mothers of sons are more likely to perceive their marriage as unstable because fathers spend less time with daughters than with sons, leading mothers of daughters to perceive their marriage as less equitable, and thus less satisfying and less stable.

3. DATA

The General Social Survey (GSS) is a cross-sectional and nationally representative survey conducted by Statistics Canada on a yearly basis. Its purpose is "to gather data on social trends in order to monitor changes in the living conditions and well-being of Canadians over time; and to provide immediate information on specific social policy issues of current or emerging interest" (Statistics Canada). Each year of the GSS covers a specific focus, along with the usual demographic variables. The fifteenth cycle of the GSS, conducted from February through December 2001, is entitled 'Family History' and is the third GSS cycle to provide detailed information on families in Canada. This cycle focuses on respondents' family origins, marriages and common-law unions, fertility, and values and attitudes. The survey is conducted at the individual level and covers individuals aged 15 and over living in the ten provinces of Canada. The survey is voluntary and the sample for the 2001 GSS includes 24,310 observations.

The population of interest consists of respondents aged 18 and over who have borne or fathered at least one child. Respondents who did not report their first-born child's gender, age or household status were excluded from the sample, as well as those who did not report their education level. Respondents were only included in the sample if their oldest reported child was their biological child; otherwise, respondents were included if any children older than their oldest biological child were step-children or adopted children who joined the household after the birth of the oldest biological child. In total, 11,138 observations were dropped, representing a loss of approximately 50 percent of the GSS sample for 2001. Seventy-five percent of these omitted observations were respondents who had never borne or fathered any children, while respondents who did not report their first-born child's gender, age or household status accounted for 15 percent of the dropped observations. There are 13,172 observations in this restricted sample. Additional sample restrictions are made for each outcome studied and are discussed in Section 4.

The weighted summary statistics of the restricted sample are presented in Table 1. Fifty-five percent of respondents are women. Forty-nine percent of respondents have a first-born daughters, and 20.5 percent of respondents were unmarried at the time of their first-born's birth. Approximately 3 percent of respondents had their first child in the 1940s, 10 percent in the 1950s, 15 percent in the 1960s, 20 percent in the 1970s, 25 percent in the 1980s, and 24 percent in the 1990s. The average age of respondents is 48 years, while the average age of respondents' first-born children is 22 years. The sample is composed 19 percent of university graduates and 27 percent of college graduates. Respondents with some postsecondary education, with high school diplomas and with some secondary school education represent

10 percent, 20 percent and 23 percent of the sample, respectively. Five percent of respondents are widowed. Table 2 provides the same information as Table 1, but separates variables by gender.

4. METHODOLODY & RESULTS

The econometric model takes the form of a linear probability function:

$$\text{Outcome}_i = \beta_0 + \beta_1(\text{FBC is a girl})_i + X_i\gamma + \varepsilon \tag{1}$$

where the explanatory variable of interest $(\text{FBC is a girl})_i$ is a binary variable equal to one if respondent *i*'s first-born biological child is a girl. X_i is a vector of characteristics, including respondent *i*'s age and quadratic age, the age of respondent *i*'s first-born child, dummies for respondent *i*'s level of education, dummies for respondent *i*'s province of residence, and dummies for the decade of birth of respondent *i*'s first-born child. The dependent variable, denoted $\text{Outcome}_{i,}$ is a binary variable pertaining to respondent *i*'s family structure – seven specifications are regressed in order to measure the effect of a first-born child's gender on seven different outcomes. Firstly, this paper estimates the impact of child gender on the probability of a respondent's first-born child residing in a biological father's household. Secondly, this paper investigates a respondent's probability of being a lone parent. The third outcome studied is a respondent's likelihood of experiencing a shotgun marriage. Next, the paper looks at a respondent's probability of ever marrying following a non-marital birth. The fifth outcome is a respondent's likelihood of being divorced at the time of the survey. This is followed by an investigation of a respondent's likelihood of having his or her first marriage end in divorce or separation. Lastly, this paper estimates the impact of first-born child gender on a respondent's probability of remarrying after a divorce. All regressions are weighted, and robust standard errors are used. The remainder of this section includes, for each outcome, a description of the binary dependent variable and sample restrictions used, as well as a discussion of regression results.

1. Probability of a respondent's first-born child residing in his or her biological father's household

In order to estimate the effect of a first-born child's gender on his or her likelihood of residing in a biological father's household, the sample is further restricted to parents aged 18 and over whose first-born child lives with at least one biological parent on a full-time or part-time basis. Widows are excluded. There are 9,156 observations in this restricted sample, consisting of 5,538 female respondents and 3,692 male respondents. The dependent variable used is a binary variable equal to one if the respondent's first-born child lives with his or her biological father on a full-time or part-time basis, and zero otherwise. More specifically, for male respondents, this variable is equal to one if the first-born child lives in the respondent's household. For female respondents, this variable is equal to one if the respondent's household includes the first-born child as well as the first-born child's biological father, if the first-born child lives outside of the respondent's household with their other biological parent, or if the first-born child lives in the respondent's household part-time and lives with their biological father the rest of the time. The coefficient on $(\text{FBC is a girl})_i$ is expected to be negative because Dahl and Moretti (2008) estimate that households with first-born daughters were found to be 3.1 percent less likely to include a resident father than those with first-born sons. Moreover, Mott (1994) finds that sons are more likely to reside with their biological father than are daughters.

Column 1 of Table 3 shows that the proportion of parents whose first-born child lives with a biological father is 1.3 percentage points (or 1.7 percent) lower among parents whose first-born child is a girl. This estimate has the expected negative sign. Although not statistically significant, this estimate is essentially unchanged when control variables are added in Column 2 and Column 3, indicating that the result is robust. The estimate in Column 3 implies that first-born daughters are 1.2 percentage points (or 1.6 percent) less likely than their male counterparts to live in their biological father's household when individual characteristics are held constant.

When the sample is restricted to female respondents, the estimate on $(\text{FB is a girl})_i$ is reduced by half, compared to the estimate provided by the sample of both mothers and fathers (Table 3, Columns 4 to 6). The estimate in Column 4 of Table 3 implies that mothers' first-born children are 0.6 percentage points (or 0.7 percent) less likely to live with a biological father if they are girls. This estimated effect is not statistically significant, nor important in size; however, the estimate is robust, as it is unchanged when control variables are added in Column 5 and Column 6.

As seen in Column 7 of Table 3, among male respondents, the proportion of fathers whose first-born lives in their household is 2.2 percentage points (or 3.2 percent) lower when the first-born is a girl, but this difference is not statistically significant. However, when control variables are included in Column 8 and Column 9, the coefficient on $(\text{FBC is a girl})_i$ becomes significant at the 5 percent level. The estimate in Column 9 indicates that fathers of first-born daughters are 2.5 percentage points less likely to have their first-born in their household than fathers of sons, when individual characteristics are held constant. Since 68 percent of male respondents live with their first-born child, this means that male respondents are 3.5

percent less likely to live with their first-born if their first-born is a daughter rather than a son. Because this estimate is not statistically significant when the specification does not include variables which control for any paternal or child characteristics, this means that the effect of child gender on a father's probability of residing with his first-born child is conditional on factors such as the child's age.

The results shown in Table 3 indicate that first-born girls are slightly less likely to live with their biological father than first-born boys. Although a statistically significant estimate is only found when the sample is restricted to men and the specification includes control variables (Table 3, Columns 8 and 9), the estimates are robust. That is, for each of the three samples used (i.e. the sample of both male and female parents, the sample of mothers, and the sample of fathers), the estimated coefficient on $(\text{FBC is a girl})_i$ is unchanged when control variables are included in the specification. These results therefore support the findings of Dahl and Moretti (2008) and Mott (1994).

2. Respondent's probability of being a lone parent

A lone-parent is defined as a parent whose household includes at least one biological child (on a full-time or part-time basis), and does not include their child(ren)'s other biological parent, or a step-parent. For this outcome, the sample includes the 11,760 respondents who live with at least one biological child, regardless of whether the first-born child lives at home. Again, widowed respondents are excluded. Within this sample, there are 6,855 female respondents and 4,905 male respondents. Lone parents account for 6.7 percent of this sample of parents; 10 percent of the female respondents are lone parents, while only 2.5 percent of the males are lone parents. The dependent variable is a binary variable equal to one if respondent i is a lone parent, and zero otherwise. The coefficient β_1 is expected to be positive, because parents of first-born daughters have been shown to be less likely to marry following a non-marital birth (Lundberg and Rose, 2003) and more likely to divorce (Dahl and Moretti, 2008; Morgan, Lye and Condran, 1988) than parents of first-born sons. Additionally, it is expected that the effect of child gender on the likelihood of being a lone parent be larger among mothers than among fathers. It has been shown that sons are more likely than daughters to live with their biological father (Dahl and Moretti, 2008; Mott, 1994), indicating that sons may be less likely than daughters to have a mother who is a lone-parent; however, there is no evidence that gender affects a child's likelihood of having a lone-parent father.

Within a sample of both male and female respondents, the proportion of lone-parents is 0.2 percentage points (or 3 percent) higher among parents of girls (Table 4, Column 1). When individual characteristics are held constant, parents with a first-born girl are 0.3 percentage points (or 4.5 percent) more likely to be lone parents than those with a first-born boy, (Table 4, Columns 2 and 3). However, these estimates are not statistically significant at any standard level of confidence.

When the sample is restricted to female respondents, the proportion of lone parents is 0.4 percentage points (or 4 percent) higher among mothers of first-born daughters (Table 4, Column 4). Although the estimate is not statistically significant, it barely changes when control variables are added in Columns 5 and 6. The impact of a first-born child's gender on a mother's likelihood of being a lone parent is estimated at 0.5 percentage points (or 5 percent) and remains statistically insignificant in Columns 5 and 6. Restricting the sample to male respondents provides a similar statistically insignificant result; regardless of whether control variables are included in the specification, fathers of first-born girls are 0.1 percentage points (or 4 percent) more likely to be lone parents, compared to fathers of first-born boys (Table 4, Columns 7 to 9).

The estimates shown in Table 4, although not statistically significant, have the expected sign. For each of the three samples, the estimated coefficient on $(\text{FBC is a girl})_i$ is robust with the inclusion of control variables. The results shown in Table 4 are therefore consistent with the hypothesis that parents of first-born daughters are more likely to be lone parents than those of first-born sons.

3. Respondent's probability of experiencing a shotgun marriage

A shotgun marriage is commonly defined as a marriage that is contracted during an unplanned pregnancy in order to avoid a non-marital birth. The effect of child gender on a respondent's probability of experiencing a shotgun marriage is estimated by restricting the sample to respondents who were single (never married) four months before the birth of their first child. The four-month time period is used because the gender of a child *in utero* can be accurately determined approximately twenty weeks into a pregnancy (Schwärzler et al, 1999). Therefore, parents who married more than four months prior to the birth of their child would not have been influenced by the gender of their child in their decision to marry. However, it cannot be shown how many of these parents were aware of the gender of their children four months before delivery – some parents may become aware of their child's gender later into the pregnancy, while some remain unaware until delivery. The dependent variable used to estimate the effect of first-born child gender on the likelihood of a shotgun marriage is a binary variable equal to one if respondent *i* was married at the time of birth of the first born child, and zero otherwise. The sample mean of this variable indicates that 21 percent of respondents who were single four months before the birth of their first-born child were married at the time of birth of their first-born child.

The estimated coefficient for β_1 is expected to be negative because of the finding that first-time mothers who presumably know the gender of their child before delivery are more likely to be married at delivery if they give birth to a boy rather than a girl (Dahl and Moretti, 2008). However, for a sample of both male and female respondents, the estimate is slightly positive and is not statistically significant at any standard confidence level (Table 5, Columns 1 to 3). Within a sub-sample of female respondents, the coefficient is positive and is not statistically significant (Table 5, Columns 4 to 6). The coefficient has the expected negative sign when the sample is restricted to male respondents and control variables are used; however, these estimates are not statistically significant (Table 5, Columns 8 and 9). These findings do not support those of Dahl and Moretti (2008).

4. Respondent's probability of ever marrying following a non-marital birth

To estimate the effect of child gender on a respondent's probability of ever marrying following a non-marital birth, the sample consists of respondents aged 18 and over who were single (never married) at the time of birth of their first child. This sample includes a total of 2,963 respondents, 1,907 of which are women, and 1,056 of which are men.

The effect of child gender on marriage probability following a non-marital birth is expected to vary significantly between a single mother and a single father. In theory, child gender affects a couple's likelihood of marriage following a non-marital birth because a son leads a couple to place a higher value on marriage relative to single parenthood; this may be because an unmarried couple who gives birth to a son rather than a daughter will attribute a higher cost to the absence of the child's father, assuming men are (or are at least believed to be) more effective parents of sons, or more crucial to the social development of a son than to that of a daughter. Similarly, child gender is expected to affect a single mother's likelihood of marrying someone other than the biological father following a non-marital birth; single mothers of sons may have a greater demand for a husband than single mothers of daughters because they perceive a larger cost to the absence of a father-figure for their child, and they may face a larger supply of husbands if potential husbands prefer to assume parenthood of step-sons over step-daughters. However, according to this theory, child gender is not expected to affect a never-married father's probability of marrying someone other than the biological mother.

The dependent variable is a binary variable equal to one if respondent *i* ever married following their first-born child's non-marital birth, and zero otherwise. The estimated coefficient for β_1 is expected to be negative because of the finding that mothers are more likely to marry following a non-marital birth if they give birth to a boy (Lundberg and Rose, 2003). Moreover, Dahl and Moretti (2008) find that mothers with first-born daughters are less likely to ever marry than those with first-born sons.

Column 1 of Table 6 shows that, using the sample of both male and female respondents who had a non-marital first birth, the proportion of parents who ever married is 2.4 percentage points (or 4.2 percent) higher among parents of first-born girls, compared to parents of first-born boys. This result is not statistically significant at any standard level of confidence. When control variables are included in Column 2 and Column 3, the estimated coefficient on $(\text{FBC is a girl})_i$ becomes slightly negative, and remains statistically insignificant; the effect of child gender on a parent's probability of marrying following a non-marital birth is estimated at -0.1 percentage points.

It can be seen in Column 4 of Table 6 that, among female respondents who were never married at the time of their first child's birth, the proportion of mothers who married following their first child's birth is 1.2 percentage points (or 2.2 percent) lower among those whose first-born is a girl. This estimate has the expected negative sign, but is not significant at any standard confidence level. When control variables are added in Column 5 and Column 6, the estimated coefficient on $(\text{FBC is a girl})_i$ more than doubles in size, but remains statistically insignificant. The estimate indicates that mothers of first-born daughters are 3.1 percentage points (or 5.7 percent) less likely to ever marry following a non-marital first birth, compared to mothers of first-born sons. These results do not support the findings of Lundberg and Rose (2003) or that of Dahl and Moretti (2008), which suggest that mothers of daughters are less likely to marry following a non-marital birth than those of sons.

For male respondents, a statistically significant estimate is found when control variables are excluded (Table 6, Column 7). However, this estimate does not have the expected negative sign, and it implies that the proportion of fathers who have married following a non-marital birth is 7.3 percentage points (or 18.2 percent) higher among fathers of first-born girls, compared to those of first-born boys. When control variables are included in Column 8 and in Column 9, the estimated coefficient on $(\text{FBC is a girl})_i$ is reduced by half and is no longer statistically significant – this implies that the estimate displayed in Column 7 was biased due to one or more omitted variable(s). This lack of a robust estimate with a negative sign for β_1 is expected, since child gender is not expected to affect fathers' probability of marrying anyone other than the biological mother after a non-marital birth.

5. Respondent's probability of being divorced or separated at the time of the survey

The sample consists of respondents aged 18 and over who have borne or fathered at least one child, and who have been married at least once. Widows are excluded. The dependent variable is a binary variable equal to one if respondent *i* was divorced or separated at the time of the survey, and zero otherwise. The estimated coefficient for β_1 is expected to be positive because of the finding that parents of first-born daughters are more likely to be divorced than parents of first-born sons (Dahl and Moretti, 2008).

Using a sample of 10,801 observations, including both female and male respondents, the proportion of parents who are separated or divorced is 1.1 percentage points (or 7.5 percent) higher among parents of first-born girls (Table 7, Column 1). When control variables are included in Columns 2 and 3, the estimated coefficient on $(\text{FBC is a girl})_i$ implies that respondents with a first-born girl are 1 percentage point (or 6.8 percent) more likely to be divorced or separated at the time of the survey. The estimates in Columns 1 to 3 are not statistically significant at any standard confidence level. When the sample is restricted to the 6,209 ever-married female respondents who have borne at least one child, the estimated coefficient on the variable $(\text{FBC is a girl})_i$ is only slightly positive, and is not statistically significant (Table 7, Columns 4 to 6).

Among a sample of 4,529 ever-married male respondents who have fathered at least one child, fathers of first-born girls are more likely to be separated or divorced. In Column 7 of Table 7, the coefficient on $(\text{FBC is a girl})_i$ has the expected sign, and is significant at the 5 percent level when no control variables are used. This estimate indicates that the proportion of fathers who are separated or divorced at the time of the survey is 2.1 percentage points (or 17.6 percent) higher for fathers of first-born girls, compared to fathers of first-born boys. When control variables are added in Column 8 and Column 9, the estimated coefficient on $(\text{FBC is a girl})_i$ is virtually unchanged and remains significant at the 5 percent level. The estimate in Column 9 implies that fathers of first-born daughters are 2.2 percentage points (or 18.5 percent) more likely than fathers of boys to be separated or divorced at the time of the survey, holding constant individual characteristics such as education, age, and birth decade.

It is not clear why there is an important and statistically significant effect of child gender on a male respondent's probability of being separated or divorced at the time of the survey, but no such effect for female respondents. The effect of child gender on a respondent's likelihood of experiencing marital dissolution is not expected to vary between mothers and fathers. One explanation is that the effect of child gender on the likelihood of remarrying is different between mothers and fathers. Child gender may equally affect both mothers' and fathers' probability of divorce or separation, but it may also affect mothers' probability of remarrying after a divorce or separation. If mothers of daughters are somehow more likely to remarry than mothers of sons, they become less likely to be separated or divorced at the time of the survey. In other words, even if mothers of daughters are more likely to divorce than mothers of sons, this likelihood would not be reflected in their marital status at the time of the survey if they also have a higher likelihood of remarriage.

Although a gender differential in the likelihood of remarriage could explain the results of this analysis, Lundberg and Rose (2003) find no effect of child gender on mothers' probability of remarriage after a divorce. Moreover, these authors hypothesize that mothers of first-born daughters should be *less* likely to remarry than mothers of first-born sons, because mothers of sons have a greater demand for, and face a greater supply of, potential stepfathers for their child. Alternatively, child gender may have no significant effect on the likelihood of marital dissolution of mothers or fathers, but may have a significant effect on fathers' likelihood of remarriage. That is, if divorced fathers of first-born daughters are less likely to remarry than those of first-born sons, then ever-married fathers of first-born daughters would be more likely to be divorced at the time of the survey than those of first-born sons.

If the difference between mothers and fathers in the effect of child gender on parents' likelihood of being divorced is due entirely to a the difference between mothers and fathers in the effect of child gender on divorced parents' likelihood of remarriage, then the effect of child gender on respondents' likelihood of a *first* marriage ending in divorce should not differ between male and female respondents. The following outcome perhaps more accurately captures a parent's likelihood of experiencing marital dissolution.

6. Respondent's probability of a first marriage ending in divorce or separation

Dahl and Moretti (2008) note that a first-born child's gender may affect parents' likelihood of being divorced at the time of the survey by affecting parents' likelihood of remarrying following a divorce. Therefore, in order to more accurately measure the impact of first-born child gender on a parent's likelihood of marital dissolution, Dahl and Moretti (2008) estimate the impact of first-born child gender on parents' likelihood of a *first* marriage ending in divorce, and find that parents of first-born girls are more likely to end their first marriage than those of first-born boys.

The sample is restricted to respondents whose first marriages were intact at the time of birth of their first child. This sample includes 3,949 men and 5,854 women, totalling 9,803 respondents. The dependent variable is a binary variable equal to one if respondent *i*'s first marriage ended in divorce or separation, and zero otherwise. The estimated coefficient for β_1 is expected to be positive because, among respondents who had their first child during their first marriage, parents of first-born daughters should be more likely to divorce or separate than those of first-born sons.

Column 1 of Table 8 displays the regression results for a sample of both male and female respondents, when control variables are excluded. The proportion of parents whose first marriage ended in separation or divorce is 1.4 percentage points (or 9.5 percent) higher among parents of first-born daughters compared to those of first-born sons, and this estimate is statistically significant at the 10 percent level. When the full set of control variables is included in Column 3, the estimate is unchanged and remains statistically significant at the 10 percent level.

For a sample of female respondents, the estimated coefficient for (FBC is a girl)$_i$ is positive, but statistically insignificant in each specification. Column 4 of Table 8 shows that the proportion of mothers whose first marriage ended is 1 percentage point (or 6.2 percent) higher among mothers of first-born daughters. When the full set of control variables is included in Column 6, the estimate implies that mothers of first-born daughters are 0.9 percentage points (or 5.6 percent) more likely than mothers of first-born sons to have their first marriage end.

When the sample is restricted to male respondents, the estimate shown in Column 7 of Table 8 indicates that the proportion of fathers whose first marriage ended is 1.9 percentage points (or 14.6 percent) higher among those who have a first-born girl instead of a first-born boy. This estimate is significant at the 10 percent level. When the full set of control variables is used in Column 9, the estimated coefficient on (FBC is girl)$_i$ slightly increases, and remains statistically significant at the 10 percent level. Among male respondents, those with a first-born girl are 2 percentage points (or 15.4 percent) more likely to have their first marriage end than those with a first-born son, when fathers' individual characteristics are held constant.

Overall, the analysis of the effect of child gender on this particular outcome supports the finding of Dahl and Moretti (2008) – parents of first-born girls are more likely than parents of first-born boys to have their first marriage end in divorce or separation. For each sample in Table 8, the estimated coefficient on (FBC is a girl)$_i$ has the expected positive sign and is robust with the inclusion of control variables. Although the sample of all parents provides a statistically significant estimate for β_1, the sample of women alone does not provide a statistically significant estimate. When the sample is restricted to male respondents, the estimate is statistically significant, and twice as large as that provided by the sample of female respondents. It is not clear why such a pattern exists; because a divorce or separation is experienced by both members of a couple, the effect of child gender on the probability of a first marriage ending is not expected to differ between a sample of men and a sample of women.

7. Respondent's probability of remarriage following a divorce

The sample is restricted to respondents aged 18 years and over whose first marriages ended in divorce, and whose first-born children were born during their first marriages. The dependent variable is a binary variable equal to one if individual *i* has been married more than once, and zero otherwise. This sample includes 1,741 observations, consisting of 1,092 females and 649 males. Within this sample of respondents, 49 percent of the males had remarried, and 43 percent of females had remarried.

The effect of first-born child gender is expected to differ between divorced mothers and divorced fathers, just as the effect of first-born child gender on marriage probabilities was expected to vary between single mothers and single fathers. In theory, a divorced mother's likelihood of remarrying is a function of her demand for a step-father for her child, and her supply of potential step-fathers. If divorced mothers of sons perceive the lack of a step-father more detrimental to their child's social development than mothers of daughters, or if potential step-fathers would prefer a step-son over a step-daughter, then divorced mothers have both a higher demand and a higher supply of step-fathers if their first-born child is a boy rather than a girl (Lundberg and Rose, 2003). This same theory does not explain any possible effect of first-born child gender on divorced fathers' likelihood of remarriage; therefore, first-born child gender is expected to have an effect on remarriage for female respondents only. Although Lundberg and Rose (2003) find no effect of first-born child gender on mothers' probability or remarriage following a divorce, the estimated coefficient on $(\text{FBC is a girl})_i$ is expected to be negative because mothers of first-born girls are expected to be less likely to remarry than those of first-born boys.

Using the sample of both male and female respondents, the estimated coefficient on $(\text{FBC is a girl})_i$ does not have the expected sign, and is not statistically significant at any standard level of confidence (Table 9, Columns 1 to 3). The estimate in Column 1 of Table 9 implies that the proportion of parents who remarried after their first marriage ended is 1.8 percentage points (or 4 percent) higher among parents whose first-born child is a girl. When control variables are added in Column 2 and Column 3, the estimated coefficient slightly decreases. The estimate in Column 3 implies that parents whose first marriages ended are 1.3 percentage points (or 2.9 percent) more likely to ever remarry if their first-born child is a girl rather than a boy (Table 9, Column 3).

The sign of the coefficient for (FBC is a girl)$_i$ differs between the male and female subgroups, and the estimates remain statistically insignificant in both the male and female samples. When the sample is restricted to female respondents, the estimated coefficients are greater than zero. The proportion of mothers who remarried is 6 percentage points (or 14.1 percent) higher among mothers of first-born girls compared to mothers of first-born boys (Table 9, Column 4). When individual characteristics are held constant, the estimated coefficient indicates that divorced mothers are 4.8 percentage points (or 11.3 percent) more likely to remarry if their first-born child is a girl (Table 9, Column 6). Within the group of male respondents, having a (FBC is a girl)$_i$ has a negative effect on fathers' probability of remarriage following a divorce. The estimate in Column 7 indicates that the proportion of fathers who remarried is 4.6 percentage points (or 9.3 percent) lower for fathers whose first-born child is a girl rather than a boy. When the full set of control variables is included in the specification, the estimate implies that fathers of first-born girls are 5.3 percentage points (or 10.8 percent) less likely to remarry after a divorce from a first marriage (Table 9, Column 9).

Although not statistically significant, the results shown in Table 9 are consistent with the suggested explanation for the mother-father differential in the effect of child gender on the likelihood of being separated or divorced at the time of the survey. The results in Table 7 indicate that having a first-born girl has a statistically significant effect on a respondent's probability of being separated or divorced at the time of the survey within a sample of male respondents, but has no such effect within a sample of female respondents – this suggests that mothers of daughters may be more likely to remarry than mothers of sons, or that fathers of daughters may be less likely to remarry than fathers of sons.

5. CONCLUSION

This paper examines the link between child gender and family structure using Canadian data from the 2001 General Social Survey (GSS). Using a linear probability model, this paper estimates the effect of a first-born child's gender on his or her probability of living with a biological father, and on his or her parent's probability of being a lone parent, experiencing a shotgun marriage, marrying following a non-marital birth, being divorced or separated, ending a first marriage, and remarrying following a divorce.

This paper finds that first-born daughters are 1.6 percent less likely to live with their biological father; although this result was not statistically significant, it was found to be robust with the inclusion of control variables. The only statistically significant effect of child gender on this outcome was estimated when the sample was restricted to male respondents; when control variables were included to account for individual characteristics, fathers of first-born daughters were found to be 3.5 percent less likely to live with their first-born, compared to fathers of first-born sons. In addition, although statistically insignificant, a robust estimate indicated that parents of first-born daughters are 4.5 percent more likely to be lone parents than parents of first-born sons.

A sample of ever-married parents was used to study the effect of first-born child gender on parents' probability of being separated or divorced at the time of the survey; it was found that parents are 6.8 percent more likely to be separated or divorced when their first-born is a girl. Among ever-married fathers, those with a first-born girl were 18.5 percent more likely to be separated or divorced at the time of the survey; this result was both robust and statistically significant. However, child gender had no effect on ever-married mothers' probability of being separated or divorced at the time of the survey. Using a sample of parents who had their first-born child during their first marriage, it was estimated that parents of first-born girls are 9.5 percent more likely than parents of first-born boys to have their first marriage end in divorce; this estimate was robust and significant. When the sample was restricted to male respondents, it was found that fathers of first-born girls are 15.4 percent more likely to have their first marriage end, and this estimate remained significant. The sample of mothers provided a smaller and statistically insignificant estimate – mothers of first-born girls were 5.6 percent more likely than those of first-born boys to have their first marriage end. It is not clear why the effect of child gender on a respondent's likelihood of being separated or divorced, and of having a first marriage end, differs between male and female respondents.

This study found no effect of child gender on a parent's probability of experiencing a shotgun marriage. Moreover, using a sample of parents who were unmarried at the time of their first-born child's birth, no evidence was found to support the hypothesis that child gender affects parents' probability of ever marrying. Lastly, although estimates were not statistically significant, having a first-born daughter had a positive effect on divorced mothers' likelihood of remarrying, and a negative effect on divorced fathers' likelihood of remarrying.

This study provides some support for the findings of Dahl and Moretti (2008). The most notable finding is that the gender of a first-born child affects parents' likelihood of experiencing separation or divorce. It was estimated that parents of first-born girls are 6.8 percent more likely to be separated or divorced and 9.5 percent more likely to have their first marriage end in separation or divorce, compared to parents of first-born boys; however, the effect of child gender on these outcomes differs between mothers and fathers. Also consistent with Dahl and Moretti (2008) are the finding that first-born girls are less likely to live with a biological father, and the finding that parents of first-born girls are more likely to be lone parents. However, because the child gender effect is found to be small and statistically insignificant in most cases, a larger sample may be required to provide more statistically significant estimates of the effect of child gender on family structure.

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Table 1: Summary Statistics, Male and Female Respondents

Variable	**Mean (Std. Dev.)**
Woman	0.55 (0.50)
First-born child is a girl	0.49 (0.50)
Age	48.26 (13.60)
Age of first-born	22.17 (14.38)
Unmarried at birth of first-born child	0.21 (0.40)
Separated or divorced	0.13 (0.33)
Never married	0.09 (0.28)
Widowed	0.05 (0.22)
University degree	0.19 (0.39)
College diploma	0.27 (0.45)
Some post-secondary	0.10 (0.30)
High school diploma	0.20 (0.40)
High school dropout	0.23 (0.42)
First-born child born in 1940s	0.03 (0.16)
First-born child born in 1950s	0.10 (0.30)
First-born child born in 1960s	0.15 (0.35)
First-born child born in 1970s	0.20 (0.40)
First-born child born in 1980s	0.25 (0.43)
First-born child born in 1990s	0.24 (0.43)
N	13,172

Note: All results are weighted. Standard deviations are in parentheses.

Table 2: Summary Statistics, by Gender of Respondent

Variable	Female Respondents	Male Respondents
First-born child is a girl	0.49 (0.50)	0.48 (0.50)
Age	47.83 (13.95)	48.79 (13.15)
Age of first-born	23.06 (14.75)	21.09 (13.84)
Unmarried at birth of first-born child	0.22 (0.41)	0.19 (0.39)
Separated or divorced	0.14 (0.35)	0.11 (0.31)
Never married	0.10 (0.30)	0.08 (0.27)
Widowed	0.08 (0.26)	0.02 (0.14)
University degree	0.17 (0.37)	0.22 (0.42)
College diploma	0.28 (0.45)	0.27 (0.44)
Some post-secondary	0.10 (0.30)	0.10 (0.30)
High school diploma	0.21 (0.41)	0.18 (0.39)
High school dropout	0.24 (0.43)	0.23 (0.42)
First-born child born in 1940s	0.03 (0.18)	0.02 (0.12)
First-born child born in 1950s	0.11 (0.31)	0.08 (0.28)
First-born child born in 1960s	0.15 (0.36)	0.15 (0.35)
First-born child born in 1970s	0.19 (0.40)	0.20 (0.40)
First-born child born in 1980s	0.25 (0.43)	0.24 (0.43)
First-born child born in 1990s	0.22 (0.42)	0.26 (0.44)
N	7,996	5,176

Note: All results are weighted. Standard deviations are in parentheses.

Table 3: Regression Results, Probability of a respondent's first-born child residing in his or her biological father's household

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Variable	**Men and Women**			**Women**			**Men**		
First-born is a girl	-0.013 (0.010)	-0.011 (0.009)	-0.012 (0.009)	-0.006 (0.011)	-0.006 (0.011)	-0.006 (0.010)	-0.022 (0.017)	-0.027** (0.012)	-0.025** (0.011)
Age		0.027*** (0.003)	0.060*** (0.004)		0.025*** (0.003)	0.049*** (0.004)		0.009*** (0.003)	-0.002 (0.004)
Age squared		0.000*** (0.000)	-0.001*** (0.000)		0.000** (0.000)	0.000*** (0.000)		0.000* (0.000)	0.000 (0.000)
Age of FBC		-0.012*** (0.001)	-0.009*** (0.002)		-0.008*** (0.001)	-0.010*** (0.002)		-0.030*** (0.001)	-0.024*** (0.002)
University degree		-0.014 (0.014)	-0.014 (0.014)		0.026 (0.016)	0.014 (0.016)		-0.021 (0.018)	-0.014 (0.016)
College diploma		-0.022 (0.013)	-0.021 (0.013)		-0.014 (0.015)	-0.020 (0.015)		-0.020 (0.018)	-0.016 (0.016)
Some post-secondary		-0.039** (0.018)	-0.035** (0.017)		-0.052** (0.021)	-0.062*** (0.020)		-0.018 (0.024)	-0.022 (0.023)
High school dropout		-0.035** (0.016)	-0.033** (0.016)		-0.020 (0.016)	-0.032** (0.016)		-0.033* (0.019)	-0.015 (0.019)
Other education		-0.036 (0.088)	-0.037 (0.087)		-0.160 (0.114)	-0.155 (0.107)		0.020 (0.076)	-0.029 (0.077)
Newfoundland		-0.002 (0.023)	-0.006 (0.022)		0.048** (0.021)	0.037* (0.020)		-0.013 (0.031)	-0.010 (0.028)
Prince-Edward Island		-0.005 (0.028)	0.000 (0.028)		-0.010 (0.031)	-0.015 (0.030)		-0.010 (0.037)	0.012 (0.036)
Nova Scotia		-0.029 (0.023)	-0.033 (0.022)		0.005 (0.021)	0.004 (0.021)		-0.049 (0.031)	-0.072*** (0.026)
New Brunswick		-0.035 (0.027)	-0.037 (0.025)		0.018 (0.024)	0.018 (0.024)		-0.074** (0.032)	-0.069*** (0.023)
Quebec		-0.019 (0.013)	-0.020 (0.012)		-0.024 (0.015)	-0.019 (0.014)		0.002 (0.016)	-0.007 (0.015)
Manitoba		-0.014 (0.021)	-0.013 (0.020)		-0.009 (0.023)	-0.010 (0.023)		-0.008 (0.027)	-0.015 (0.025)
Saskatchewan		-0.031 (0.021)	-0.021 (0.021)		0.014 (0.023)	0.015 (0.023)		-0.068*** (0.025)	-0.073*** (0.021)
Alberta		-0.048*** (0.018)	-0.048*** (0.017)		-0.006 (0.020)	-0.007 (0.020)		-0.067*** (0.022)	-0.054*** (0.020)
British Columbia		0.003 (0.015)	-0.003 (0.014)		-0.010 (0.016)	-0.013 (0.016)		-0.005 (0.019)	-0.008 (0.018)
FBC born in 1940s			0.403*** (0.109)			0.215** (0.098)			0.053 (0.122)
FBC born in 1950s			0.095 (0.081)			0.071 (0.084)			-0.052 (0.104)
FBC born in 1960s			-0.163** (0.064)			-0.003 (0.069)			-0.132 (0.082)
FBC born in 1970s			-0.255*** (0.047)			-0.114** (0.056)			-0.088 (0.064)
FBC born in 1980s			-0.133*** (0.031)			-0.243*** (0.039)			0.315*** (0.036)
FBC born in 1990s			-0.071*** (0.019)			-0.140*** (0.026)			0.165*** (0.023)
N	9,230			5,538			3,692		

Note: All regressions are weighted. Robust standard errors are in parentheses. *Significant at 10%; **significant at 5%; ***significant at 1%.

Table 4: Regression Results, Respondent's probability of being a lone parent

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Variable	**Men and Women**			**Women**			**Men**		
First-born is a girl	0.002 (0.005)	0.003 (0.005)	0.003 (0.005)	0.004 (0.008)	0.005 (0.008)	0.005 (0.008)	0.001 (0.004)	0.001 (0.005)	0.001 (0.005)
Age		-0.008*** (0.001)	-0.017*** (0.002)		-0.011*** (0.002)	-0.022*** (0.003)		0.003** (0.001)	0.000 (0.001)
Age squared		0.000** (0.000)	0.000*** (0.000)		0.000** (0.000)	0.000*** (0.000)		0.000** (0.000)	0.000 (0.000)
Age of FBC		0.003*** (0.000)	0.003*** (0.001)		0.002** (0.001)	0.002 (0.002)		0.000 (0.000)	-0.001 (0.001)
University degree		-0.009 (0.007)	-0.005 (0.007)		-0.016 (0.013)	-0.011 (0.013)		-0.005 (0.007)	-0.004 (0.007)
College diploma		-0.001 (0.007)	0.001 (0.007)		-0.002 (0.011)	0.000 (0.011)		0.001 (0.007)	0.002 (0.007)
Some post-secondary		0.018* (0.010)	0.019* (0.010)		0.025* (0.015)	0.027* (0.015)		0.012 (0.012)	0.012 (0.012)
High school dropout		0.006 (0.007)	0.009 (0.007)		0.014 (0.012)	0.018 (0.012)		0.004 (0.007)	0.006 (0.007)
Other education		0.083 (0.063)	0.079 (0.062)		0.264** (0.122)	0.266** (0.119)		-0.027*** (0.006)	-0.031*** (0.006)
Newfoundland		-0.017* (0.009)	-0.016* (0.009)		-0.028* (0.015)	-0.025* (0.015)		-0.001 (0.007)	-0.001 (0.007)
Prince-Edward Island		0.002 (0.014)	0.002 (0.014)		-0.024 (0.018)	-0.022 (0.019)		0.044** (0.022)	0.043** (0.021)
Nova Scotia		0.008 (0.010)	0.007 (0.010)		0.005 (0.016)	0.004 (0.016)		0.011 (0.010)	0.010 (0.010)
New Brunswick		-0.001 (0.012)	-0.001 (0.012)		-0.018 (0.018)	-0.017 (0.018)		0.021 (0.016)	0.021 (0.016)
Quebec		0.011* (0.006)	0.010 (0.006)		0.005 (0.010)	0.004 (0.011)		0.016** (0.007)	0.016** (0.007)
Manitoba		-0.008 (0.010)	-0.009 (0.010)		-0.013 (0.017)	-0.013 (0.017)		-0.003 (0.006)	-0.004 (0.007)
Saskatchewan		0.002 (0.011)	0.000 (0.011)		-0.011 (0.017)	-0.011 (0.017)		0.024* (0.013)	0.022* (0.013)
Alberta		0.002 (0.009)	0.003 (0.009)		0.002 (0.015)	0.002 (0.015)		0.009 (0.008)	0.009 (0.008)
British Columbia		-0.001 (0.007)	0.000 (0.007)		-0.005 (0.011)	-0.004 (0.011)		0.005 (0.006)	0.005 (0.006)
FBC born in 1940s			-0.001 (0.044)			-0.001 (0.072)			0.044 (0.044)
FBC born in 1950s			0.031 (0.038)			0.046 (0.062)			0.034 (0.038)
FBC born in 1960s			0.052* (0.031)			0.073 (0.052)			0.030 (0.031)
FBC born in 1970s			0.091*** (0.026)			0.127*** (0.043)			0.039 (0.026)
FBC born in 1980s			0.108*** (0.018)			0.144*** (0.029)			0.048*** (0.018)
FBC born in 1990s			0.064*** (0.012)			0.103*** (0.020)			0.019 (0.012)
N		11,760			6,855			4,905	

Note: All regressions are weighted. Robust standard errors are in parentheses. *Significant at 10%; **significant at 5%; ***significant at 1%.

Table 5: Regression Results, Respondent's probability of experiencing a shotgun marriage

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Variable	**Men and Women**			**Women**			**Men**		
First-born is a girl	0.019	0.009	0.005	0.028	0.019	0.013	0.005	-0.004	-0.008
	(0.016)	(0.015)	(0.015)	(0.020)	(0.019)	(0.019)	(0.026)	(0.025)	(0.024)
Age		0.021***	0.010**		0.021***	0.017***		0.021***	0.000
		(0.003)	(0.005)		(0.004)	(0.006)		(0.005)	(0.008)
Age squared		0.000***	0.000		0.000***	0.000**		0.000***	0.000
		(0.000)	(0.000)		(0.000)	(0.000)		(0.000)	(0.000)
Age of FBC		0.005***	0.004		0.007***	0.003		0.003	0.005
		(0.002)	(0.003)		(0.002)	(0.004)		(0.002)	(0.005)
University degree		-0.013	-0.012		-0.033	-0.032		0.018	0.020
		(0.030)	(0.030)		(0.038)	(0.038)		(0.048)	(0.048)
College diploma		-0.020	-0.017		-0.032	-0.032		0.002	0.012
		(0.022)	(0.022)		(0.028)	(0.028)		(0.035)	(0.035)
Some post-secondary		-0.011	-0.009		-0.015	-0.016		-0.002	0.003
		(0.028)	(0.028)		(0.033)	(0.033)		(0.047)	(0.047)
High school dropout		-0.038*	-0.037*		-0.039	-0.038		-0.035	-0.030
		(0.022)	(0.022)		(0.029)	(0.029)		(0.035)	(0.035)
Other education		-0.141***	-0.125***		-0.190***	-0.165***		-0.103**	-0.096**
		(0.033)	(0.030)		(0.050)	(0.047)		(0.044)	(0.044)
Newfoundland		0.097***	0.094**		0.100**	0.096**		0.090	0.086
		(0.037)	(0.037)		(0.043)	(0.042)		(0.063)	(0.062)
Prince-Edward Island		0.019	0.015		0.004	-0.001		0.028	0.016
		(0.045)	(0.045)		(0.053)	(0.054)		(0.074)	(0.074)
Nova Scotia		0.055	0.051		0.030	0.024		0.104	0.115
		(0.040)	(0.040)		(0.041)	(0.041)		(0.080)	(0.079)
New Brunswick		-0.032	-0.034		-0.035	-0.034		-0.029	-0.034
		(0.040)	(0.041)		(0.046)	(0.044)		(0.073)	(0.076)
Quebec		-0.099***	-0.095***		-0.073***	-0.071***		-0.134***	-0.120***
		(0.020)	(0.020)		(0.025)	(0.024)		(0.033)	(0.033)
Manitoba		-0.050	-0.048		-0.009	-0.011		-0.114**	-0.102*
		(0.034)	(0.034)		(0.044)	(0.045)		(0.052)	(0.052)
Saskatchewan		0.014	0.012		0.000	-0.002		0.034	0.037
		(0.033)	(0.033)		(0.042)	(0.042)		(0.056)	(0.055)
Alberta		0.065**	0.062*		0.089**	0.094**		0.032	0.026
		(0.032)	(0.032)		(0.042)	(0.043)		(0.049)	(0.051)
British Columbia		0.052*	0.053*		0.037	0.037		0.069	0.075
		(0.029)	(0.029)		(0.038)	(0.038)		(0.048)	(0.046)
FBC born in 1940s			-0.228			0.007			-0.492*
			(0.171)			(0.215)			(0.277)
FBC born in 1950s			-0.095			0.121			-0.368*
			(0.137)			(0.171)			(0.223)
FBC born in 1960s			0.045			0.162			-0.106
			(0.104)			(0.130)			(0.171)
FBC born in 1970s			0.025			0.037			0.015
			(0.079)			(0.099)			(0.128)
FBC born in 1980s			-0.047			-0.036			-0.065
			(0.050)			(0.061)			(0.085)
FBC born in 1990s			-0.047			-0.046			-0.050
			(0.030)			(0.037)			(0.052)
N		3,665			2,352			1,313	

Note: All regressions are weighted. Robust standard errors are in parentheses. *Significant at 10%; **significant at 5%; ***significant at 1%.

Table 6: Regression Results, Respondent's probability of ever marrying following a non-marital birth

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Variable	**Men and Women**			**Women**			**Men**		
First-born is a girl	0.024 (0.021)	-0.001 (0.018)	-0.001 (0.018)	-0.012 (0.027)	-0.031 (0.022)	-0.031 (0.022)	0.074** (0.035)	0.036 (0.030)	0.036 (0.029)
Age		0.027*** (0.004)	0.013** (0.006)		-0.023*** (0.004)	-0.013* (0.007)		-0.025*** (0.006)	0.008 (0.010)
Age squared		0.000*** (0.000)	0.000*** (0.000)		0.000*** (0.000)	0.000*** (0.000)		0.000*** (0.000)	0.000 (0.000)
Age of FBC		-0.024*** (0.002)	-0.020*** (0.004)		-0.028*** (0.002)	-0.025*** (0.005)		-0.022*** (0.003)	-0.018*** (0.006)
University degree		0.035 (0.036)	0.043 (0.036)		0.043 (0.043)	0.050 (0.043)		0.046 (0.060)	0.057 (0.061)
College diploma		0.041 (0.027)	0.041 (0.027)		-0.004 (0.033)	-0.002 (0.033)		-0.104** (0.044)	-0.101** (0.045)
Some post-secondary		0.005 (0.033)	0.009 (0.033)		-0.022 (0.040)	-0.018 (0.040)		0.048 (0.057)	0.050 (0.056)
High school dropout		0.012 (0.026)	0.013 (0.025)		-0.032 (0.032)	-0.033 (0.032)		0.062 (0.042)	0.066 (0.042)
Other education		-0.328*** (0.082)	-0.357*** (0.083)		-0.376*** (0.145)	-0.399*** (0.147)		-0.299*** (0.064)	-0.346*** (0.068)
Newfoundland		0.025 (0.034)	0.026 (0.034)		0.035 (0.046)	0.031 (0.046)		0.000 (0.052)	0.012 (0.050)
Prince-Edward Island		0.045 (0.048)	0.040 (0.048)		0.019 (0.060)	0.015 (0.059)		0.046 (0.080)	0.042 (0.080)
Nova Scotia		0.006 (0.038)	0.003 (0.038)		-0.022 (0.045)	-0.023 (0.046)		0.083 (0.063)	0.087 (0.062)
New Brunswick		-0.045 (0.045)	-0.047 (0.045)		-0.014 (0.055)	-0.010 (0.056)		-0.111 (0.075)	-0.124* (0.074)
Quebec		-0.211*** (0.025)	-0.212*** (0.025)		-0.195*** (0.031)	-0.199*** (0.031)		-0.235*** (0.041)	-0.228*** (0.042)
Manitoba		-0.006 (0.043)	-0.009 (0.043)		-0.038 (0.055)	-0.045 (0.055)		0.029 (0.066)	0.031 (0.066)
Saskatchewan		-0.045 (0.041)	-0.054 (0.041)		-0.031 (0.049)	-0.035 (0.049)		-0.090 (0.071)	-0.099 (0.071)
Alberta		-0.006 (0.035)	-0.002 (0.035)		0.024 (0.042)	0.025 (0.042)		-0.060 (0.057)	-0.047 (0.059)
British Columbia		-0.004 (0.031)	-0.006 (0.031)		-0.037 (0.039)	-0.041 (0.039)		0.033 (0.050)	0.035 (0.050)
FBC born in 1940s			0.135 (0.185)			0.112 (0.231)			0.202 (0.301)
FBC born in 1950s			0.189 (0.154)			0.139 (0.191)			0.270 (0.251)
FBC born in 1960s			-0.242** (0.122)			0.181 (0.150)			0.319 (0.201)
FBC born in 1970s			-0.221** (0.097)			0.151 (0.118)			-0.319** (0.161)
FBC born in 1980s			-0.236*** (0.064)			-0.194** (0.078)			-0.285*** (0.108)
FBC born in 1990s			-0.169*** (0.037)			-0.114*** (0.043)			-0.229*** (0.066)
N		2,963			1,907			1,056	

Note: All regressions are weighted. Robust standard errors are in parentheses. *Significant at 10%; **significant at 5%; ***significant at 1%.

Table 7: Regression Results, Respondent's probability of being separated or divorced at time of survey

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Variable	**Men and Women**			**Women**			**Men**		
First-born is a girl	0.011 (0.007)	0.010 (0.007)	0.010 (0.007)	0.002 (0.010)	0.000 (0.010)	0.000 (0.010)	0.021** (0.010)	0.021** (0.009)	0.022** (0.010)
Age		0.011*** (0.002)	0.004 (0.003)		0.010*** (0.003)	0.003 (0.004)		0.012*** (0.002)	0.006* (0.003)
Age squared		0.000*** (0.000)	0.000*** (0.000)		0.000*** (0.000)	0.000*** (0.000)		0.000*** (0.000)	0.000*** (0.000)
Age of FBC		0.009*** (0.001)	0.009*** (0.001)		0.010*** (0.001)	0.008*** (0.002)		0.006*** (0.001)	0.009*** (0.002)
University degree		-0.014 (0.011)	-0.012 (0.011)		0.011 (0.016)	0.013 (0.016)		-0.038** (0.015)	-0.037** (0.015)
College diploma		0.005 (0.011)	0.006 (0.011)		0.019 (0.015)	0.020 (0.015)		-0.010 (0.015)	-0.010 (0.015)
Some post-secondary		0.027* (0.015)	0.027* (0.015)		0.026 (0.020)	0.027 (0.020)		0.029 (0.022)	0.028 (0.022)
High school dropout		-0.008 (0.011)	-0.005 (0.011)		-0.005 (0.016)	-0.002 (0.016)		-0.012 (0.016)	-0.010 (0.016)
Other education		0.034 (0.062)	0.035 (0.062)		0.039 (0.121)	0.047 (0.118)		0.027 (0.064)	0.028 (0.064)
Newfoundland		-0.055*** (0.013)	-0.055*** (0.013)		-0.061*** (0.019)	-0.061*** (0.019)		-0.043** (0.017)	-0.044*** (0.017)
Prince-Edward Island		0.004 (0.022)	0.004 (0.022)		-0.030 (0.029)	-0.028 (0.029)		0.044 (0.033)	0.041 (0.033)
Nova Scotia		0.011 (0.016)	0.011 (0.016)		0.001 (0.021)	0.001 (0.021)		0.024 (0.023)	0.023 (0.023)
New Brunswick		0.005 (0.019)	0.005 (0.019)		-0.026 (0.023)	-0.023 (0.023)		0.042 (0.032)	0.041 (0.032)
Quebec		0.059*** (0.010)	0.059*** (0.010)		0.057*** (0.015)	0.056*** (0.015)		0.065*** (0.015)	0.064*** (0.015)
Manitoba		-0.018 (0.014)	-0.019 (0.014)		-0.023 (0.020)	-0.024 (0.020)		-0.011 (0.018)	-0.011 (0.018)
Saskatchewan		-0.026* (0.014)	-0.027* (0.014)		-0.042** (0.020)	-0.042** (0.020)		-0.008 (0.019)	-0.010 (0.019)
Alberta		0.018 (0.013)	0.018 (0.013)		0.020 (0.019)	0.020 (0.019)		0.018 (0.017)	0.018 (0.017)
British Columbia		0.016 (0.010)	0.016 (0.010)		0.018 (0.015)	0.019 (0.015)		0.013 (0.014)	0.014 (0.014)
FBC born in 1940s			-0.013 (0.069)			0.128 (0.101)			-0.174* (0.094)
FBC born in 1950s			-0.018 (0.057)			0.100 (0.083)			-0.149* (0.077)
FBC born in 1960s			0.017 (0.046)			0.124* (0.067)			-0.102* (0.062)
FBC born in 1970s			0.046 (0.035)			0.131** (0.053)			-0.044 (0.045)
FBC born in 1980s			0.052** (0.023)			0.103*** (0.034)			-0.001 (0.030)
FBC born in 1990s			0.030** (0.013)			0.065*** (0.020)			-0.005 (0.018)
N		10,801			6,209			4,592	

Note: All regressions are weighted. Robust standard errors are in parentheses. *Significant at 10%; **significant at 5%; ***significant at 1%.

Table 8: Regression Results, Respondent's probability of a first marriage ending in divorce or separation

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Variable	**Men and Women**			**Women**			**Men**		
First-born is a girl	0.014* (0.008)	0.013* (0.008)	0.014* (0.008)	0.010 (0.011)	0.010 (0.011)	0.009 (0.011)	0.019* (0.011)	0.019* (0.011)	0.020* (0.011)
Age		0.014*** (0.002)	0.004 (0.003)		0.014*** (0.002)	0.003 (0.004)		0.013*** (0.003)	0.005 (0.004)
Age squared		0.000*** (0.000)	0.000*** (0.000)		0.000*** (0.000)	0.000*** (0.000)		0.000*** (0.000)	0.000*** (0.000)
Age of FBC		0.015*** (0.001)	0.014*** (0.002)		0.017*** (0.001)	0.014*** (0.002)		0.014*** (0.001)	0.015*** (0.002)
University degree		0.007 (0.012)	0.008 (0.012)		0.019 (0.016)	0.019 (0.017)		-0.005 (0.017)	-0.004 (0.017)
College diploma		0.013 (0.012)	0.014 (0.012)		0.017 (0.016)	0.018 (0.016)		0.008 (0.017)	0.008 (0.017)
Some post-secondary		0.070*** (0.018)	0.070*** (0.018)		0.084*** (0.024)	0.084*** (0.024)		0.053** (0.026)	0.053** (0.026)
High school dropout		-0.028** (0.012)	-0.025** (0.012)		-0.032** (0.017)	-0.028* (0.017)		-0.026 (0.019)	-0.025 (0.019)
Other education		-0.012 (0.052)	-0.012 (0.051)		0.037 (0.126)	0.040 (0.121)		-0.046 (0.037)	-0.046 (0.038)
Newfoundland		-0.067*** (0.015)	-0.068*** (0.015)		-0.075*** (0.021)	-0.075*** (0.021)		-0.061*** (0.022)	-0.061*** (0.022)
Prince-Edward Island		-0.010 (0.027)	-0.012 (0.027)		-0.035 (0.033)	-0.038 (0.033)		0.017 (0.044)	0.016 (0.044)
Nova Scotia		-0.018 (0.017)	-0.018 (0.017)		-0.016 (0.022)	-0.016 (0.022)		-0.020 (0.025)	-0.020 (0.025)
New Brunswick		-0.018 (0.020)	-0.016 (0.020)		-0.016 (0.026)	-0.012 (0.026)		-0.018 (0.031)	-0.019 (0.031)
Quebec		0.040*** (0.011)	0.039*** (0.011)		0.037** (0.015)	0.036** (0.015)		0.047*** (0.017)	0.046*** (0.017)
Manitoba		-0.042*** (0.015)	-0.044*** (0.014)		-0.051** (0.021)	-0.053*** (0.021)		-0.033 (0.020)	-0.035* (0.020)
Saskatchewan		-0.019 (0.017)	-0.022 (0.017)		-0.044** (0.022)	-0.045** (0.022)		0.010 (0.025)	0.006 (0.025)
Alberta		0.004 (0.014)	0.004 (0.014)		0.004 (0.019)	0.004 (0.019)		0.004 (0.019)	0.004 (0.019)
British Columbia		0.018 (0.012)	0.019 (0.012)		0.019 (0.016)	0.021 (0.016)		0.018 (0.017)	0.018 (0.017)
FBC born in 1940s			-0.004 (0.076)			0.110 (0.104)			-0.119 (0.114)
FBC born in 1950s			0.009 (0.063)			0.094 (0.087)			-0.083 (0.092)
FBC born in 1960s			0.043 (0.051)			0.123* (0.070)			-0.055 (0.074)
FBC born in 1970s			0.075** (0.038)			0.140*** (0.054)			-0.004 (0.053)
FBC born in 1980s			0.044* (0.024)			0.075** (0.034)			0.006 (0.035)
FBC born in 1990s			-0.002 (0.010)			0.018 (0.015)			-0.024 (0.015)
N		9,803			5,854			3,949	

Note: All regressions are weighted. Robust standard errors are in parentheses. *Significant at 10%; **significant at 5%; ***significant at 1%.

Table 9: Regression Results, Respondent's probability of remarriage following a divorce

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Variable	**Men and Women**			**Women**			**Men**		
First-born is a girl	0.018	0.011	0.013	0.060	0.045	0.048	-0.046	-0.051	-0.053
	(0.029)	(0.028)	(0.028)	(0.037)	(0.035)	(0.035)	(0.046)	(0.044)	(0.043)
Age		-0.003	-0.003		-0.023	-0.016		0.018	0.018
		(0.011)	(0.016)		(0.014)	(0.021)		(0.017)	(0.026)
Age squared		0.000	0.000		0.000	0.000		0.000**	0.000
		(0.000)	(0.000)		(0.000)	(0.000)		(0.000)	(0.000)
Age of FBC		0.020***	0.011*		0.027***	0.022***		0.022***	0.000
		(0.003)	(0.006)		(0.004)	(0.007)		(0.005)	(0.009)
University degree		0.100**	0.103**		-0.008	-0.008		0.205***	0.220***
		(0.046)	(0.046)		(0.057)	(0.057)		(0.071)	(0.071)
College diploma		0.028	0.030		0.041	0.043		-0.006	-0.002
		(0.039)	(0.039)		(0.049)	(0.049)		(0.065)	(0.064)
Some post-secondary		0.129**	0.126**		0.174***	0.171***		0.056	0.055
		(0.052)	(0.052)		(0.065)	(0.065)		(0.084)	(0.083)
High school dropout		0.043	0.043		0.052	0.050		-0.009	-0.001
		(0.042)	(0.042)		(0.053)	(0.054)		(0.069)	(0.067)
Other education		-0.327**	-0.329**		-0.328***	-0.330***		-0.160	-0.098
		(0.131)	(0.143)		(0.111)	(0.118)		(0.319)	(0.346)
Newfoundland		-0.260***	-0.259***		-0.336***	-0.334***		-0.187*	-0.173
		(0.068)	(0.070)		(0.075)	(0.076)		(0.104)	(0.107)
Prince-Edward Island		0.085	0.088		0.051	0.061		0.133	0.110
		(0.100)	(0.099)		(0.127)	(0.127)		(0.169)	(0.170)
Nova Scotia		-0.072	-0.066		-0.089	-0.083		-0.034	-0.023
		(0.062)	(0.062)		(0.075)	(0.075)		(0.103)	(0.101)
New Brunswick		-0.026	-0.016		0.027	0.033		-0.118	-0.092
		(0.076)	(0.076)		(0.087)	(0.088)		(0.142)	(0.144)
Quebec		-0.218***	-0.224***		-0.178***	-0.179***		-0.264***	-0.281***
		(0.037)	(0.037)		(0.047)	(0.047)		(0.061)	(0.059)
Manitoba		-0.044	-0.056		-0.041	-0.040		-0.023	-0.076
		(0.076)	(0.078)		(0.103)	(0.104)		(0.113)	(0.112)
Saskatchewan		0.010	-0.002		0.051	0.046		-0.063	-0.090
		(0.063)	(0.064)		(0.088)	(0.089)		(0.098)	(0.099)
Alberta		-0.023	-0.023		-0.083	-0.081		0.067	0.075
		(0.052)	(0.052)		(0.062)	(0.062)		(0.079)	(0.079)
British Columbia		0.033	0.033		0.051	0.053		-0.006	-0.024
		(0.043)	(0.043)		(0.055)	(0.055)		(0.066)	(0.067)
FBC born in 1940s									1.227***
									(0.378)
FBC born in 1950s			-0.124			-0.118			0.954***
			(0.099)			(0.132)			(0.304)
FBC born in 1960s			-0.247*			-0.193			0.658***
			(0.138)			(0.181)			(0.238)
FBC born in 1970s			-0.294*			-0.223			0.530***
			(0.174)			(0.227)			(0.179)
FBC born in 1980s			-0.356*			-0.237			0.317***
			(0.210)			(0.272)			(0.117)
FBC born in 1990s			-0.535**			-0.333			
			(0.247)			(0.313)			
N		1,741			1,092			649	

Note: All regressions are weighted. Robust standard errors are in parentheses. *Significant at 10%; **significant at 5%; ***significant at 1%.