PARENTAL LEAVE POLICIES AND THEIR CONSEQUENCES FOR INEQUALITY

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Ankita Patnaik

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Ankita Patnaik, Ph.D.

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In order to increase parents' participation in parental leave, in 20006 Quebec reformed its paid leave program to to offer higher benefits and institute a 'daddy-only' quota that reserved 5 weeks for fathers. In this dissertation I investigate the effects of this landmark reform on various dimensions of inequality.

In chapter 1, I analyze whether QPIP affected the gender gap in parents' leave participation. Using data on benefit claims, I find that QPIP had an immediate program effect of increasing fathers' leave participation by 53 percentage points and leave duration by 3.1 weeks, with no immediate effect on mothers' leave behavior. I find evidence that the 'daddy-only' quota produces an intra-household fly-paper effect: even though the quota does not change a binding constraint for most families in Quebec, the 'daddy' benefits stick to fathers. This suggests that one of the reasons that daddy quotas are effective is that they produce a labeling effect from the 'daddy-only' label.

In chapter 2, I investigate how this exogenous increase in fathers' leave taking under QPIP may have affected household sex-specialization in the long-term. I utilize data from time-diaries, and exploit variation in exposure to QPIP across provinces, time and the age of one's children. I find that QPIP had a large and persistent effect on the division of household labor. In exposed households, fathers experience decreased time in market work and personal income, while mothers experience increased time at the workplace, labor supply, and personal income. The organization of non-market work also changes: Fathers increase time in housework, while mothers move time away from housework and towards childcare

instead. Overall, households exposed to QPIP are found to be less sex specialized.

Chapter 3 explores whether QPIP reduced inequalities in leave-utilization across socioeconomic strata, and in turn reduced health inequalities. I find that QPIP increased mothers' leave participation, particularly among never-married mothers and low-income mothers, who previously took significantly less leave. On average QPIP increased breastfeeding initiations by 6% and increased the duration of breastfeeding and exclusive breastfeeding, but the program effects favored married, educated, high-income mothers, suggesting increasing health inequalities.

BIOGRAPHICAL SKETCH

Ankita Patnaik completed her undergraduate education at the University of Edinburgh while receiving the Undergraduate India Scholarship for 2006-2010. It was while writing her Senior Honors Thesis, which decomposed the motherhood wage penalty in the United Kingdom, that she developed research interests in issues related to gender in the labor market. She graduated in 2010 with Honors and First-Class Distinction and was awarded the Mary Teresa Rankin Prize for Best Economics Graduate at the University.

In August 2010 Ankita enrolled in the graduate program at the Economics department of Cornell University. Under the guidance of Francine D. Blau, John M. Abowd and Michael F. Lovenheim, she specialized in the fields of labor economics, applied microeconomics, and family and social welfare policy. Ankita received her M.A. in Economics in December 2013 and her Ph.D. in Economics in August 2015.

For Priyabrata and Chirasmita Patnaik

I would like to dedicate this dissertation to my father, who ingrained in me the idealism and silly optimism that convinces one to pursue a Ph.D - and my mother, without whom I would not have the pragmatism and discipline that it takes to finish one.

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These analyses use restricted-access data from Statistics Canada's Employment Insurance Coverage Survey (2002-2010), General Social Survey (2005, 2010) and Canadian Community Health Survey (2005, 2007-2013). All computations, use and interpretation of these data are entirely that of the author. All results have been checked for confidentiality disclosure and released by Statistics Canada. I thank the staff at Statistics Canada and the Research Data Center at York University for their technical support.

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CHAPTER 1

MAKING PATERNITY LEAVE EASIER: BETTER COMPENSATION AND DADDY-ONLY ENTITLEMENTS

1.1 Introduction

Job-protected parental leave mandated are common in developed countries, with the aim of promoting the welfare of infants and parents. The leave provisions vary considerably internationally - they tend to be long, universal and generously compensated in European countries, whereas they are short, restricted and unpaid in most of the United States.[1] The central aim of maternity leave is to allow mothers to fully recover from childbirth and to form a bond with their babies. Other rationales for providing parental leave include maintaining a productive economy by retaining female workers, sustaining birth rates, decreasing unemployment and relieving some of the parenting deficit that is growing alongside the increasing incidence of dual-earner parents with long working hours (Haas, 1992). Further, as the single breadwinner model increasingly gives way to the dual-earner household, another increasingly common objective of parental leave reforms is to promote gender equality. There has been a trend in policy-making, beginning in Scandinavia but now catching on in other countries, towards promoting equality by encouraging fathers to take parental leave. Such policies aim to increases fathers' contact with their infants, invest in men's caregiving skills, reduce work-family frictions by labeling working men as fathers, and offer a supportive home environment for working mothers by reducing the burden of childcare and domestic work that falls on them. These policies thus aim to strengthen the ties of fathers to their family and simultaneously the ties of mothers to working life.

[1] By state mandate, pregnancy- and childbirth-related leave are available to eligible employees in California, Colorado (for public employees), Hawaii, New Jersey, New York, and Rhode Island via Temporary Disability Insurance. Further, California, New Jersey and Rhode Island have established Paid Family Leave Programs to augment their existing TDI Programs

As fathers' participation in parental leave programs has become a notable area of policy debate in many OECD countries, this begs the question: what kinds of schemes are successful in getting fathers to take leave, and what are the mechanisms behind that success? In this study, I explore this question while investigating a landmark reform to parental leave in Canada. I investigate the Quebec Parental Insurance Program (QPIP), a system of parental leave benefits introduced in Quebec in 2006 that explicitly sought to boost fathers' participation in paid parental leave. From 2001 to 2005, eligible parents in all Canadian provinces could claim parental leave benefits from the government through the Employment Insurance (EI) Program. Prior to the reform, fathers only had access to 'shared' parental leave with their spouses, and leave-takers were compensated with a little over half their wages up to a strict cap so that household incomes were hit hard when fathers took leave. Consequently, fathers' leave participation in Quebec never exceeded 22% prior to QPIP. Notably, the majority of families did not exhaust their total amount of leave prior to the reform, such that families were leaving benefits 'on the table' even as fathers declined to participate. In 2006, Quebec left the EI system and established the Regime Quebecois D'assurance Parentale or the Quebec Parental Insurance Plan. This new scheme lowered eligibility criteria, increased income replacement, and established a 5-week 'daddy quota' of leave for fathers (Doucet et al., 2010). Due to QPIP's 'daddy quota', Quebec is the only province in Canada in which fathers enjoy an individual and non-transferable right to parental leave.

This study makes several contributions to the literature. It is the first study to explore the causal mechanisms behind why daddy quotas may be effective. That is, I am able to not only analyze the extent of QPIP's impact on parents' leave behavior, but also investigate whether quotas can succeed in getting fathers involved by forcing their hand, or by eliciting a behavioral response to the 'daddy-only' label. Consequently, this paper is the first to explore the possibility of an intra-household flypaper effect (IHFE) in parental leave, where leave that is labeled as daddy-only 'sticks' to fathers even if the quota does not change a binding constraint. Second, this is the first study to date to study how this Canadian policy episode affected parents' leave behavior. This is interesting because Canada offers a political and social context that is quite different from the previously-studied Scandinavian countries, since the latter have some of the most generous welfare provisions and family-friendly policies worldwide.Third, since only the province of Quebec deviated from the national policy, this study is also unique in utilizing regional variation in policy rather than a nationwide change in policy. This results in improved study design: by using other Canadian provinces where policy did not change as a natural control group, I can provide causal estimates that are robust to various trends.

I use data on benefit claims from the 2002-2010 rounds of the Employment Insurance Coverage Survey (EICS). I use a sharp regression discontinuity design to identify the local mean impact of QPIP at the point when it was introduced, and a difference-in-differences approach to estimate the average treatment effect of QPIP since it has been introduced. Both sets of results show that QPIP was very effective in achieving its goal of boosting fathers' involvement. The introduction of QPIP was associated with an increase in fathers' claim rates of 53 percentage points and an increase in fathers' leave duration of 3 weeks. There is some evidence that QPIP also increased mothers' participation, but the effect is much smaller than that for fathers, in both absolute and relative magnitude. My results suggest that fathers responded to not only the higher benefits but also the 'daddy-only' label associated with the quota. Since the majority of families did not exhaust their leave before the reform, the new daddy quota did not alter a binding constraint for them. Nevertheless, reserving some weeks as 'daddy-only' shifted the distribution of leave towards fathers - that is, the new program induced fathers to take leave that they would have had available even prior to QPIP. More tellingly, the average father in post-reform Quebec consumed exactly 5 weeks of paid leave- they did not increase their consumption beyond the amount allocated by the quota even when there were unused weeks of parental leave syill available. This paper thus provides novel evidence of an intra-household flypaper effect in parental leave, whereby labeling some weeks as 'daddy-only' can make those weeks 'stick' to fathers. This is an odd and important finding in terms of policy design, as it suggests that labeling may play an important role in influencing program participation.

The rest of the paper is structured as follows. Section 1.2 provides details on the Canadian reform and discusses the expected effects on leave behavior. Section 1.3 reviews prior literature on the effectiveness of policies promoting paternity leave. Section 1.4 describes the data used in my analysis. Section 1.5 explains the methods and results from an analysis of the immediate impact of QPIP, while Section 1.6 does the same for an analysis of the average treatment effect of QPIP. Section 1.7 concludes.

1.2 Background

1.2.1 Parental Leave Programs in Canada and the QPIP Reform

In every Canadian province, at least a year of job-protected parental leave is available to every parent who has worked 52 weeks or more with their current employer.[2] Further, every parent who meets certain eligibility criteria can claim benefits, converting some of this leave into paid leave. The Employment Insurance (EI) Program, which all Canadian provinces used until 2005, offers maternity benefits that mothers can take in the weeks immediately succeeding the birth as well as parental benefits that mothers and fathers must decide how to share between them. Most provinces continue to subscribe to the EI Program, with the

[2]The length of job-protected leave in Canadian provinces does not change over the period of my analysis.

notable exception of Quebec. On the 1st of January 2006, Quebec instituted the Quebec Parental Insurance Plan (QPIP), to which employees now contribute and claim benefits from instead of the EI system. It should be noted that both the EI and QPIP program are financed through payroll taxes. The details of the EI program, currently offered to residents of other provinces, and the QPIP program, currently offered to residents of Quebec, are shown in Table 1.1.

QPIP's features were designed to offer an improvement over the older EI system by easing some of the barriers that parents face to taking leave, namely, inflexibility, ineligibility, financial feasibility, and gendered attitudes. First, the new system was designed to be more flexible, offering parents a choice between the Basic Plan or a Special plan that offers higher benefits for a shorter duration, thereby letting parents select the combination of benefit amount and duration that best suited their needs. Second, the reform lowered the eligibility criteria in order to improve coverage and ease access to benefits. The EI system requires a claimant to have worked 600 hours of insurable employment. This makes it difficult for workers from seasonal, temporary, part-time or otherwise non-standard employment, who tend disproportionately to be low-income mothers, to qualify for benefits. In comparison, QPIP uses an earnings-based threshold that is easier to meet, such that any parent who has at least 2000CAD of insurable earnings can qualify. Third, QPIP offers more generous compensation for foregone income. By both increasing the maximum replacement rate (from 55% to 70%) and raising the ceiling of maximum insurable earnings on which one can claim (from 39,000CAD to 57,000CAD in 2006), QPIP ensures that a greater portion of foregone wages can be recovered via benefits while on parental leave.

QPIP also introduced the nation's first of its kind 'daddy quota', whereby 5 weeks of leave (or 3 weeks under the Special Plan) were set aside for the father and could not be transferred to the mother. This important feature of the reform stands in stark contrast to the EI Program, where fathers enjoy no individual right to paternity leave and may only access benefits through shared parental leave. More generally, QPIP changed the distribution of benefits within the household. QPIP abolished the 2-week waiting period that EI claimants are subject to. The amount of gender-neutral leave to be shared between parents was reduced and some weeks were reallocated to individual non-transferable leave for each parent. The net result was that mothers retained access to the same amount of potential leave as before (50 weeks of paid leave) but a larger share now came through maternity leave rather than shared parental leave. Fathers gained access to more leave than they had earlier: 37 potential weeks under QPIP (5 of which are 'daddy-only') versus 35 weeks under the EI Program. QPIP increased the amount of paid leave available to a family from 50 weeks to 55 weeks, such that total leave increased by the amount equivalent to the 'daddy-only' weeks.[3]

1.2.2 Expected Impact of QPIP on Parents' Leave Behavior

QPIP's choice of two programs and easier eligibility criteria are not expected to impact fathers significantly. Since the majority of fathers are full-time, full-year workers, they face no difficulty qualifying for benefits under either the EI or the QPIP scheme.[4] Further, since under QPIP the whole family had to act on either the Basic Plan or the Special Plan once the choice was made, few families selected the Special Plan, which limited their duration in return for higher compensation. Therefore, the two changes most likely to influence the decision for fathers to take leave were that of improved benefits and the daddy quota.

First, I consider how the representative parent might respond to increased income re-

[3]It should be noted that under either the EI or QPIP program, parents can take leave simultaneously so the mother does not have to resume work in order for the father to participate in parental leave.

[4]In my data, I find no statistically significant change in the proportion of husbands reported ineligible for parental leave benefits between 2005 and 2007 (the years surrounding the reform).

placement. By making benefits more generous, QPIP reduced the opportunity cost of taking leave, i.e. the difference between foregone wages and benefits. Assuming leave is a normal good, the price reduction should result in positive income and substitution effects, leading to an unambiguous increase in the amount of leave consumed. However, it should be noted that benefits increased for both mothers and fathers, so there is no reason to believe fathers should respond more strongly than mothers to the improved benefits. To the contrary, given the evidence showing that married women have more elastic labor supplies than men (Juhn and Murphy, 1997; Blau and Kahn, 2007), if anything we would expect mothers to respond more strongly to the improved benefits.[5] Further, as benefits are capped at a certain threshold, lower-income parents experience a larger marginal reduction in the price of leave under QPIP since they face a smaller wage-benefit differential. Therefore, with respect to the increase in financial benefits under QPIP, we should expect mothers, who tend to be the lower-earning spouse and have more elastic labor supplies, to respond more strongly than fathers.

Second, I consider the reservation of the daddy quota. A daddy quota could make the difference between a father participating or not participating if, absent the quota, his wife consumed the total amount of leave allocated to the family. In that case, the addition of 5 daddy-only weeks would make it necessary for the father to participate for the family to continue exhausting total family leave. However, Quebec presents an interesting case because prior to the quota, most families did not use all of their leave. Figure 1.1 presents the probability density function of the distribution of maternal leave duration in Quebec in the period before the reform. Although there was bunching at the cap of 52 weeks, i.e. 12 months, a significant portion of mothers were not consuming all the paid leave available to the household. Figure 1.2 shows the cumulative distribution function of mothers' leave duration in Quebec and the other provinces in the pre-reform period (2002-2005). Even when

[5] Blau and Kahn (2007) report that accounting for the presence of young children diminishes estimates of the own-wage elasticity of married women only very slightly.

the father did not take leave (the majority of families), over 60% of mothers reported taking 11 months or fewer of leave, leaving at least one full month of benefits unused. Furthermore, since the EICS survey asks mothers about all leave taken and not specifically paid parental leave, their answers include the 2 weeks of unpaid 'waiting period' under the EI program, and may even include other kinds of paid leave such as vacation or sick days. For example, a mother who reports taking 11 months of leave is at most taking 10.5 months of paid parental leave, a full 1.5 months less than the maximum the family is entitled to. This implies that for these families, a considerable amount of leave was always available to fathers even though they chose not to use it. Therefore, for the majority of families who were not consuming at the cap pre-reform, the newly imposed constraint of the daddy quota should not have been binding. Accordingly, any increase in total family leave under QPIP should have been considered an ordinary extension of family leave since the additional weeks were essentially fungible between parents.

Given that the constraint does not bind for most families, I investigate whether the daddy quota could alter the parents' maximization problem in a different way. To do so, I consider the existence of a non-monetary cost of leave. In addition to the opportunity cost of taking leave (wages net of benefits), parents may face a non-monetary 'stigma' cost that causes them to discount benefit income compared to wage income. This 'stigma' cost could encompass any number of things, for example, personal distaste to taking leave, peer pressure or workplace hostility to leave-takers. Further, this cost may differ across individuals, for example, stigma may be higher for men than women, or for men working in blue-collar environments versus those working in white-collar environments. Differences in stigma may therefore contribute to the differences in leave participation rates across genders as well as income or education groups. One mechanism through which a daddy quota can have an impact even when the constraint does not bind is if the 'daddy-only' label for the quota reduces this stigma cost for men. The daddy-only label establishes a father's individual right to leave, removes the need to negotiate with his wife, and improves his bargaining position with employers and co-workers who may be more sympathetic to him using leave specifically designated for him. Moreover, the quota sends a clear public message that promotes fathers' involvement, which may reduce social stigma against taking leave and possibly even introduce stigma against those who do not utilize this generous opportunity to spend time with their children. The idea that fathers respond to reduced social or workplace stigma is consistent with the finding of Dahl et al. (2014) that fathers are more likely to take parental leave if their brothers or coworkers have done so. Therefore, under QPIP fathers may have experienced a reduction in not only the opportunity cost of taking leave but also the stigma cost. Mothers, however, only experienced the former. This difference might lead to fathers responding more strongly to the reform than mothers.

We thus can consider two alternate hypotheses:

(i) H_0: The daddy-only label does not affect stigma costs, or no stigma cost exists

In this case, parents would respond only to the opportunity cost of taking leave. Since QPIP's improved benefits lowered the opportunity cost for both males and females, we should see leave consumption increase for both parents. Given that, relative to men, married women have higher elasticities of labor supply (Juhn and Murphy, 1997; Blau and Kahn, 2007) and tend to be the lower-earning spouse (thereby experiencing a greater marginal increase in benefits), we expect mothers to increase their leave consumption by at least as much as fathers.

(ii) H_1: A stigma cost exists, and the daddy-only label reduces it for men

If there exists a stigma cost that is higher for fathers than mothers, which QPIP's 'daddy-only' label reduces for men, then fathers may increase their leave consumption by more than mothers. That is, if we observe an intra-household flypaper effect, whereby benefits stick to the fathers even though the additional weeks are fungible for most families, this would be evidence consistent with the existence of a stigma cost.

1.3 Previous Research on the Effectiveness of Paternity Leave Policies

Despite the considerable evidence that fathers' involvement in childcare is positively associated with children's social, emotional, physical, and cognitive development (Allen and Daly (2007) provide a useful summary), leave participation rates of fathers worldwide remain much lower than those of mothers. Since the father is often the higher-earning parent, financial compensation plays a significant role in their decision to take leave. Studies have shown that loss of earnings is an important factor in fathers' decisions to not take parental leave (Zhelyazkova, 2013). It is also common for fathers to cite workplace attitudes as an obstacle to utilizing leave even when they are entitled to it, out of fear it could damage their careers (Bygren and Duvander, 2006). Social and psychological factors also may play a role: it is possible that men have a lower taste for childcare, that social constructs push men to see themselves as the primary breadwinner who must prioritize paid work, or that they are rarely exposed to role models in the form of men who care for infants.

Several studies have exploited cross-country variation in policies to determine how easing these barriers can improve fathers' leave-taking. Fathers' leave take-up tends to be higher in countries with generous compensation rates (Moss and O'Brien, 2006) and is especially low in countries like the United States where leave is unpaid (Han et al., 2007). O'Brien (2009) compares 24 countries and finds fathers' use of statutory leave is greatest when high income replacement (fifty percent or more of earnings) is combined with extended duration (more than fourteen days). It also matters whether fathers' access to leave is derived via a family right or an individual right. Several cross-country comparisons have shown that fathers are more likely to utilize leave in countries that have a daddy quota in place (Bruning and Plantenga, 1999; O'Brien, 2009; Haas and Rostgaard, 2011). However, while these findings provide suggestive associations between different kinds of leave policies and fathers' behavior, they suffer from endogeneity issues since the assignment of each country to a specific policy regime is non-random. That is, a country may offer high income replacement precisely because parents are highly motivated or concerned about parental leave.

More recently a few studies have exploited natural experiments, where leave policy was changed suddenly, to identify causal effects by comparing births just before and just after the reform. Dahl et al. (2014) report that the introduction of a daddy quota in Norway had an impact on fathers' takeup of 32 percentage points. Duvander and Johansson (2012) and Ekberg et al. (2013) study Sweden and find a strong effect on parental leave use resulting from the reservation of the first 'daddy month'.[6] These studies present causal estimates of the impact of daddy quotas and provide some evidence of the success of such schemes. However, the specific nature of these reforms present limitations on the ways in which we can interpret the program effects. In the case of Sweden, the daddy quota did not represent the addition of a new month of leave, but instead a transfer from total family leave to 'daddy-only' leave. Thus, if the mother had previously exhausted the total leave, the quota now made it necessary for the father to participate to simply maintain the status quo amount of family leave. In Norway, the introduction of the quota did not decrease mothers' potential leave, but since most mothers took the entire amount of family leave prior to the reform (Dahl et al. (2014), pp.2), the family could only use the additional leave if the father utilized

[6] Duvander and Johansson (2012) detect a smaller but clear effect resulting from the second daddy month, and no effect from the 'gender equality bonus' which rewards couples for sharing leave equally.

his month. Therefore, in the case of both Norway and Sweden, the introduction of the quota altered a binding constraint. We therefore cannot be sure about the mechanisms behind the reforms' success, i.e., were fathers responding to their individual right and the 'daddy-only' label - or were families simply trying to maximize leave, which made it necessary for fathers to participate?

1.4 Data

To analyze the immediate impact of QPIP on parents' leave behavior, I use data on benefit claims collected through the Employment Insurance Coverage Survey (EICS) (Statistics Canada, 2002-2010). The target population for this annual survey is a subset of the target population for the Labor Force Survey, and comprises individuals who, given their recent status in the labor market, could potentially be eligible for employment insurance. Mothers of infants less than one year old, who I will focus on in this study, fall into this last category, since they could potentially be eligible for benefits via maternity or parental leave. The EICS is conducted annually, and I focus on mothers in a nine-year window framing the QPIP reform, from 2002 to 2010. Specifically, I use data from 2002-2005 as the pre-reform period (roughly 42% of the observations), and 2006-2010 as the post-reform period.[7] It should be noted that I use restricted-access versions of this data which can only be accessed on-site at a Statistics Canada Remote Data Center, as the Public Use Microdata do not have detailed information on month of birth, fathers' leave duration and household income.

The primary sample comprises 8,907 observations of mothers aged 18-40 who have a child

[7]There were nation-wide reforms to both job-protected and paid parental leave in late 2000, and Quebec also extended its publicly subsidized childcare to children aged 0 to 1 in 2001. This motivates me to exclude data from survey years prior to 2002 as well as any observations in survey year 2002 which report the birth year as 2001.

under one year old and identify as part of a married or cohabitating couple.[8] Approximately one-fifth of the observations are from Quebec. The rest of the observations come from the control group, which comprises the five largest other provinces, i.e. Ontario, Alberta, British Columbia, Atlantic Region, and Manitoba and Saskatchewan, where the EI system remained in place over the entire period of the analysis.

The outcomes regarding leave participation are measured by indicators taking value 1 if the respondent (or her spouse) has claimed or plans to claim maternity/parental/paternity benefits through the EI or QPIP system. Parents' leave duration is measured by mothers' reports of total weeks of actual or planned leave taken by her and her spouse. My measures of leave duration are not conditional on participation and so include zeros, thus offering a summary measure that takes into account both changes in participation and changes in duration conditional on participation. There are two important things to note about the measures of mothers' leave duration. First, mothers who are still on leave at the time of survey offer responses about planned leave duration while mothers who have returned to work report their completed leave duration. There is therefore concern that mothers may report planned duration that is either shorter or longer than the actual length of leave the parent ends up taking. However, since the EICS only covers mothers who have an infant under a year old, limiting our sample to mothers who have already returned to work would lead to the systematic over-representation of mothers who took shorter leaves, skewing the distribution of leave durations to the left. Consequently, I treat duration of leave to be length of completed leave for those who have returned, and length of planned leave for mothers still on leave. Second, the EICS survey asks new mothers about the duration of all

[8] I exclude single parents for three reasons. First, given the more limited financial resources of single parents, they are likely to respond differently to changes in the generosity of benefits than their partnered counterparts. Second, since they have no partner to share the gender-neutral parental leave with, there is no consideration of allocation decisions, which is an important component of this analysis. Third, there is concern that their behavior may be influenced by other policy changes that occurred in that period, such as enhancements of the National Child Benefit that particularly targeted lower-income single parents. Small sample sizes preclude a separate analysis of single mothers as well.

leave (not specifically paid parental leave) taken by them, and could capture unpaid leave or paid sick or vacation leave mothers take in lieu of paid parental leave. This means that the EICS measures of mothers' leave will represent the higher bound for the duration of paid parental leave taken by mothers. However, given the generous benefits available during paid parental leave and the lack of stigma to maternal leave-taking, mothers are unlikely to use other kinds of leave except to supplement paid parental leave once they have exhausted their weeks of benefits - and as mentioned earlier, the majority of families do not exhaust their total allowed weeks of benefits. For the minority of families that do, I assume that mothers use paid parental leave for the first X weeks of leave they report, where X is the family's total allowed minus any weeks reportedly consumed by the father. The measures of fathers' leave duration refer specifically to the number of weeks of paid parental leave that the mother reports that her spouse has claimed or plans to claim.

1.5 The Immediate Impact of QPIP

1.5.1 Regression Discontinuity Method

To evaluate the immediate impact of QPIP at the point that it was introduced, I adopt a sharp regression discontinuity (RD) design. Since the reform was introduced on 1st January 2006 with no gradual phase-in period, this provides a sharp cutoff after which a birth was eligible for QPIP. Moreover, there was limited certainty about the timing or the details of the reform until only a few months prior to its implementation.[9] The final details of QPIP,

[9] The idea of QPIP was discussed several years before the program came into place, but there were several bottlenecks in the policy process that prevented the program from being implemented. In June 2000, Quebec introduced legislation to establish its own parental leave program and in 2001 the Quebec National Assembly passed an Act that led to the development of a plan for Quebec's own program. However, the implementation of this legislation stalled because the federal government would not agree on the funds that

such as benefit amounts and the date of implementation, were only announced in mid-2005. Given that it takes some time to conceive a baby, it is reasonable to think that parents who gave birth around the cutoff were already pregnant at the time of announcement. Therefore, whether a birth occurred a few days prior to rather than a few days after January 1st 2006 was essentially random, allowing me to cleanly identify the local mean impact of QPIP through a regression discontinuity framework.

For each mother I have information on the year and month of birth of her youngest child, and the running variable for the RD is the distance in months from the cutoff date.[10] The model for each outcome is given by

$$Y_i = f(m_i) + \beta(m_i \geq Jan2006), \tag{1.1}$$

where $Y_{i,t}$ represents the outcome of mother i and m_i is the running variable which is the distance between the birth month and the cutoff of January 2006. β, the parameter of interest, represents the local mean impact of QPIP at the moment it was introduced. $f(m_i)$ is an unknown continuous function of the month of birth. I assume a flexible form for $f(m_i)$ and estimate it non-parametrically. I estimate equation 1.1 using local linear regressions (LLR) as Hahn et al. (2001) show that LLR performs better than kernel estimations at avoiding the boundary problem and obtaining a higher order of convergence at boundary points. The choice of bandwidth, i.e. the time window around the reform, is important since it determines the smoothing of the data and there is a tradeoff between variance and bias when choosing the optimal bandwidth. I select the bandwidth using the plug-

the Quebec government would be able to keep in order to finance its own program. In an effort to force the federal government to act, the Quebec government asked the Quebec Court of Appeal to rule on the constitutionality of the EI provisions on maternity and parental benefits. Only once the court ruled that the Employment Insurance Act regarding maternity and parental benefits encroached on provincial jurisdiction and exceeded the powers of the Canadian Parliament, did negotiations begin between the two governments in 2004. It was not until the middle of 2005, more than four years after the initial act regarding the program had been passed, that news emerged that the two governments had finally reached an agreement. QPIP officially came into place on the 1st of January 2006.

[10]The reported analyses use a bin width of one month. I also tried to widen the bins to 2-month intervals but rejected this method using an F-test since it over smoothed the data.

in method proposed by Imbens and Kalyanaraman (2009).[11] I provide the White (1980) heteroskedastic-consistent estimates of OLS standard errors, and in some specifications I allow for the clustering of standard errors within birthmonth, as suggested by Lee and Card (2008).

To confirm internal validity I verify that the pre-cutoff and post-cutoff group are balanced in characteristics. Table 1.2 presents results from regression discontinuity analyses on the personal and educational characteristics of households in the EICS Sample, showing no statistically significant discontinuities at the cutoff. Related RD graphs can be seen in Appendix Figures A.1 and A.2 . However, even though the discontinuities in sample characteristics are not statistically significant, a few are large enough to be economically significant, and therefore warrant concern. Accordingly, I present results from RD specifications with and without controls for these characteristics, to check that the program effect is not being confounded by changes in sample composition.

1.5.2 Regression Discontinuity Results

Table 1.3 presents results from regression discontinuity analyses to identify the the immediate effect that QPIP had on parents' leave participation rates and duration. Panel I shows results for Quebec in a simple RD specification that does not control for any personal or household characteristics. Column 1 of Panel I reports that the introduction of QPIP in January 2006 is associated with a jump of 53.6 percentage points in the probability that a father claims parental leave benefits in Quebec. This point estimate is highly statistically and economically significant, representing more than 250% of the pre-reform participation rate

[11] In Appendix Table A.1 I show my results are robust to the use of a parametric regression discontinuity analysis instead of a non-parametric approach.

of 21.3%. Column 2 indicates that QPIP resulted in a jump of 3.088 weeks in fathers' leave duration. This is also an economically and statistically significant effect, representing a 150% increase from the pre-reform average of 2.011 weeks. For mothers' leave participation, the RD detects a jump of 15 percentage points, but the estimate is not statistically significant. It appears there is no immediate jump in mothers' leave duration at the cutoff. The inclusion of personal and household characteristics in the RD analysis do not affect the results, as can be seen in Panel II. This is reassuring, confirming that the program effect of QPIP detected by the RD is not being biased by sharp changes in sample composition. Panel III presents RD analyses where the standard errors have been clustered at the level of the monh of birth, and the only notable finding is that under this specification the jump in mothers' participation rates becomes statistically significant. For comparison purposes, Panel IV presents RD results for the control group of other provinces and show that there was no change in parents' behavior around the time of the reform in provinces that were not treated.

Figure 1.3 provides visual support for these results, graphing the local polynomial for each of the simple RD analyses. Clear discontinuities can be seen in both fathers' participation rates and leave duration at the cutoff. Mothers' leave participation does jump at the cutoff but falls back down again. Figure 1.3 confirms there is no discontinuity in mothers' leave duration at the cutoff.

In aggregate, the RD results show that mothers' leave duration did not change while fathers' leave duration shot up. Furthermore, the increase in fathers' participation of 53 percentage points, is considerably more than the share of families for whom QPIP loosened a binding constraint by extending total family leave. This evidence is therefore consistent with an intra-household flypaper effect, whereby the daddy quota induced participation from fathers even in families where the new 'daddy weeks' were essentially fungible.

1.5.3 Threats to Identification

The identification strategy in a sharp Regression Discontinuity framework depends crucially on the assignment to treatment being based on an exogenous measure. Since in this case assignment is based on the month of birth, clean identification requires that the timing of pregnancies and births is exogenous to the introduction of QPIP. Several studies have shown that decisions of fertility and timing of birth can respond to financial incentives (Gans and Leigh, 2009; Tamm, 2013). This leads to concerns that citizens may have known about QPIP sufficiently in advance and in detail in order to time their births so that they could utilize the new program. However, I present several pieces of supportive evidence that the strategic manipulation of births is not a significant concern confounding my estimates. First, details about the date and features of the reform were not officially announced until only a few months prior to its implementation. [12] There were relatively few searches for the program until January 2006 when QPIP came into place, consistent with the idea that details of QPIP were not commonly known sufficiently in advance of 2006 such that parents could plan their pregnancies accordingly.

Second, it is necessary to check whether residents of Quebec who were already pregnant when they learned of QPIP may have delayed their births until after January 2006 in order to be eligible for QPIP. Since RD analyses identify a jump at the cutoff and it is naturally infeasible to delay a birth by more than a few days, this is equivalent to checking that our RD estimates are not biased by pregnant women who were originally due in late December who may have been able to delay the delivery by a few days in order to qualify for QPIP instead of EI. As a check against this, I drop all observations in the one month surrounding the reform, and re-estimate the RD on this trimmed window to check how sensitive my

[12] Appendix Figure A.3 presents a 'Google Trends' graph tracking searches for the word 'QPIP' around the time of the reform.Google Trends Searches using the full English and French names of the program present similar patterns.

results are to the exclusion of observations in the immediate vicinity of the cutoff (Barreca et al., 2011). Results from this 'trimmed' RD (shown in Appendix A.2) provide consistent estimates, with the exception that the effect on fathers' leave duration appears smaller.

1.6 The Average Treatment Effects of QPIP

1.6.1 Difference-in-differences Method

While the RD provides a clean estimate of the local mean impact of QPIP at the point at which it was introduced, it tells us nothing about whether QPIP continued to have an effect in the months and years to follow. To investigate the average treatment effect of QPIP over the period it has been in place, I use a longer span of data from 2002 to 2010, and employ a difference-in-differences method which exploits variation over provinces and time. It should be noted that another advantage of the difference-in-differences method is that there is less concern that it is biased by the manipulation of births around the cutoff. I estimate: I estimate:

$$Y_{ijt} = \alpha + \beta I[j = Quebec] * I[t >= 2006] + \theta I[t >= 2006] + \phi Z_{ijt} + \lambda_j + \delta_t + \epsilon_{ijt} \quad (1.2)$$

where subscript i denotes the individual, subscript j denotes province and subscript t denotes the year of last birth. Y_{ijt} therefore represents the outcome of mother i observed in province j who gave birth in year t. As outcomes, I explore whether the parent claims parental leave benefits and the duration of their actual or planned leave.[13] $I[t >= 2006]$ is an indicator variable taking the value 1 if the birth-year t is 2006 or greater, i.e., if the observation is from

[13] Mothers were asked about spouse's leave participation in all years, but about spouse's leave duration only in survey years 2004 and later. This does not affect my regression discontinuity framework because it only uses observations in a narrow window around the reform. For difference-in-differences, I conduct robustness checks using a uniform sample from 2004-2010 and find consistent results (Appendix Table A.5).

the post-reform period. The coefficient θ represents the change in the value of the outcome that is shared by all provinces. The term $I[j = Quebec] * I[t >= 2006]$ takes the value of 1 if the individual lives in Quebec and gave birth in a post-reform year, and otherwise takes the value 0. The coefficient β therefore represents the DD estimate of primary interest as it captures the change in the value of the outcome post-reform that is unique to Quebec. Under the assumption that no other policy changes were enacted to affect it, β represents QPIP's average treatment effect. λ_j and δ_t denote the fixed province and year effects. It should be noted that I do not control for all province-year interactions, but instead collapse them into the term $I[j = Quebec] * I[t >= 2006]$. [14]

The term Z_{ijt} is a vector of personal characteristics including age, education, legal marital status and immigrant status as well as household characteristics such as family size, number of children aged 0-1 and 1-5 and 6-17. Including these as regressors controls for changes in group composition. ϵ_{ijt} is the error term. I calculate cluster-robust standard errors that generalize the White (1980) heteroskedastic-consistent estimates of OLS standard errors to the clustered setting in order to account for possible heteroskedasticity and within-province dependence of standard errors, which are particularly a concern in difference-in-difference estimations since the regressor of interest is highly correlated within clusters (Bertrand et al., 2004). However, the small number of province-level clusters available in my sample leads to concerns regarding statistical inference since asymptotic tests have been shown to over-reject with too few clusters. Accordingly, I use the wild bootstrap-t procedures suggested by Cameron et al. (2008) to provide asymptotic refinement of standard errors.[15] All analyses are conducted using ordinary least squares regressions despite the binary nature of some of

[14] I do not include controls for province-specific time trends as these would be highly collinear with the program effect of QPIP. As one would expect, in supplementary regressions I confirm that the inclusion of a Quebec-specific time trend absorbs some of the program effects, leading to smaller but consistent point estimates of QPIP's impact on leave behavior.

[15] It is possible , though rare, for wild-bootstrapped errors to be smaller than regular standard errors in some cases with very few cluster groups. However, I confirm this was never the case for any regression in my analysis, and I only report the larger, wild-bootstrapped, standard errors.

the indicators because they resulted in very similar estimates as those from logit estimates.

It is important to discuss the assumptions under which this difference-in-differences identification strategy is valid, offering estimates of the true program effect. The first assumption is that no other programs or laws were enacted in Quebec at the same time which may have affected our outcome, such that the coefficient on $Quebec * Post - Reform$ may pick up effects of those other events instead. Notably, Quebec has a publicly subsidized childcare system while the rest of Canada does not. I verify that no policy changes were made to this program around the years of the QPIP reform. Further, I am careful to exclude data from before 2001, when the last expansion of the childcare program occurred. I also verify that Quebec did not make changes to child tax benefits or supplements that affect my sample around the time of the reform. The second necessary assumption of the difference-in-differences identification strategy is that of 'parallel trends' between the treatment and control group. That is, that the two groups should ideally have experienced similar trends prior to the introduction of the program, such that the control group offers a good proxy for the rate at which the outcome may have changed in the treatment province absent the treatment. I verify that this is the case for Quebec and other Canadian provinces in Figure 1.4.

Table 1.4 presents mean sample characteristics for the full 2002-2010 EICS sample as well as differences-in-means between the treatment and control groups over time. There are four difference-in-differences in characteristics which merit mention. First, the average age of new mothers grew more in Quebec than in other provinces, with a difference of 0.909 years. Second, the proportion of mothers that are legally married also grew more in Quebec than in other provinces. Third, the education levels of new parents changed more in Quebec than it did in other provinces, with an increase in mothers who have a high school education or less, and a decrease in mothers whose spouses only have a high school education or less. Lastly, though the difference is not statistically significant, the increase in proportion of immigrant mothers in Quebec by 6 percentage points may be economically significant. Interestingly, it should be noted that the increase in older, married, foreign-born couples with bigger husband-wife education differentials should be correlated with more traditional beliefs about gender roles, biasing me against finding more equal sharing of parental leave responsibilities. Nevertheless, I explore the issue of changing sample characteristics in two main ways. First, to account for compositional changes, I present results from DD estimations with and without controlling for such personal characteristics as age, education and immigrant status of mother and spouse, as well as household characteristics such as family size and number of children - and show that the point estimates are unaffected by these controls. Second, in Section 1.6.3 I discuss in detail whether these changes in sample composition could threaten identification, and I provide several robustness checks to allay any concerns.

1.6.2 Difference-in-differences Results

Table 1.5 presents results from difference-in-difference estimations of QPIP's impact on leave behavior. For fathers' leave outcomes, the DD estimates are very close to those obtained through the RD analysis. Column 1 of Panel I reports a significant program effect of 53.1 percentage points on fathers' participation rates and of 3.2 weeks in leave duration. For mothers' leave participation rates, the DD finds an average program effect of 12.1 percentage points, or a 16% increase from the pre-reform baseline. Column 4 explores the effect of QPIP on mothers' leave duration and reports a point estimate of 2.84 weeks, though the estimate is not statistically significant. Panel II presents results from DD estimates that also control for personal covariates and province and year-fixed effects, and we see that this does not affect the point estimates, though it affects inference in the case of fathers' leave duration.

Note that, just as in the case of the RD, the DD results show that fathers' leave duration responded more strongly to QPIP than did mothers' leave duration. This is especially so when we consider the effects in relative terms: since mothers' took an average of 43 weeks of leave prior to the reform, a program effect of even 3 weeks would represent an increase of approximately 7%, whereas for fathers a program effect of 3 weeks represents an increase of 150%. This larger response in fathers' leave-taking to the reform is consistent with QPIP's daddy quota producing an intra-household flypaper effect due to a reduction in stigma cost.

To examine this pattern in closer detail, Table 1.6 presents the impact of QPIP on the joint distribution of parental leave. Each cell represents a particular combination of mothers' and fathers' leave. The coefficients are estimated through difference-in-difference regressions where the outcome variable is an indicator for a family choosing that particular combination of mothers' and fathers' leave. The negative coefficients in Row A show that QPIP reduced the likelihood of any combination where the father took 0 weeks of leave. The positive coefficients in Row B show that QPIP increased the likelihood that the average father took between 1 and 5 weeks of leave, i.e. consumes from his quota. The coefficients in Column 5 are consistent with families responding to the relaxation of the total family leave constraint: when the mother is consuming a full year of leave, adding 5 daddy-only weeks makes it less likely that the father consumes no leave, and makes it more likely that Dad consumes between 1 to 5 of the newly available weeks of leave. However, the coefficients in Columns 3 and 4 find increases for fathers even in families not constrained by the cap, i.e., where mothers consumed less than a year of leave such that fathers always had weeks available to them.[16] That is, even in families where the father always had some amount of leave available to him, the introduction of QPIP made it less likely that father consumed no leave, and made

[16]It should be remembered that the EICS reports of maternal leave duration may include other kinds of leave (as well as the mandatory 2-week waiting period under EI), representing the higher bound for paid parental leave consumed by mothers. Families in Column 4 are not facing a binding constraint because even if the mother reported 11 months of leave, the father had at least 1 full month of paid parental leave available to him.

it more likely that he consumed some leave.

Even more telling, the majority of movement in this table brings fathers from Row A (consuming no leave) to row B (consuming from their daddy quota), but not into Row C (consuming more than their quota). In fact, the average father in post-reform Quebec consumes 5.23 weeks of leave, almost exactly his quota. This means that in families that have slack, QPIP increases fathers' consumption but only by the amount of his new quota, meaning some weeks of leave remain unconsumed. For example, consider a family in Column 4 where the mother takes 10-11 months of leave, meaning the father has at least 1-2 months of paid leave available to him under the EI program. Table 1.6 shows that the introduction of QPIP, which adds 5 extra 'daddy weeks' to the family's budget, makes it less likely that the father does not participate and makes it more likely that the father participates but only consumes 1-5 weeks of leave - even though he now has 2-3 months of leave available to him. The overall pattern is that QPIP made fathers more willing to participate in parental leave, but only to consume from their 'daddy quota' allocation. This is highly suggestive evidence that labeling 5 weeks of leave to be 'daddy-only' makes these weeks "stick" to fathers - a flypaper effect.

It should be noted that overall, even though fathers responded more strongly to QPIP than did mothers, the QPIP program did not eliminate gender differences in leave behavior. Even under QPIP, only 75% of fathers claimed parental leave benefits compared to 83% of mothers, and fathers on average took a little over 5 weeks of leave, in stark comparison to mothers' average 46.37 weeks of leave. In aggregate QPIP was able to shrink but not close the gap in parental leave participation between men and women.

1.6.3 Threats to Identification

In difference-in-differences frameworks, it is common to challenge the exogeneity of reforms by questioning whether policies are in fact endogenously implemented as a response to trends in the outcome in the first place. For example, one may wonder whether policymakers in Quebec instituted a daddy quota because they were concerned by falling participation rates among fathers. Figure 1.4 plots fathers' participation rates in the treatment and control provinces over time. It shows that fathers' participation rates were not falling or rising more quickly in Quebec than other provinces prior to the reform. This confirms that the key assumption of the difference-in-differences identification, that of parallel trends between treatment and control groups, is satisfied. Prior to 2006, both Quebec and other provinces experienced slightly increasing but parallel trends in fathers' participation. Thus, the control group offers a good proxy for the trajectory that Quebec would have followed absent the treatment, such that the program effects from DD regressions offer a good estimate of the level shift in participation rates due to the introduction of QPIP.

A second possibility is that even though details of QPIP were not available until mid-2005, the basic idea had been proposed in 2001 and citizens may have heard that a reform was being discussed that would offer generous incentives for fathers to participate. It is possible that couples who were particularly keen to have fathers take leave chose to delay pregnancy until the new program was in place, building pent-up demand that could only be released in early 2006. This seems unlikely as this would suggest families may have been willing to postpone a pregnancy by many years (five years passed between the program proposal and final announcement of implementation) in order to gain a few weeks of 'daddy' leave. Nevertheless, I check this by conducting an event study analysis to see how the program effect differed in the years following the reform. If my program effects were driven by pent-up demand which was released in 2006, we would expect fathers' uptake to jump up in 2006 but then to fall back down in later years once this pent-up demand was relieved. Figure 1.5 presents results from an event-study analysis of the reform and shows that the program effect did not fall over time but remained constant or may have grown stronger as years passed.

Another threat to identification is that of selective migration, i.e., that people may have moved to Quebec specifically to give birth there and avail themselves of the generous benefits. However, the Population Estimates Program at Statistics Canada reports that Quebec experienced negative net migration every year over the decade in question, and moreover, that the numbers of out-migrants actually increased over the years surrounding the reform, i.e. from 2004 to 2008 (Milan, 2011).

Lastly, I consider the possibility that changes in the characteristics of the treatment group over time may be driving my program effects. I address these concerns in two ways. First, I conduct a summary test of composition bias by regressing leave outcomes on personal and household characteristics using the pre- treatment sample, obtaining predicted outcomes, and then running a simple difference-in-difference regression on the predicted outcomes (See Appendix Table A.3). The results for fathers' outcomes strongly support a causal interpretation of QPIP, since the reform has no 'program impact' on the predicted participation rates or durations. In the case of mothers' outcomes, there are small negative 'effects', suggesting that if anything, changes in sample composition should have biased us against finding a program effect of increased mothers' leave participation and duration. As a further robustness check, I conduct difference-in-difference analyses on subsamples of the data based on these characteristics, e.g., subsamples of non-immigrant mothers or young mothers etc (Appendix Table A.4). The results for each subsample are very consistent to the results for the main sample, offering little reason to believe that QPIP's average program effects are largely driven by heterogeneous effects on these sub-groups that may have increased or decreased in prominence in the sample over time.

1.7 Conclusion

This paper provides the first comprehensive study of the short-run effects of the QPIP reform and offers an important contribution to the literature on parental leave. It is the first to provide causal evidence that daddy quotas may influence behavior even when they do not relax a binding constraint, suggesting that the 'daddy-only' label produces an intra-household flypaper effect that makes leave stick to fathers. The results of this study thus have important policy implications. First, they suggest that 'daddy-only' quotas may help fathers overcome such barriers to taking leave as social stigma and perceived professional penalties.

Table 1.1: Details of Parental Leave Programs in Canada

	Employment Insurance	QPIP Basic Plan	QPIP Special Plan
Eligibility Requirement	600 hours of insurable employment	2000CAD of insurable earnings	2000CAD of insurable earnings
Basic Replacement Rate	55%	70% for all maternity, & paternity leave, first seven weeks of parental leave and 55% thereafter	75%
Max insurable earnings	39,000CAD	57,000CAD	57,000CAD
Waiting Period	2 weeks	None	None
Duration	Total 50 weeks = 15 weeks maternity leave + 35 weeks parental leave + no paternity leave	Total 55 weeks = 18 weeks maternity leave + 32 weeks parental leave + 5 weeks paternity leave	Total 40 weeks = 15 weeks maternity leave + 25 weeks parental leave + 3 weeks paternity leave

Source: Table constructed by author using information from the Digest of Benefit Entitlement Principles, available at http://www.servicecanada.gc.ca/eng/ei/digest/chp12_appendix.shtml. For features that may change on a yearly basis, such as the amount of maximum insurable earnings, figures provided are for 2006.

Table 1.2: Regression Discontinuities in Personal & Educational Characteristics

	Mother's Age	Fathers' Age	Children aged 0-1	Children aged 1-5	Family Size	Immigrant
RD Estimate	0.138	0.624	0.012	-0.130	-0.324	0.110
	[0.894]	[0.610]	[0.340]	[0.432]	[0.118]	[0.410]
Bandwidth (months)	16.69	17.14	4.06	9.65	9.80	8.147

	Mother's Educ $\leq$ High School	Mother's Educ = Some College	Mother's Educ = College	Fathers' Educ $\leq$ High School	Fathers' Educ = Some College	Fathers' Educ = College
RD Estimate	0.014	-0.100	0.091	-0.101	0.222	-0.123
	[0.913]	[0.540]	[0.484]	[0.264]	[0.191]	[0.464]
Bandwidth (months)	8.46	8.94	8.80	8.28	8.94	8.76

Notes: Table presents results from non-parametric RD analysis of Quebec, using local linear regressions to detect discontinuities in personal and household characteristics between parents who experienced a birth before versus after the cutoff of January 2006. Heteroskedasticity-robust p-values are presented in square brackets. The optimal bandwidth was selected using the plug-in procedure suggested by Imbens and Kalyanaraman (2009). Sample spans 2004-2007 of the EICS data and comprises mothers aged 18-40 in cohabitating or married relationships who have experienced a birth in the last year.

Table 1.3: Regression Discontinuities in Parents' Leave Behavior

OUTCOMES:	(1) Fathers' Participation Rates	(2) Fathers' Leave Duration (Weeks)	(3) Mothers' Participation Rates	(4) Mothers' Leave Duration (Weeks)
Baseline for Quebec (2002-2005 Average)	0.213	2.011	0.725	43.140
I. RD analysis of Quebec				
Jump at Cutoff	0.536***	3.088*	0.155	-0.277
	[0.00]	[0.06]	[0.18]	[0.93]
Bandwidth (months)	8.692	18.968	8.568	24.874
II. RD analysis of Quebec, including personal covariates				
Jump at Cutoff	0.531***	3.126**	0.188	0.930
	[0.00]	[0.04]	[0.15]	[0.93]
Bandwidth (months)	8.791	18.968	8.568	24.874
III. RD analysis of Quebec, including personal covariates and clustered errors				
Jump at Cutoff	0.530***	3.125**	0.187**	0.930
	[0.00]	[0.01]	[0.01]	[0.78]
Bandwidth (months)	8.692	18.968	8.568	24.874
IV. RD analysis of control provinces				
Jump at cutoff	-0.006	-1.332	-0.008	-0.207
	[0.92]	[0.11]	[0.93]	[0.93]
Bandwidth (months)	6.143	14.733	7.11	20.669

*** p<0.01, ** p<0.05, * p<0.1, Heteroskedasticity-robust p-values in parentheses

Notes: Table presents results from non-parametric RD analysis of Quebec, using local linear regressions to detect discontinuities in leave outcomes between parents who experienced a birth before versus after the cutoff of January 2006. The running variable is month of birth, with bin size of 1 month each. Sample spans 2004-2007 of the EICS data and comprises mothers aged 18-40 in cohabitating or married relationships who have experienced a birth in the last year. Optimal Bandwidth chosen by the plug-in procedure suggested by Imbens and Kalyanaraman (2009). When errors are clustered, they are done so at the level of the month of birth (assignment variable).

Table 1.4: Sample Means in Household Characteristics in EICS Data

	Control 2002-2005	Control 2006-2010	Quebec 2002-2005	Quebec 2006-2010	Difference in Differences
Age of Mother	30.307	30.456	29.071	30.134	0.909***
Age of Spouse/Partner	32.732	32.851	31.900	32.586	0.552
Legally Married	0.862	0.846	0.366	0.389	0.040**
Immigrant	0.213	0.204	0.113	0.169	0.063
Family Size	3.784	3.830	3.711	3.760	0.006
Number of children aged 0-1	1.013	1.019	1.011	1.015	0.001
Number of children aged 1-5	0.526	0.582	0.497	0.528	-0.021
Number of children aged 6-17	0.243	0.251	0.205	0.249	0.036
Mother has high school degree or less	0.244	0.162	0.163	0.162	0.034***
Mother has some college	0.415	0.417	0.552	0.466	-0.087*
Mother has college degree	0.339	0.371	0.284	0.371	0.053
Father has high school degree or less	0.279	0.261	0.252	0.176	-0.060*
Father has some college	0.434	0.432	.512	0.515	-0.005
Father has college degree	0.285	0.306	0.235	0.308	0.054

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Notes: Sample spans 2002-2010 of the EICS data and comprises mothers aged 18-40 in cohabitating or married relationships who have experienced a birth in the last year. Difference-in-differences are identified across provinces and time.

Table 1.5: Difference-in-Differences in Parents' Leave Outcomes

	(1)	(2)	(3)	(4)
OUTCOMES:	Fathers' Participation Rates	Fathers' Leave Duration (Weeks)	Mothers' Participation Rates	Mothers' Leave Duration (Weeks)
Baseline for Quebec (2002-2005 Average)	0.213	2.011	0.725	43.140
I. Simple D-in-D Specification				
Quebec * Post-Reform	0.531***	3.225*	0.121*	2.845
	[0.00]	[0.08]	[0.07]	[0.25]
N	8907	7157	8907	6172
II. D-in-D Specification with Personal Controls and Province & Year- Fixed Effects				
Quebec * Post-Reform	0.527***	3.241	0.125**	2.765
	[0.00]	[0.15]	[0.05]	[0.14]
N	8905	7156	8905	6441

*** p<0.01, ** p<0.05, * p<0.1, Heteroskedasticity-robust province-clustered p-values in parentheses

Notes: Table presents difference-in-difference estimates of parents' leave behavior between Quebec and Other Provinces before and after the introduction of QPIP in 2006. Sample spans 2002-2010 of the EICS data and comprises mothers aged 18-40 in cohabitating or married relationships who have experienced a birth in the last year. Heteroskedasticity-robust p-values, clustered at the province level and calculated using wild bootstrap procedures, are presented in brackets. Statistics Canada's cell-size requirements for data disclosure prevent me from slicing the data more finely in terms of fathers' leave duration.

Table 1.6: Program Effect of QPIP on Joint Distribution of Parental Leave

	(1) Mother takes 0 months of leave	(2) Mother takes 1-5 months of leave	(3) Mother takes 6-9 months of leave	(4) Mother takes 10-11 months of leave	(5) Mother takes 12+ months of leave
(A)					
Father takes 0 weeks of leave	0.002	-0.049	-0.070	-0.128**	-0.336***
	[0.95]	[0.12]	[0.71]	[0.03]	[0.00]
(B)					
Father takes 1-5 weeks of leave	0.003***	-0.003***	0.089***	0.211***	0.258***
	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]
(C)					
Father takes 6+ weeks of leave	-0.001	-0.004	0.017**	0.021	-0.008
	[0.57]	[0.21]	[0.00]	[0.29]	[0.89]

*** p<0.01, ** p<0.05, * p<0.1, Heteroskedasticity-robust province-clustered p-values in parentheses

Notes: Table shows the difference-in-differences across provinces and time in the likelihood of various combinations of mothers and fathers' leave duration. Columns represent mothers' leave duration in months and rows represent fathers' leave duration in weeks; hence each cell represents a different outcome which is an indicator for a family choosing a particular combination of mothers' and fathers' leave durations. Sample spans 2002-2010 of the EICS data and comprises mothers aged 18-40 in cohabitating or married relationships who have experienced a birth in the last year. All regressions include controls for personal and household characteristics and province- and year-fixed effects.

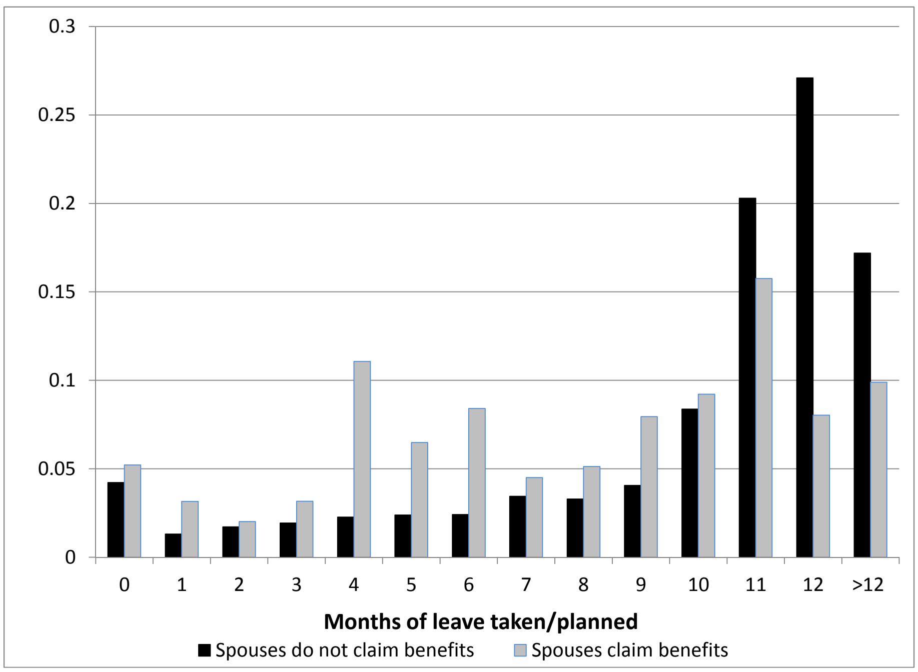


Figure 1.1: P.D.F of Maternity Leave Duration in Quebec, 2002-2005

Source: Graph of probability density function created by author using raw EICS data on mothers' leave duration in Quebec for survey years 2002-2005.

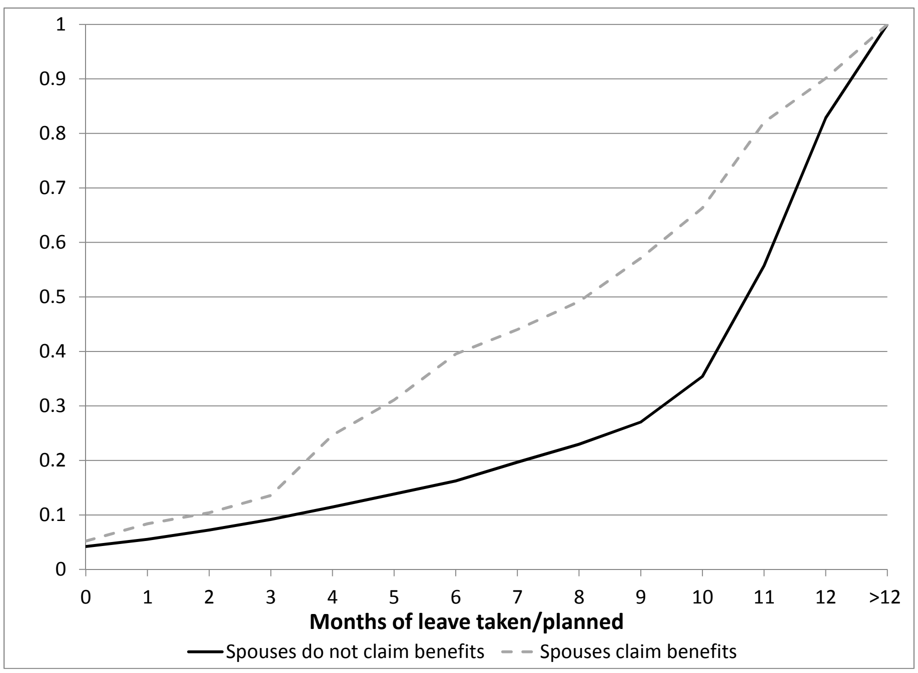


Figure 1.2: C.D.F of Maternity Leave Duration in Quebec, 2002-2005

Source: Graph of cumulative distribution function created by author using raw EICS data on mothers' leave duration in Quebec for survey years 2002-2005.

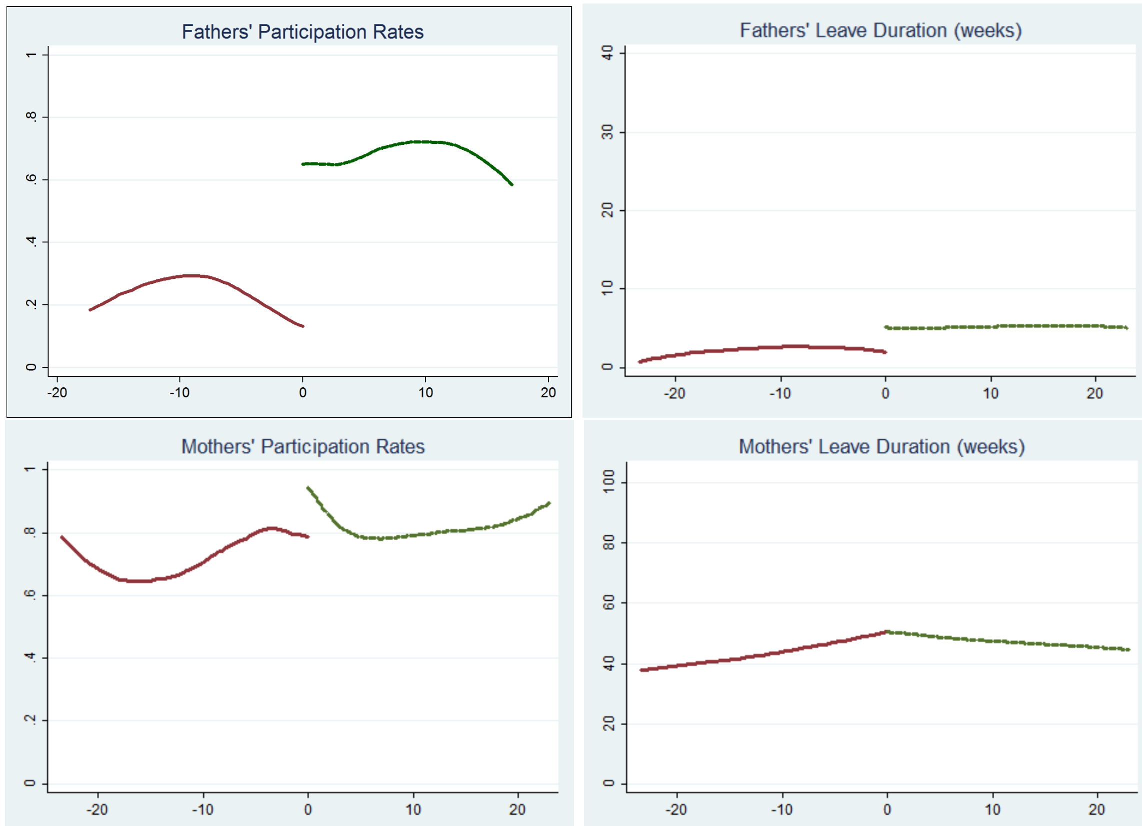


Figure 1.3: Discontinuities in Parents' Leave Outcomes

Source: Graphs created by STATA program to plot local polynomials from non-parametric local linear regressions that identify jumps in outcomes at the cutoff (January 2006), using EICS data for Quebec. Graphs correspond to the regression discontinuity results presented in Panel I of Table 1.3. Statistics Canada's disclosure requirements for restricted-access Master data prevent the plotting of raw data in these graphs.

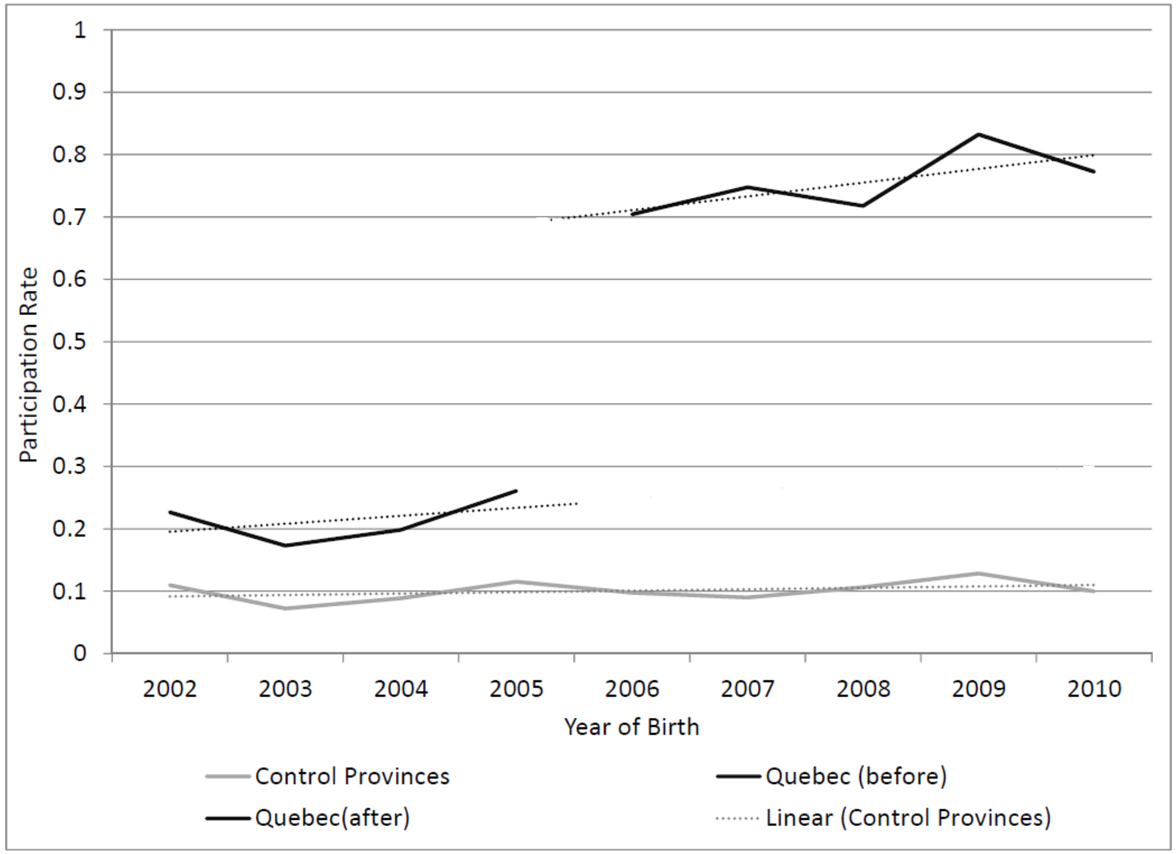


Figure 1.4: Trends in Fathers' Leave Participation

Source: Graph constructed by author using Employment Insurance Coverage Survey data from 2002-2010. Graph plots average participation rates for fathers in Quebec and other provinces, as well as corresponding linear trend lines.

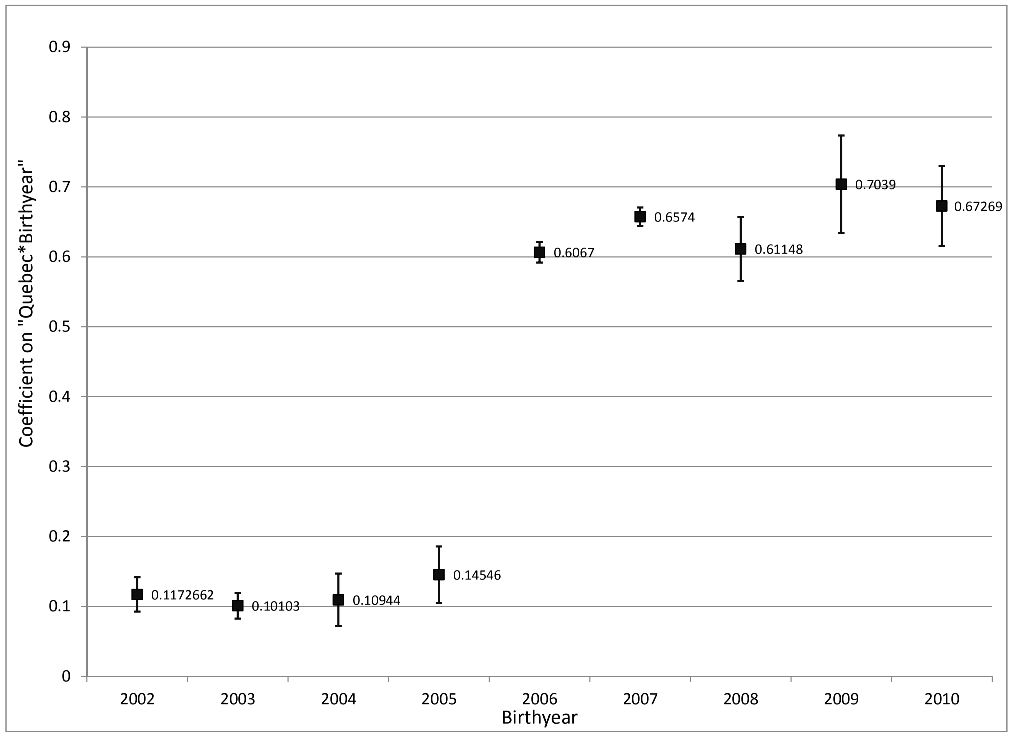


Figure 1.5: Event Study of Fathers' Leave Participation

Source: Graph constructed by author. Event study was conducted by regressing the outcome on indicators for birth year as well as interaction terms between an indicator for Quebec and indicators for birth year Graph plots coefficients from the interaction terms of Quebec*birth year, and thus presents the difference between the average father in Quebec and that in other provinces for every year. Vertical lines show 95% confidence intervals.

CHAPTER 2

MERGING SEPARATE SPHERES: DOES PATERNITY LEAVE REDUCE SEX SPECIALIZATION IN THE LONG RUN?

2.1 Introduction

Despite a dramatic reduction in the gender gap in labor force participation and wages, a large and persistent gap remains in the realm of care work. Mothers are much more likely than fathers to take parental leave in the first months of a child's life, which has the potential to hurt mothers' careers.[1] Consistent with this pattern, a cross-country analysis by the OECD reports that the length of paid parental leave available is correlated with a higher pay differential by gender (OECD, 2012). Mothers also perform considerably more unpaid work in the home than do fathers (Hochschild and Machung, 1989; Blair and Lichter, 1991; Bianchi, 2011, 2012). In turn, this disproportionate amount of housework done by women, particularly time-inflexible and routine work, has been shown to contribute to the gender pay gap (Hersch and Stratton, 2002). Becker (1981) used the idea of comparative advantage to argue that the traditional division of household labor may be efficient due to women's lower market wages and biological advantages in care-giving such as the ability to breastfeed. However, recent decades have witnessed considerable growth in women's wages and the advent of technology to minimize biological differences, without a corresponding reduction of the same magnitude in household sex specialization. These patterns suggest that sticky social norms about gender roles may be perpetuating higher levels of specialization than necessarily efficient for the household.[2]

[1]Several studies note that the provision of extended maternity leave delays women's return to work (Schönberg and Ludsteck, 2007; Lalive and Zweimller, 2009) and lowers the probability of upward occupational moves (Evertsson and Duvander, 2011). In addition, Blau and Kahn (2013) find that generous parental leave policies seem to encourage women's part-time work and employment in lower level positions.

[2]Sex specialization may be undesirable for reasons beyond inefficiency, as argued by Blau et al. (2014). Since it reduces a woman's wages and increases her marriage-specific investments, female specialization in

Research has consistently shown that the birth of a child marks the beginning of a more gendered division of labor (Baxter et al., 2005; Shelton, 2000) because mothers and fathers have different initial experiences of parenting, which can serve to reinforce traditional gender roles. New mothers often absent themselves from the labor market for varying lengths of time, which erodes their labor market positions and increases investments in household work, particularly routine and time-inflexible chores. Fathers instead maintain or even strengthen their ties to the workforce in the first years of a child's life (Sanchez and Thomson, 1997), taking on greater financial responsibilities and reaffirming their roles as breadwinners. So it can be argued that if mothers and fathers were to experience the transition to parenthood in more similar ways it may lead to a more equitable division of labor within the household in the long run. One means by which gender differences in the initial parenting experience could be reduced is through fathers' participation in parental leave. Although the primary goal of most parental leave policies remains the protection of the health of the child and mother, there has been a recent trend in policy-making, beginning in Scandinavia but now catching on in other countries, towards promoting gender equality by incentivizing fathers to take parental leave.

It was with this goal in mind that policymakers in Quebec decided to increase compensation for paid parental leave and create a 'daddy quota' for fathers. From 2001 to 2005, eligible parents in all Canadian provinces could claim parental leave benefits from the government through the Employment Insurance (EI) Program. In 2006, Quebec left the EI system and established the Regime Quebecois D'assurance Parentale or the Quebec Parental Insurance Plan (QPIP). This new scheme offered easier eligibility criteria, increased income replacement, and established a 5-week 'daddy quota' of leave for fathers (Doucet et al., 2010). Prior to the reform, fathers only had access to 'shared' parental leave with their spouses,

care work lowers her bargaining power. The costs of interdependence are also disproportionately borne by women, as in the case of the displaced homemaker who upon being widowed or divorced must support herself from a weakened market position.

and leave-takers were compensated with a little over half their wages up to a strict cap so that household incomes were hit hard when fathers took leave. Consequently, fathers' leave participation in Quebec never exceeded 22% prior to 2005. The QPIP reform was extremely successful in boosting fathers' involvement in parental leave: Chapter 1 of this dissertation reported that QPIP was associated with a sharp increase in father's participation rates of 53 percentage points and an increase in fathers' leave duration by 3 weeks. In this study, I exploit the exogenous variation in fathers' leave participation created by QPIP to investigate the causal impact of paternity leave on household sex specialization in the long run.

This study makes several important contributions to the literature. First, this study offers the first comprehensive examination of the long-term causal effects of paternity leave. By simultaneously investigating multiple outcomes related to gender equality, such as the division of household labor, patterns of housework and spouses' time in the workplace, I offer unique evidence of the impact on overall household dynamics. Second, since only the province of Quebec deviated from the national policy, this study is unique in utilizing regional variation in policy rather than a nationwide change in policy. This results in improved study design: by using other Canadian provinces where policy did not change as a natural control group, I can provide causal estimates that are robust to various trends. Third, this is the first causal study of paternity leave to use data from time-diaries, increasingly considered the gold standard for information about non-market production (Sevilla, 2014). Consequently, this analysis explores more detailed measures of parent's daily behavior, and is able to glean a more nuanced and accurate insight into the long-term effects of paternity leave on household division of labor where other studies have been unable to. Lastly, this is the first study of how this Canadian policy episode affected gendered patterns in market and non-market production. This is interesting because Canada offers a political and social context that is quite different from the previously-studied Scandinavian countries, since the latter have some of the most generous welfare provisions and family-friendly policies worldwide. In sum, this study contributes to the literature by utilizing better data and improved methodology in a new context, as well as by exploring the bigger picture of sex specialization across the breadth of parental responsibilities.

I use variation in exposure to QPIP as a proxy for exogenous variation in paternity leave, and study the impact on the division of labor within a household 1-3 years after it experienced a birth. I use time-diary data from the 2005 and 2010 rounds of the General Social Survey, and a difference-in-differences approach that exploits variation in exposure to QPIP across time, provinces and children's ages. I find strong evidence that by altering the initial experience of parental leave, QPIP had a large and persistent impact on gender dynamics within households. Exposure to QPIP moved households towards a dual-earner, dual-caregiver model wherein fathers and mothers contribute more equally to home and market production. I find that exposed mothers spent more time in paid work, more time physically at the workplace and were more likely to be full-time employed, compared to their counterparts who were not exposed. In the realm of non-market production, I find that exposure to QPIP increased both parents' contributions - although exposed fathers increased their time by more than exposed mothers. Specifically, exposed fathers spent more time in housework per day, while exposed mothers decreased their housework and spent more time in childcare instead. Moreover, exposed fathers spent more time physically at home while exposed mothers spent less time in the home. Overall, a clear pattern of reduced sex specialization emerges among exposed households. Taken together, these results suggest that small changes in the initial parenting experience can have lasting effects on parents' behavior. More broadly, my findings highlight that there need not be a trade-off between gender equality and parental time with children: paternity leave can distribute household responsibilities more equally and increase time investments in children.

The rest of this Chapter is organized as follows: Section 2.2 describes the background for the study, providing the theoretical context and a review of the literature. Section 2.3 discusses the research design, data and empirical methods used in this study. Section 2.4 presents the findings and discusses the results as well as possible confounding factors and limitations of the study. Section 2.5 concludes.

2.2 Background

2.2.1 Expected effects of QPIP on the Household Division of Labor

Paternity leave policies promote gender equality by intervening at a crucial time for renegotiating household work (Hook, 2010) and facilitating a re-allocation of parents' resources across professional and domestic spheres. They make fathers available for time-inflexible housework and childcare, enabling mothers to return to the workforce sooner and invest in their careers. Thus, reduced sex specialization during the period of paternity leave can be explained both by fathers' increased time availability for non-market work and mothers' increased bargaining power.[3] A main objective of this study, however, is to explore the *long-term* causal effects of paternity leave on the household division of labor, which no previous study has been able to establish. I seek to answer the following question: do the effects of paternity leave on sex specialization persist after the leave period, or does the household revert to traditional gender roles afterward? There are several channels through which paternity

[3]If bargaining power is proportional to an individual's contribution to the household income (as proposed by e.g. (Lundberg and Pollak, 1996)), fathers on leave have reduced earnings and thereby diminished ability to bargain away from doing unpleasant domestic chores, while their wives who have returned to work would have higher earnings and therefore bargaining power. This argument relies on the parental leave being compensated at less than 100% of usual earnings, which is the case in Canada and in many nations offering extensive paid leave to fathers.

leave may lead to a *permanent* reduction in sex specialization, as I argue below.

The first explanation builds on the theory of Becker (1981), in which a household uses productivity differentials that may differ across genders to determine an efficient allocation of resources. Since men earn higher market wages on average and women have some biological advantages in childcare, the theory of comparative advantage suggests that men allocate more time to market work while women take on more domestic responsibilities because it is efficient for the household. However, being on leave increases fathers' time in childcare and housework, especially time-inflexible tasks, in which they gain experience and competence. Fathers on leave undergo on-the-job training, which increases their domestic productivity and reduces differentials between their returns to non-market and market work. If fathers are penalized by employers for taking leave through lower wages or fewer promotions, this further reduces their productivity differential by lowering returns to market work.[4] Women whose husbands take paternity leave can return to work earlier and enjoy greater job continuity, increasing their returns to market work. As the ratio of the returns to market versus non-market work for mothers and fathers converge, this should lead to a within-family time allocation that is less sex-specialized, whereby fathers contribute more to unpaid work and mothers contribute more to market work.

Another mechanism through which paternity leave may influence behavior in the long-term is that of habit persistence in preferences. Individuals may have utility over different kinds of work that is non-separable over time. Under such a model, lifetime utility would take the form of $U(C) = \sum_{t=0}^{T} u(c_t - \alpha c_{t-1})$, where $u(.)$ is a concave utility function, c_t is consumption in period t and α denotes the intensity of habit formation. Due to the concavity of the utility function $u(.)$, we would then have that $u'(c_t) \leq 0$ and $u'(c_{t-1}) \geq 0$, that is,

[4] Over time, if social norms and expectations about gender roles change, we may also see men and women make different decisions about human capital investments as well as occupation and industry choices, which would have an impact on their returns to market work.

that marginal utility is decreasing in current consumption as per usual but is increasing in past consumption. Thus, an increase in current consumption, c_t, will lower the marginal utility of consumption in the current period but increase marginal utility in the next period. For example, paternity leave could increase a father's initial consumption of childcare, which increases the marginal utility of childcare in the next period, such that he demands more contact with his child even after the leave period ends. The same logic would hold if we modeled either childcare or housework as a 'bad' rather than a 'good'. For example, paternity leave may increase a father's participation in cooking, which lowers his marginal dis-utility from the task in future periods.

Paternity leave also may create a pattern of household behavior during the period of leave that is costly to reverse later. One potential cost of changing behavior after the leave period is that of learning. Parents who take leave simultaneously may divide up non-market tasks and each invest in task-specific human capital. After the leave period ends, it becomes costly for either parent to learn how to perform the other's designated task and to avoid this cost they may continue to share chores as they did while they were on leave. In addition, there may be utility costs associated with reversion. For example, the wives of men who take leave may enjoy the experience of committing to their careers while being supported by a helpful spouse at home, and they may perceive dis-utility from returning to traditional gender roles where their career is subordinate to their spouse's.[5]

Paternity leave should also limit the possibility of strategic shirking, since fathers cannot credibly claim to be incompetent in certain childcare and housework tasks any longer. Lastly, the public message promoting active fathering behind a daddy quota, as well as the actual experience of taking paternity leave, could influence the identity of fathers and their spouses

[5] In that case, under a unitary model of maximization reverting to the traditional division of labor would be sub-optimal for the household. Alternatively, under a non-cooperative model, since mothers benefit from the non-traditional division of labor they can use their improved bargaining power to enforce continuation of this behavior even after the leave period ends.

(Akerlof and Kranton, 2010).

2.2.2 Previous Research on the Long-Run Effects of Paternity Leave

Much of the extant research on the long-term effects of paternity leave has examined variation in actual leave-taking among fathers or cross-country variation in leave policies. Such studies have found that fathers who take leave are more involved in childcare (Haas, 1990; Brandth and Kvande, 1998; Haas and Hwang, 1999; Tanaka and Waldfogel, 2007; Nepomnyaschy and Waldfogel, 2007) and that the average father's time in childcare is higher in countries with generous paternity leave policies (Fuwa and Cohen, 2007; Sullivan et al., 2009; del Carmen Huerta et al., 2013; Boll et al., 2014). Although studies find no evidence of an association between paternity leave and fathers' average time in housework, some do find evidence consistent with increased male participation in time-inflexible and typically-female housework (Brandth and Kvande, 1998; Hook, 2006, 2010). Moreover, paternity leave is correlated with shorter work hours for fathers (Haas and Hwang, 1999; Duvander et al., 2010), and shorter career breaks, longer work hours and improved labor market positions for mothers (Brandth and Kvande, 1998; Pylkkänen and Smith, 2003). Taken together, these cross-sectional studies suggest that paternity leave is correlated with a less traditional division of labor within the household. However, these associations are vulnerable to endogeneity issues. Cross-country studies may be biased upwards by the omission of country-level variables such as institutional or normative contexts. Similarly, studies using cross-sectional variation in actual leave-taking among fathers cannot control for unobserved heterogeneity in preferences, beliefs, motivation, and workplace constraints. Thus, their findings can only be interpreted as informative associations rather than as causal estimates.

More recently a few studies have sought to identify causal effects of paternity leave by comparing the behavior of parents before and after a change in policy that led to a sudden increase in fathers' leave-taking, thus exploiting exogenous variation in leave-experience. Interestingly, these studies of 'natural experiments' have not been able to confirm the results from cross-sectional research. First, several studies fail to detect a significant causal impact of paternity leave on the distribution of childcare between parents (Kluve and Tamm, 2009; Rieck, 2012; Ekberg et al., 2013; Ugreninov, 2013). One study did report that paternity leave leads to more equal sharing of housework, but could only detect a significant effect for the chore of laundry (Kotsadam and Finseraas, 2012). Second, these studies are not consistent in their findings on the causal effect on parents' labor market outcomes. While some studies report that paternity leave reduces fathers' earnings (Johansson, 2010; Rege and Solli, 2013), others find no impact on fathers' earnings or work hours (Cools et al., 2011). Similarly, some studies find no causal effect on mothers' labor supply or earnings (Rege and Solli, 2013; Kotsadam et al., 2011), while others report that paternity leave leads to higher or lower maternal earnings (Johansson (2010) and Cools et al. (2011) respectively). Thus, the results from these quasi-experimental studies are not conclusive, but do confirm the suspicion that the relationship between paternity leave and parental behavior may not be as straightforward as the cross-sectional evidence suggests.

The inability of the quasi-experimental literature to reach a conclusive result may be explained by three critical shortcomings of the studies conducted thus far. First, the data on non-market production used in these studies is far from ideal. Some studies explore narrow measures of parental involvement, e.g., Ekberg et al. (2013) and Ugreninov (2013) use the share of sick days taken to care for ailing children as their measure of a parent's childcare work. Other studies use broader measures, but rely on data vulnerable to measurement error and reporting bias. For example, Kluve and Tamm (2009) ask parents 1.5 years after the birth to report the proportion of childcare performed by them during the first year of the child's life, even though respondents may incorrectly remember their contributions. Kotsadam and Finseraas (2012) use data from a survey that asks respondents, for example, "*Who does the chore of laundry in your household? - you? -your partner? -you share equally?*". The limited range of possible answers means these measures lack precision. More importantly, since these questions explicitly hint at evaluating gender relations these data are susceptible to response bias, wherein respondents purposefully understate or exaggerate their behavior to align with cultural norms about gender equality rather than reporting their true actions. Second, previous studies have focused on one or at most two dimensions of parental responsibility, and are therefore unable to identify substitutions between tasks. Some studies only considered outcomes for one parent (e.g. Kotsadam et al. (2011); Ugreninov (2013)) and so cannot capture the fact that mothers' and fathers' time may be complements or substitutes in household production. Third, the quasi-experimental studies to date exploited nation-wide changes in policy to compare fathers who experienced a birth before a reform to those who experienced a birth after. Analyses using only one period of observation (e.g. Kotsadam and Finseraas (2012)) thus necessarily compare fathers of older children to those of younger children, whose behaviors may differ inherently - and further, the differences may simply reflect cohort trends in parent's behavior. Studies with multiple periods of observation produce difference-in-difference estimates by comparing parents' behavior across children's ages and and time (e.g. Rege and Solli (2013)), but it can be argued that parents of older children are not an ideal comparison group due to the strong identifying assumption of parallel trends in parents' behavior across children's ages. Keeping these issues with the prior literature in mind, the present study is designed to make careful use of time-diary data that offer precise, unbiased, comprehensive measures of each parent's behavior, and to exploit variation across provinces and time in order to control for trends and provide clean identification of causal links.

2.3 Research Design and Data

2.3.1 The Natural Experiment

In order to identify a causal mechanism, I exploit variation in exposure to a policy that is positively correlated with fathers' participation in parental leave, but unlikely to be correlated with the parent's personal or employer characteristics. The policy reform in question is Quebec's move in 2006 to leave the national 'Employment Insurance' (EI) program and set up its own agency, the Quebec Parental Insurance Program (QPIP). QPIP's features were designed to offer an improvement over the older EI system by easing some of the barriers that parents face to taking leave, namely, inflexibility, ineligibility, financial feasibility, and gendered attitudes. First, the new system was more flexible, offering a choice between the Basic Plan or a Special plan that offers higher benefits for a shorter duration, thereby letting parents select the combination of benefit amount and duration that best suited their needs. Second, QPIP lowered the eligibility criteria. The EI system requires a claimant to have worked 600 hours of insurable employment, making it difficult for workers from seasonal, temporary, part-time or otherwise non-standard employment, who tend disproportionately to be low-income mothers, to qualify for the program. In comparison, under QPIP any parent who has at least 2000CAD of insurable earnings can qualify for benefits. Third, QPIP offers more generous financial compensation to leave-takers, by both increasing the maximum replacement rate (from 55% to 70%) and raising the ceiling of maximum insurable earnings on which one can claim (from 39,000CAD to 57,000CAD in 2006). QPIP also established a 'daddy quota', whereby 5 weeks of leave were set aside for the father and could not be transferred to the mother. In contrast, under the EI program fathers enjoy no individual right to leave and may only access benefits through shared parental leave. Thus, QPIP increased the amount of paid leave available to a family from 50 weeks to 55 weeks, such that total leave increased by the amount equivalent to the 'daddy-only' weeks.[6]

The validity of using exposure to QPIP as an instrument for fathers' leave-taking hinges on the success of the reform in boosting fathers' participation rates. Recall that in Chapter 1, I applied a sharp regression discontinuity approach to benefit claims data and found that the introduction of QPIP was associated with an immediate jump in fathers' participation rates of 53 percentage points (250% increase from pre-reform mean) and in the average duration of fathers' leave by 3.1 weeks (150% increase from pre-reform mean). Referring back to Figure 1.4 from Chapter 1, we see that fathers' leave participation rates in Quebec show a sharp jump in levels between 2005 and 2006 while participation in other provinces remained stable over the decade, confirming their validity as a control group.

It is important to note that although QPIP did not increase the weeks of leave available to women, it did increase financial compensation for all parents and therefore offered increased incentives for mothers to take leave as well.[7] Chapter 1 of this dissertation reports that QPIP had an average treatment effect of increasing mothers' participation rates by 12 percentage points - but had no significant effect on mothers' leave duration. Given the nature of the reform, it is not possible to delineate the long-term effects of increased fathers' leave-taking from those related to increased mothers' leave-taking. However, if we are concerned by the possibility that the changes in household division of labor might be driven by mothers' rather than fathers' leave behavior, we must remember two important things. First, QPIP's program effects on fathers' leave behavior dwarfed the program effects on mothers' leave behavior. For example, the increase in mothers' participation rates represented a 15% increase from the pre-reform mean for women. In comparison, the in-

[6] It should be noted that under either the EI or QPIP program, parents can take leave simultaneously so the mother does not have to resume work in order for the father to participate in parental leave.

[7] While in theory mothers may also have changed their leave behavior in response to their husband's increased leave-taking rather than directly in response to the reform, I explored these interaction effects and did not find a statistically significant association.

crease in fathers' participation represented a 250% increase relative to the pre-reform mean for men. So the program effects of QPIP on men were much larger, both in absolute and relative terms, than those on women. Second, since longer maternity leaves make mothers more available for time-inflexible non-market work and weaken their labor market position, they are associated with higher levels of sex specialization (Hook, 2006). That is, while fathers' increased leave participation is posited to reduce sex specialization, mothers' leave participation is expected to increase it. Therefore, if the small increase in mothers' leave participation under QPIP did have long-run effects on parental behavior, we would expect it to bias me against finding reduced sex specialization in households exposed to QPIP. In that case, we should consider my estimates of the long-term effects of QPIP on sex specialization to be underestimates of the true causal effect that paternity leave alone may have.

2.3.2 Data

To analyze the long-run effects of QPIP on the division of household labor, I use time-diary data from Canada's General Social Survey (GSS) (Statistics Canada, 2005, 2010). In this time-diary survey, the respondent is asked to record his or her activities as well as corresponding details such as locations and whether other people were present every 7 minutes over the 24-hour survey window. These data offer extremely precise measurements of time allocations, and are robust to response bias since the survey does not hint at gender issues. According to Kotsadaam and Finseraas (2012, pg 1619), who study the effect of a Norwegian daddy quota on the sharing of housework, "under ideal settings, we would exploit a time-use data-set with a large enough sample of individuals who had their last child in a time period around the reform to investigate the actual sharing, but no such data-set exists" (for Norway). Fortunately, exactly such a data-set exists for Canada, and I use it in this study.

I specifically use the two most recent rounds of the GSS that collected time-diary data: cycle 19 that was conducted in 2005 and cycle 24 that was conducted in 2010. Since QPIP was introduced in Quebec in 2006, observations from 2005 are considered to be in the pre-reform period while observations from 2010 are considered to be in the post-reform period. The target population of the GSS includes all persons 15 years of age and older in Canada excluding full-time residents of institutions. Approximately one-fifth of the observations are from Quebec, the treatment group, while the rest of the observations come from the control group that consists of other provinces in Canada.[8] The sample comprises parents aged 18-50 whose youngest child is aged between 1-8 and who are in a married or cohabitating relationship. Parents whose youngest child is under one year old are excluded from the sample to eliminate the possibility they may still be on parental leave at the time of the survey. In all analyses I exclude parents whose youngest child is aged 4 because, since the GSS does not record the child's exact date of birth, it is impossible to determine whether parents of 4 year olds interviewed in 2010 experienced this birth in 2005 or 2006, and therefore whether or not they were exposed to the QPIP.

I measure time spent in various types of work in minutes per day as recorded by their time-diary. For parents' market outcomes, I investigate the time spent in paid work (which includes commuting to paid work) and time spent physically at the workplace, as well as labor market outcomes such as employment status, full-time employment, usual weekly hours, and weeks worked last year. It should be noted that the measure of full-time employment is constructed by the author, using an indicator variable taking the value 1 when the respondent reports 35 or more hours usual weekly hours worked.

For parents' non-market outcomes, I examine total time spent in non-market work (the

[8]Specifically, the control group comprises observations of parents resident in Newfoundland and Labrador, Prince Edward Island, Nova Scotia, New Brunswick, Ontario, Manitoba, Saskatchewan, Alberta, British Columbia, Yukon, Nunavut and the Northwest Territories.

sum of housework and childcare) as well as childcare and housework separately, and total time spent physically at home and time spent in the vicinity of family members. Domestic work is the sum of time in 'core' non-market work (such as meal preparation and cleanup, laundry, ironing, dusting, and indoor cleaning), time spent obtaining goods and services (such as shopping for groceries or household supplies) and 'other' home production such as maintenance and repairs, gardening, and caring for houseplants and pets. Childcare is the sum of time spent in routine childcare (such as feeding children or getting them ready for school), interactive childcare (such as helping with homework or reading to children) and also travel and communication related to childcare such as driving children to school or attending a parent-teacher conference. I also look at "total time spent at home," i.e. the sum of all time that the respondent identifies his location as his residence, and "total time spent with family members," which is the sum of all time where the respondent notes a family member was also present. It is important to note that these last two outcomes could include time spent sleeping, which could be recorded "at home" or "in the presence of a family member" without necessarily indicating any contribution to home production or 'quality' time with family.

There are two limitations of the GSS data that merit mention. First, since the data are not collected at the couple-level, I cannot track changes in spouses' behavior within the same household; instead, my results show how mothers' and fathers' behavior changed on average across households. A second limitation is that given that families can move between provinces, it is possible that families observed in Quebec experienced their last birth in one of the other provinces or vice versa. However, given that the proportion of people moving in and out of Quebec in any given year is small.[9] Therefore, the number of recent-movers in and out of Quebec in my GSS sample of parents with young children is expected to be small and

[9]For example, using inter-provincial migration rates for 2007 reported by Milan (2011), I calculate that the proportion of people migrating into and out of Quebec was 0.26% and 0.41% respectively of Quebec's population.

unlikely to bias the results significantly. Moreover, since this type of cross-contamination should reduce differences between observations in Quebec and other provinces, my results would underestimate the true causal effect of paternity leave on long-term behavior.

2.3.3 Identification Strategy

To analyze difference-in-differences in outcomes over provinces and time, I estimate:

I estimate:

$$Y_{ijt} = \alpha + \beta I[j = Quebec] * I[t >= 2006] + \theta I[t >= 2006] + \phi Z_{ijt} + \lambda_j + \delta_t + \epsilon_{ijt} \qquad (2.1)$$

where subscript i denotes the individual, subscript j denotes province and subscript t denotes the year of last birth. Y_{ijt} therefore represents the outcome of mother i observed in province j who gave birth in year t. $I[t >= 2006]$ is an indicator variable taking the value 1 if the birth-year t is 2006 or greater, i.e., if the observation is from the post-reform period. The coefficient θ represents the change in the value of the outcome that is shared by all provinces. The term $I[j = Quebec] * I[t >= 2006]$ takes the value of 1 if the individual lives in Quebec and gave birth in a post-reform year, and otherwise takes the value 0. The coefficient β therefore represents the DD estimate of primary interest as it captures the change in the value of the outcome post-reform that is unique to Quebec. Under the assumption that no other policy changes were enacted to affect it, β represents QPIP's average treatment effect. λ_j and δ_t denote the fixed province and year effects. It should be noted that I do not control for all province-year interactions, but instead collapse them into the term $I[j = Quebec] * I[t >= 2006]$.

The term Z_{ijt} is a vector of personal characteristics including age, education, legal marital status and immigrant status as well as household characteristics such as family size, number of children aged 0-1 and 1-5 and 6-17. Including these as regressors controls for changes in group composition. ϵ_{ijt} is the error term. I calculate cluster-robust standard errors that generalize the White (1980) heteroskedastic-consistent estimates of OLS standard errors to the clustered setting in order to account for possible heteroskedasticity and within-province dependence of standard errors, which are particularly a concern in difference-in-difference estimations since the regressor of interest is highly correlated within clusters (Bertrand et al., 2004).

To identify the long-term causal effects of QPIP, I apply this difference-in-differences method to a GSS sample of parents whose youngest child is aged 1-3. However, since Time Use GSS data is only available every 5 years, the above method compares changes between 2005 and 2010 among Quebecois parents of children aged 1-3, compared to identical parents in other provinces. This gives rise to the concern that something else may have changed over that period in Quebec such that a simple double-difference could confound a Quebec-wide trend with a change in behavior causally related to the QPIP program. To check that the results from my DD regressions are not picking up Quebec-specific trends over time, I devise a robustness check that utilizes a placebo group of parents whose youngest child is aged 5-8.[10] These parents form a convenient placebo group because even if they are observed in the treated province in the post-treatment year (2010), their children are slightly too old for them to have been eligible for QPIP. These robustness checks thus use a difference-in-difference-in-differences (DDD) identification strategy that exploits variation in exposure to paternity leave across provinces, time, and age-group of the child. In this setup, a parent is only considered to be exposed to QPIP if they are observed in Quebec in 2010 and their youngest child was born since 2006, i.e., the child is aged 1-3.

[10] I vary the age restrictions on the placebo group, using the same minimum age of 5 but increasing the maximum from 8 up till 14 and get similar results.

For the robustness check, I run triple-differencing regressions that estimate:

$$\begin{aligned} Y_{ijta} = \alpha + \delta I[j = Quebec] * I[t >= 2006] * I[a <= 3] \\ +\beta I[j = Quebec] * I[t >= 2006] + \sigma I[a <= 3] * I[t >= 2006] \\ +\theta I[j = Quebec] * I[a <= 3] \\ +\gamma I[t >= 2006] + \chi I[a <= 3] + \phi Z_{ijta} + \lambda_j + \epsilon_{ijta}, \end{aligned} \tag{2.2}$$

where subscript i denotes the individual, subscript j denotes province and subscript t denotes the year. Y_{ijta} represents the outcome for parent i in province j in year t in the child's age group a. $I[t >= 2006]$ is an indicator variable taking the value 1 if the observation is from after 2006, i.e., if the observation occurred after the reform was introduced in Quebec. The coefficient γ represents the change in the value of the outcome that is shared by parents in all provinces. An interaction term, $I[j = Quebec] * I[t >= 2006]$, is included to capture the change in the value of the outcome post-reform that is unique to Quebec. $I[a <= 3]$ is an indicator taking value 1 when the age of the parent's youngest child is less than 3 years old, and taking value 0 if the child is older. The parameter of interest is δ, the coefficient on $I[j = Quebec] * I[t >= 2006] * I[a <= 3]$, which captures the effect of being in the treated province in the post-treatment period and having a child young enough that the parent was eligible for the treatment. The term Z_{ijta} is a vector of personal characteristics including age, spouse's age, marital status, nation of birth, as well as household characteristics such as family size, number of children, and age of youngest child.[11] I control for Province fixed effects through the term λ_j. ϵ_{ijta} is an i.i.d error term. Here again I calculate heteroskedasticity-robust standard errors which are clustered at the province level.

The inclusion of the placebo group of parents of slightly older children offers a good robustness check as it differences out changes that have occurred in Quebec over time that are

[11] I do not control for educational characteristics for two reasons. First, for the education questions the GSS Time Use survey has lower response rates, imposing significant restrictions on sample size. Second, leaving education controls out of the regressions allows the education of each parent to be endogenous.

unrelated to exposure to QPIP. For example, regional economies may have fared differently in the recent recession, but the parents of children of different ages within the same province would have faced the same economic opportunities. Thus including placebo fathers of older children accounts for the fact that Quebec's economy may have suffered more or less than other provinces in the years between 2005 and 2010. Moreover, if one were concerned that the introduction of QPIP is endogenous to a Quebecois culture that increasingly values gender equality, then including Quebecois fathers of slightly older children will control for trends in egalitarian beliefs in Quebec. Therefore, when evaluating the validity of my main DD results one can use the following protocol. The validity of DD results as a causal link is supported whenever the triple-difference results have the same sign and a similar or larger magnitude. But in cases where the triple-difference results have an opposite sign or a smaller magnitude, we must be careful in interpreting the DD results as they may be picking up some Quebec-specific trends. In discussing my DD results later in the paper, I focus only on the changes in parents' behavior which are supported by both the double-difference and triple-difference results.[12]

Table 2.1 presents sample characteristics for the GSS data across treatment, control and placebo groups in the years 2005 and 2010, as well as the differences across the groups over time. Reassuringly, I detect only one significant difference across the groups over time: between 2005 and 2010 the proportion of fathers who were not born in Canada grew more rapidly in the 'exposed' group (i.e. fathers in Quebec with a youngest child aged 1-3) than in the other groups. However in every other characteristic such as age, spouse's age, number and age of children, family size, there are no reasons to believe the triple-differencing identification strategy would be mistakenly picking up changes in group composition. Nevertheless, I include controls for these parental and household characteristics in each triple-differencing regression.

[12] In addition, I run difference-in-difference regressions separately on the sample of placebo parents in Appendix Tables B.1 and B.2.

2.4 Results

Before beginning my discussion of the long term effects of QPIP, it is worth establishing a baseline for how households behaved prior to the existence of QPIP. Table 2.2 shows mean outcomes for parents with youngest children aged 1-3 in the year 2005; it shows that household responsibilities were clearly divided along gendered lines. Mothers spent more time in non-market work including housework and childcare, especially in time-inflexible chores such as cooking, housekeeping and routine childcare. The only household chore in which fathers spend more time than mothers is that of maintenance and repairs, which is flexible, not routine and in line with norms of masculinity. The ratio is reversed when we consider market work and labor market outcomes. Fathers spend considerably more time in paid work and physically at the workplace. Moreover, fathers are more likely to be employed, full-time workers, work longer hours per week, and work more weeks per year. A clear pattern of sex specialization within the household is evident, providing us with a baseline against which to evaluate the magnitude of any program effects.

Table 2.3 presents results exploring parents' involvement in market production. Panel I shows results from both double-difference regressions using the sample of parents whose child is aged 1-3, and to check that these DD results are not just reflecting Quebec-specific trends, Panel II shows results from triple-difference regressions including 'placebo parents' whose child is aged 5-8. For the sake of brevity, I will focus my discussion on only those changes detected by the double-difference regressions that are not eliminated when we include the placebo group in the triple-difference regressions. The DD cannot detect statistically significant changes in fathers' time in market work, though Column 1 does a decrease in time in paid work that is large enough to be economically significant, and is supported by the triple-difference results. Mothers, on the other hand, experience impressive gains in market outcomes if exposed to QPIP. Exposed mothers are found to spend more time in paid work per day. The DD analysis finds that, conditional on being employed, exposed mothers spend 79 minutes longer physically at the workplace (44% increase from baseline), and are 5.4 percentage points more likely to be full-time employed (7% increase from baseline), compared to mothers who experienced their last birth under the EI program. Exposed mothers report working more hours per week, but working fewer weeks per year. Overall, Table 2.3 presents evidence of increased female investments in market work.

In Table 2.4, I present results for parents' time in non-market work and related outcomes. The DD finds that exposed fathers spend 37 minutes longer in non-market work per day, representing a 23% increase from baseline. In housework, the DD finds exposed fathers spend 15 minutes longer per day than their counterparts, a 21% increase from baseline. In addition, exposure to QPIP is associated with an increase of 36 minutes spent physically at home. Though the DD finds an increase of 21 minutes in childcare by exposed fathers, no such change is detected by the triple-difference regression, suggesting that all fathers in Quebec are spending more time with their children in the post-reform period, regardless of exposure to the QPIP. Interestingly, exposure to QPIP is also associated with increased time spent by mothers in non-market work, although the absolute and relative magnitude of their increase is smaller than that of fathers. The DD results show that exposed mothers reduce their time in housework by 18 minutes and increased their time in childcare by 48 minutes, leading to an increase of 30 minutes in total non-market work (10% increase from baseline). Exposure to QPIP is also associated with mothers spending 30 fewer minutes physically at home per day.

Since exposed fathers increase non-market work more than do exposed mothers, Table 2.4 does suggest that in aggregate female specialization in home production is reduced. Nevertheless, it is interesting that I detect any increase at all in mothers' childcare rather than a decrease. There are several possible explanations for this. First, as previous research has suggested, mothers may be less willing to reduce time in childcare than in other household duties (Craig, 2005), so when paternity leave induces fathers to increase non-market contributions it may be efficient for them to increase time in housework that is less preferred by mothers. Alternatively, both parents may have equal preference for childcare but exposed mothers have gained bargaining power that they use to negotiate away from less-preferred housework and towards more-preferred childcare. Lastly, a possible explanation is that although parents' relative productivities in market versus non-market work may have changed, within the realm of non-market work mothers may still have a comparative advantage in childcare versus housework.

As I exploit variation in exposure to QPIP rather than actual participation, I provide evidence on 'intent-to-treat' (ITT) estimates that are preferable to estimates of 'treatment on the treated' (TOT) for several reasons. First, TOT estimates could be subject to the same bias from selection into treatment that previous cross-sectional studies have been criticized for. Second, from a policy-making perspective, ITT effects may be more relevant as they allow for feedback effects whereby the 'daddy quota' could have changed expectations and norms over and above the effects of actually using the leave option. The 'daddy quota' sent a strong public message about the importance of fathers' involvement in the home, which may have incentivized fathers who were exposed to QPIP but not treated to nevertheless change their behavior. Furthermore, a change in the behavior of treated households may change costs and incentives for neighboring households. For example, workplace expectations for all mothers may rise as treated mothers increase their career commitment and consequently, the penalty may increase for the untreated mothers who do not. Nevertheless, it is safe to assume that feedback effects on parents who were exposed but not treated are smaller than the first-order effects on parents who were treated, such that the ITT results presented here provide a lower bound for the causal effects of paternity leave on those who actually take it.

2.4.1 Threats to Identification

An important issue that must be addressed is that of changes in fertility. Since QPIP provided greater financial incentives to have children, it may have led to an increase in fertility rates in Quebec. This gives rise to the concern that the long-term program effects I detect may be driven by the marginal couples that are induced to have a child by the new program. I address this concern in multiple ways. First, it should be noted that the marginal couples do not appear different in any observable way, since I do not find significant changes in sample characteristics in Quebec compared to other provinces (see Table 2.1). For example, one could hypothesize that due to the generosity of QPIP's benefits people were concerned that the program was temporary and so couples may have rushed to have a baby sooner than they would otherwise. If this were the case, one would expect to see a unique decrease in the age of parents in Quebec, but Table 2.1 shows this was not so. Moreover, I control for personal and household characteristics in all my regressions, so the program effects I detect should not be biased by Quebecois households having larger families or more young children in the post-reform period. Second, if we consider that QPIP's effects on leave behavior were heterogeneous based on some unobservable factor, it becomes unclear what direction this bias, if it exists, would take. On the one hand, one could imagine that the fathers most responsive to paternity leave incentives are those who are more open-minded, which would bias my study in favor of finding reduced sex specialization. However, since we saw that most families did not exhaust their leave prior to QPIP, these open-minded fathers were always free to take leave. On the other hand, one could argue that QPIP's reformed features targeted more traditional families, which would bias my study against finding reduced sex specialization. For example, the increase in income replacement provided the greatest marginal benefit to low-income parents, who are more likely to have traditional beliefs. Also, QPIP's daddy-quota provided stronger incentives for fathers who were previously financially dis-incentivized from participating -

e.g., those who earned significantly more than their wives - who are also more likely to have traditional beliefs. Therefore, it cannot be known which direction this fertility bias, if it exists, would take, and an argument can be made that it may lead me to under-estimate the true program effect of QPIP on sex specialization.

2.5 Conclusion

This study also offers the first comprehensive causal analysis of the effect of a policy promoting paternity leave on the household division of labor. It provides strong evidence that by altering the initial distribution of parenting responsibilities, paternity leave can influence household decisions about how to allocate parents' resources to childcare, domestic work and paid work in later years. The results of this study have important policy implications. First, they suggest that it is possible for policies that induce changes in short-term behavior to have persistent effects on people's behavior in the long term, i.e., that a reform resulting in an increase in fathers' leave duration of 3 weeks could be sufficient to stimulate a shift in household dynamics for years to come. Second, my results suggest that there need not be a trade-off between gender equality and parental investments in children, such that paternity leave may present us with a rare win-win scenario.

Table 2.1: Mean Characteristics of GSS Data

	Parents with youngest child aged 1-3				Parents with youngest child aged 5-8				Difference-in-
	Control 2005	Control 2010	Quebec 2005	Quebec 2010	Control 2005	Control 2010	Quebec 2005	Quebec 2010	Difference-in Differences
I. Fathers' Characteristics (N=1606)									
Age	34.896	34.815	34.407	34.120	39.831	39.796	39.661	39.523	-0.102
Age of spouse	36.256	32.959	31.933	30.877	37.713	38.048	37.161	37.617	-1.103
Canadian Born	0.767	0.736	0.855	0.701	0.770	0.694	0.829	0.879	-0.248**
Born in Quebec	0.027	0.023	0.824	0.677	0.018	0.028	0.812	0.828	-0.150
Age of Youngest child	1.905	1.675	1.624	1.566	6.517	6.507	6.692	6.548	0.306
Children under 14	1.963	1.895	1.842	1.875	1.947	1.887	1.905	1.892	0.054
Children in Household	1.939	1.920	1.857	1.895	2.096	1.998	2.127	1.929	0.157
Legally Married	0.889	0.851	0.499	0.522	0.896	0.926	0.531	0.457	0.165
Household size	4.023	4.034	3.875	4.018	4.196	4.080	4.146	3.963	0.198
II. Mothers' Characteristics (N=1936)									
Age	32.465	32.600	31.748	31.111	36.945	38.748	37.042	38.485	-0.412
Age of Spouse	34.331	34.782	33.605	33.248	37.576	38.962	40.882	38.904	0.285
Canadian Born	0.728	0.674	0.786	0.814	0.729	0.692	0.803	0.754	0.094
Quebec Born	0.022	0.018	0.753	0.799	0.025	0.023	0.747	0.733	0.064
Age of Youngest child	1.841	1.606	1.853	1.534	6.400	6.483	6.610	6.232	0.369
Children under 14	1.828	1.929	1.736	1.777	1.952	2.012	2.009	1.941	0.068
Children in Household	1.827	1.924	1.730	1.800	2.160	2.075	2.086	2.038	-0.063
Legally Married	0.897	0.864	0.458	0.447	0.917	0.917	0.568	0.534	0.053
Household size	3.975	4.032	3.813	3.889	4.266	4.236	4.095	4.068	0.016

Notes: Data comprises a GSS sample of parents aged 18-50 who report being in a cohabitating or married relationship and whose youngest child is aged between one and eight years old. The last column presents difference-in-difference-in-difference in means across years, provinces and children's age groups.

Table 2.2: Baseline Sex Specialization in Quebec before the Reform

	Fathers	Mothers	$\frac{Fathers}{Mothers}$
Daily minutes in non-market work	158.70	312.83	0.51
Daily minutes in childcare	88.69	165.42	0.54
- Interactive childcare	42.56	61.96	0.69
- Routine childcare	37.77	87.72	0.43
- Travel & communication	8.36	15.73	0.53
Daily minutes in domestic Work	70.01	147.41	0.47
- Cooking	31.78	70.88	0.45
- Housekeeping	16.04	64.38	0.25
- Shopping	48.91	50.86	0.96
- Other chores	12.36	10.45	1.18
- Maintenance & repairs	9.80	1.69	5.79
Daily minutes at home	873.68	1,134.92	0.76
Daily minutes with family	409.07	509.32	0.80
Daily minutes in paid work	416.75	168.54	2.47
Employed	0.836	0.572	1.46
Daily minutes at workplace (if employed)	355.75	180.89	1.96
Full-time (if employed)	0.961	0.759	1.27
Usual weekly hours (if employed)	44.025	31.109	1.42
Weeks worked if (employed)	47.452	42.865	1.11

Notes: Table presents means from a sample of data from the 2005 GSS, comprising parents aged 18-50 who report being in a cohabitating or married relationship, living in Quebec, and having a youngest child aged 1-3 years old. Time spent at home and with family may include time spent sleeping.

Table 2.3: Exposure to QPIP and Parents' Market Outcomes

OUTCOMES:	(1) Time in Paid Work	(2) Employed	(3) Time at Workplace (if emp)	(4) Full-time Worker (if emp)	(5) Usual Weekly Hours (if emp)	(6) Weeks worked Last Year (if emp)
I. Double-Differences using parents with youngest child aged 1-3						
Fathers:						
Quebec * Post-reform	-43.296	0.002	2.293	0.035	0.016	1.945***
	[0.16]	[0.90]	[0.70]	[0.12]	[0.90]	[0.00]
N	988	988	901	888	885	888
Mothers:						
Quebec * Post-reform	60.147***	0.046*	79.908***	0.054*	1.353*	-3.257***
	0.00	[0.09]	[0.00]	[0.07]	0.07	0.00]
N	1115	1115	696	692	691	692
II. Triple-differences using all Parents with youngest child aged 1-8						
Fathers:						
Child Under 3 *	-178.046***	0.017	-122.80**	0.021	2.199	-2.03**
Quebec * Post	[0.00]	[0.44]	[0.02]	[0.40]	[0.19]	[0.04]
N	1596	1596	1468	1440	1436	1436
Mothers:						
Child Under 3 *	35.443	-0.003	87.724***	0.177**	4.295**	-3.031**
Quebec * Post	[0.23]	[0.92]	[0.00]	[0.01]	[0.03]	[0.04]
N	1939	1939	1286	1278	1273	1271

*** $p<0.01$ ** $p<0.05$ * $p<0.1$, Robust province-clustered pvalues in parentheses

Notes: Data comprises a GSS sample of married/cohabitating parents aged 18-50 whose youngest child is aged 1-8 years old.

Table 2.4: Exposure to QPIP and Parents' Non-Market Outcomes

OUTCOMES:	(1) Total Time in Non-market Work	(2) Time in Domestic Work	(3) Time in Childcare	(4) Time at Home	(5) Time with Family
I. Double-Differences using parents with youngest child aged 1-3					
Fathers:					
Quebec * Post-reform	37.031***	15.312**	21.722***	36.371**	-29.811
	[0.00]	[0.01]	[0.00]	[0.02]	[0.22]
N	988	988	988	988	988
Mothers:					
Quebec * Post-reform	30.828**	-18.025**	48.853***	-30.342**	-11.237
	[0.01]	[0.04]	[0.00]	[0.01]	[0.68]
N	1115	1115	1115	1115	1115
II. Triple-differences using all Parents with youngest child aged 1-8					
Fathers:					
Child Under 3 *	43.404*	41.851**	1.542	67.797	29.103**
Quebec * Post	[0.07]	[0.02]	[0.90]	[0.11]	[0.03]
N	1596	1596	1596	1596	1596
Mothers:					
Child Under 3 *	38.826***	-28.296***	67.123***	-33.936	-12.159
Quebec* Post	[0.00]	[0.00]	[0.00]	[0.16]	[0.17]
N	1939	1939	1939	1939	1939

*** $p<0.01$, ** $p<0.05$, * $p<0.1$, Heteroskedasticity-robust province-clustered p-values in parentheses

Notes: Data comprises a GSS sample of married/cohabitating parents aged 18-50 whose youngest child is aged 1-8 years old.

CHAPTER 3

SUBSIDIZING BREASTFEEDING: DOES PAID PARENTAL LEAVE REDUCE BREASTFEEDING INEQUALITIES?

3.1 Introduction

Many medical authorities, including the American Academy of Pediatrics and the American College of Obstetricians and Gynecologists, offer recommendations that mothers breastfeed their newborn babies, motivated by the scientific evidence of a positive association between breastfeeding and infant health. Several cross-sectional studies of child health find an association between breastfeeding and lower rates of diarrhea, respiratory tract infections, otitis media and ear infections, infectious diseases, as well as infant deaths due to these diseases (For a review, see Leon-Cava (2012) or Ip et al. (2007)). Breastfeeding selectively protects against extremes in body size and fat deposition, and is therefore negatively associated with not only childhood obesity but also obesity across the life course (Owen et al., 2005; Harder et al., 2005; Crume et al., 2012). There is even evidence to suggest that breastfeeding has a positive impact on the cognitive and behavioral outcomes of children (Borra et al., 2012). Breastfeeding has also been shown to benefit mothers' health. By delaying the return of women's fertility, it limits exposure to the health risks of multiple births within short intervals. It is also associated with reduced risk of type 2 diabetes, and breast and ovarian cancer in mothers (Ip et al., 2007). Due to these many benefits for the health of the mother and the child, the World Health Organization currently recommends that children be exclusively breastfed for the first 6 months, and breastfed in addition to other foods up until they are 2 years old ().[1]

[1]Exclusive breastfeeding refers to the practice of feeding the baby on only breastmilk, with no solid foods, liquids, or water.

In reality, however, these goals are rarely met. Statistics published for the United States in 2014 report that on average only 26.7% of mothers breastfed for a full year and only 18.8% of mothers exclusively breastfed for the first 6 months (CDC, 2014). Even more disconcerting are the vast inequalities in breastfeeding outcomes across the population. Studies consistently show that mothers who are less educated or of low socio-economic status are both less likely to start breastfeeding and less likely to continue breastfeeding as long as other women (Forste et al., 2001; MccAndrew et al., 2012). For example, the CDC (2007) reports that although the rates of exclusive breastfeeding for 3 months were 30% for the United States on average, they were significantly lower among mothers who were black (19.8%), unmarried (18.8%), resided in rural areas (23.9%), had a high school education or less (22.9%), or lived below the poverty line (23.9%). This is concerning because the practice of breastfeeding may be particularly important for families of low socio-economic status. Although substitutes for breastmilk are available, poor families are less able to purchase high-quality substitutes for breastmilk.[2] Moreover, children from disadvantaged families are exposed to greater health risks, e.g., communicable diseases from low quality daycare or childhood obesity from poor diet, and could therefore particularly benefit from the protective benefits of breastfeeding.

As employment offers a critical barrier to breastfeeding by constraining mothers' time and availability, maternity leave may facilitate breastfeeding by enabling new mothers to delay their return to work. Indeed, a significant increase in maternity leave provisions has been shown to causally increase breastfeeding durations (Baker and Milligan, 2008). Unfortunately, the way in which many parental leave programs are structured, with strict eligibility criteria and limited financial compensation, creates vast inequalities in leave-taking among mothers. In turn, these inequalities in parental leave participation may contribute to the educational and income gaps in health behaviors such as breastfeeding. This leads

[2]For example, studies show that infant formula with high protein content is associated with rapid early weight gain (which in turn is associated with later obesity), compared to infant formula with low protein content (Koletzko et al., 2009).

to an important policy question: is it possible to design a leave program that can reduce inequalities in maternal leave-taking, and if so, would that then lead to reduced inequalities in health outcomes?

In this paper I seek to answer this question while providing a causal analysis of a Canadian reform to paid parental leave. I investigate a natural experiment where on January 1st 2006 the province of Quebec left the Employment Insurance (EI) Program that the rest of Canada subscribes to, and established a system called the Quebec Parental Insurance Program (QPIP). QPIP aims to be more equitable than its predecessor in two important aspects. First, QPIP has easier eligibility criteria that allow many more women to qualify for government benefits while on job-protected leave. Second, QPIP's benefits program offers more generous financial compensation, which decreases the opportunity cost of taking leave, especially for low-earning mothers. I use this quasi-experimental setup to investigate how the QPIP reform affected inequalities in leave-taking and breastfeeding among various groups. This paper makes an important contribution to the small causal literature exploring the health consequences of parental leave policies. The few causal studies on this topic to date have examined policies which tended to favor more advantaged women and have focused on their impact on average health outcomes. In contrast, the current study examines a policy which specifically sought to make parental leave more equitable and explores whether the policy was successful in this regard and whether that in turn reduced inequalities in breastfeeding.

To analyze how QPIP affected mothers' leave-taking across various groups I use data on benefit claims from Statistics Canada's Employment Insurance Coverage Surveys (2002-2010). I utilize a difference-in-differences approach that exploits plausibly exogenous variation in exposure to the QPIP program across provinces and time. I find that on average QPIP increased mothers' leave participation rates by 19% and leave duration by 5.6 weeks.

Furthermore, I find particularly large increases in leave utilization by never-married mothers and mothers from lower-income households, suggesting that QPIP successfully reduced some inequalities in maternal leave-taking.

To study QPIP's effects on breastfeeding I use data from the Canadian Community Health Surveys 2(2005,2007-2013) and a similar difference-in-differences approach. I find that QPIP greatly improved average breastfeeding outcomes: it is associated with a 6% increase in breastfeeding initiation rates, a 23% increase in the likelihood of breastfeeding a full year, and a 25% increase in the likelihood of exclusively breastfeeding for at least 20 weeks. However, I find that QPIP did not succeed in reducing inequalities in breastfeeding. For example, though QPIP increased leave-utilization among never-married mothers more than married mothers, I only observe improvements in breastfeeding among married/cohabitating mothers, and I detect no program effect on the breastfeeding of single mothers. Similarly, even though QPIP reduced the income gradient in leave-taking for mothers, the pattern was not replicated for breastfeeding: mothers from the lowest-income households experience a smaller improvement in breastfeeding compared to mothers from the highest-income households.

In general, the positive association between parental leave and breastfeeding can be attributed to reduced opportunity costs of parents' time. However, my results show that when it comes to breastfeeding decisions, disadvantaged mothers are less sensitive to changes in opportunity costs of time compared to more advantaged mothers. This suggests that making paid parental leave accessible is not effective in reducing inequalities in breastfeeding - or at least that it may be a necessary but not sufficient condition. Policy-makers concerned with reducing disparities in breastfeeding across society should also consider policies that can increase the latent demand or preferences for breastfeeding within these disadvantaged groups.

The remainder of the Chapter is organized as follows. Section 3.2 provides background by discussing the factors affecting breastfeeding decisions, and the small existing literature on the role that parental leave can play. Section 3.3 provides details of the various leave programs in Canada and discusses the implications of the 2006 reform in Quebec. Section 3.4 describes the empirical methods and data that is used in my analysis. Section 3.5 presents results and a discussion of their implications, while Section 3.6 concludes.

3.2 Background

3.2.1 Factors Affecting the Breastfeeding Decision

Apart from a mothers' physical ability to breastfeed, which is outside the scope of public policy, two other parts make up the decision problem of whether and how long to breastfeed: the mother's latent demand for breastfeeding and the opportunity costs of doing so.

Multiple factors determine a mother's latent demand for breastfeeding. First, there is the mothers' ante-natal intention to breastfeed, which is one of the strongest predictors of breastfeeding duration (Blyth et al., 2004). This intention is influenced by how informed she is about the potential health benefits and cost savings offered by breastfeeding. Second, preference for breastfeeding increases with mothers' self-efficacy in breastfeeding (Blyth et al., 2002, 2004; Forster et al., 2006), which in turn has been shown to increase with a woman's ante-natal preparedness for breastfeeding, use of the correct suckling technique, and initial experience in initiating breastfeeding (McLeod et al., 2002; Cernadas et al., 2003). Third, attitudes within the mothers' family and society are also important. Numerous studies report that husbands' support, knowledge and attitudes to breastfeeding are a significant predictor of breastfeeding duration. (Littman et al., 1994; Kessler et al., 1995; Bar-Yam and Darby, 1997; Humphreys et al., 1998; Kong and Lee, 2004; Pisacane et al., 2005; Susin and Giugliani, 2008; Maycock et al., 2013; Bich et al., 2014; Mueffelmann et al., 2014; Abbass-Dick et al., 2015). Social pressure to breastfeed has also been shown to influence mothers (Swanson and Power, 2005). All of these psychological, social and support factors can influence a mothers' inherent demand for breastfeeding. It should be noted that these factors may vary across socio-economic groups, and are likely to favor more advantaged mothers. High-income, well-educated, and married mothers are likely to have greater latent demand for breastfeeding because they are more likely to be informed about its benefits, have access to good pre- and post-natal professional support, and be surrounded by a social network that offers encouragement as well as pressure to breastfeed longer.

In addition to latent demand for breastfeeding, a mother's decision to breastfeed is also determined by an important constraint: employment. Breastfeeding requires mother and infant to be physically together, which is often not possible when a mother works outside the home. Even when a woman can express milk manually or using a breast pump to reduce the spatial conflict, she will still face a time conflict with her job. It is therefore unsurprising that employment outside the home is associated with shorter breastfeeding durations (Fein and Roe, 1998; Kimbro, 2006; Guendelman et al., 2009; Ogbuanu et al., 2011; Skafida, 2012). On average, employed mothers breastfeed for shorter durations than non-employed mothers (Lindberg, 1996) and the timing of mothers' return to work is closely associated with the cessation of breastfeeding (Bick et al., 1998; Roe et al., 1999; Berger et al., 2005). In economic terms, we can think of foregone earnings from employment as the opportunity cost of breastfeeding. High-income mothers face greater opportunity costs of breastfeeding, as they give up more earnings during the time used to breastfeed.

3.2.2 Prior Studies on Parental Leave and Infant Health

Using the framework above, we expect parental leave policies to improve health outcomes since they reduce the opportunity cost of time spent at home and therefore some of the costs associated with investments in child and maternal health. Though only a handful of studies have examined the specific outcome of breastfeeding, there exists a broader literature on leave policies and infant and maternal health. Much of the extant literature is cross-sectional or cross-country in nature. Ruhm (1998) and Tanaka (2005) find a negative association between nations' parental leave policies and neo-natal mortality and child mortality between ages 1-5 in European nations. Roe et al. (1999) reports a positive association between the length of leave a mother takes from work and the duration of breastfeeding. Maternal return to work within the first 12 weeks from childbirth is negatively associated with breastfeeding and immunizations (Berger et al., 2005). Several studies found poorer mental health outcomes in terms of depression and anxiety among mothers who took longer maternity leaves compared to mothers who took shorter leaves (Gjerdingen and Chaloner, 1994; Chatterji and Markowitz, 2012). However, these cross-sectional studies may be biased due to the characteristics of mothers or governments that select into longer or shorter maternity leaves, so we must interpret these findings as positive associations rather than causal relationships.

More recently, a small handful of studies have used plausibly exogenous variation in leave policies to try to identify causal links between parental leave provisions and health outcomes. Their findings are consistent with the idea that a reduction in the opportunity cost of taking leave (i.e. foregone earnings) causally increases child health outcomes. Rossin (2011) examines the introduction of the Family and Medical Leave Act in the United States in 1993 to establish a causal relationship with birth and infant health outcomes. Using variation across firm size, prior state-level leave legislation, and mothers' likely eligibility status, she finds that unpaid job-protected leave due to the FMLA led to small increases in birth weight, decreases in the probability of premature births, and decreases in infant mortality among those mothers able to take advantage of the FMLA's provisions. Baker and Milligan (2008) investigate a reform to parental leave in Canada in 2000 to establish a causal link with breastfeeding behavior. The authors study an extension to parental leave under the EI program, such that mothers having children born on or after 30 December 2000 were eligible for 1 year of paid leave instead of 6 months.[3] The authors report that mothers' time away from work increased by 3-3.5 months, and relatedly, the average duration of breastfeeding increased by 1 month and the average duration of exclusive breastfeeding increased by 0.5 months.

Baker and Milligan (2008) do not investigate whether whether the reform had heterogeneous impacts on different sub-groups - in fact, they eliminate some sub-groups such as single parents and Quebecois parents entirely from their sample. In addition, it is important to note that the 2000 reform that they study was an extension to paid leave, i.e., it offered more weeks of leave to a population that, for the most part, already qualified for some leave under the old scheme, but it still left out most mothers from disadvantaged backgrounds who could not meet the eligibility criteria. Therefore, it is reasonable to interpret that the 2000 EI extension did little to mitigate inequalities in leave-taking, and that the positive effects on breastfeeding documented by Baker and Milligan (2008) were concentrated amongst relatively-advantaged mothers who were able to avail themselves of the extended EI benefits. A similar story of increasing inequality is apparent in studies of the United States, where the introduction of unpaid job-protected leave through the FMLA only increased leave-taking among college-educated and married mothers (Han et al., 2009). Relatedly, Rossin (2011)

[3]The details of financial compensation such as income replacement rates and caps for weekly benefits were not affected, such that the reform increased the length of paid leave to those who qualified but did not affect the compensation per week of leave. The work-hours eligibility criteria was also lowered slightly from 700 hours to 600 hours, increasing the proportion of women with EI-insurable employment from 70% in 2000 to 74-75% in 2001-2005. In comparison, the 2006 QPIP reform increased eligibility rates from 78% to 92%.

found that the FMLA only improved child health outcomes among college-educated and married mothers, but not among single mothers and those without college degrees. These studies show that these parental leave policies causally improved health outcomes by reducing opportunity costs. At the same time, these studies reveal that since these leave policies provided the greatest relief to more advantaged families, they may have actually increased health inequalities across society. What is not known, however, is what kinds of public policies would reduce these health disparities. Would a program that brings paid parental leave within the reach of all families then reduce inequalities in health behaviors? Or are families of low socio-economic status less sensitive to the opportunity costs of health behaviors - such that offering them paid leave would do little to improve their health outcomes? It is exactly these questions that the current study seeks to answer.

3.3 The Natural Experiment

3.3.1 The QPIP Reform

In Canada, the Employment Standards Act of 2000 established minimum standards for job-protected leave for parents, mandating that any person who has been working with their employer for at least 13 weeks prior to the birth of a child is entitled to 35 weeks of unpaid job-protected leave from work. In addition, birth-mothers qualify for another 17 weeks of unpaid job-protected leave. This means that, regardless of compensation, most working mothers have access to a year of unpaid job-protected maternity leave. Further, government benefits allow some parents to convert some of this leave into paid parental leave are available through certain programs. The Employment Insurance (EI) Program, which all Canadian provinces used until 2005, offers maternity benefits for mothers as well as parental benefits that mothers and fathers may share between them. On the 1st of January 2006, Quebec left the EI system and introduced the Quebec Parental Insurance Plan (QPIP). To review the details of the EI program, still offered to residents of other Canadian provinces, and the QPIP program, currently offered to residents of Quebec, please refer back to Table 1.1 of Chapter 1.

Recall that the 2006 reform comprised several changes that sought to tackle some of the barriers that parents faced to taking leave, namely ineligibility, financial feasibility, and in the case of fathers, social stigma. First, QPIP lowered the criteria for qualifying for parental leave benefits. The EI program requires a mother to have worked 600 hours of insurable employment with her employer in the last year, making it difficult for parents who are self-employed or worked seasonal, part-time, contractual or informal jobs - primarily low-income mothers - to qualify for benefits. In comparison, QPIP only requires a parent to have earned 2000CAD of insurable earnings in the last year.Insurable earnings and employment refers to work an EI or QPIP paycheck contribution has been made from. This change in eligibility criteria made a significant difference, introducing paid parental leave for the first time to a sizable chunk of the Quebecois population. In my data, I find that in Quebec the proportion of eligible mothers rose from 78% under the EI program to 92% under the QPIP Program.

Second, QPIP eased the financial burden of leave-taking in several ways. It removed the 2 week 'waiting period' under EI that necessitated mothers take 2 weeks of unpaid leave before they could begin receiving benefits. While under the EI program mothers can recover only 55% of their previous earnings through benefits, under QPIP mothers can recover 70% of earnings for the first 25 weeks of leave and then 55% of earnings thereafter. QPIP also raised the income ceiling on which benefits can be claimed. In 2006, EI only allowed parents to claim benefits on 39,000CAD of annual income, while QPIP allowed parents to claim benefits on 57,000CAD of annual income. QPIP also offers a Special Plan that allows families to elect into receiving higher benefits for a shorter duration of time if they would prefer.

Third, QPIP changed the way that leave was distributed within the household. The EI program comprises 15 weeks of maternity leave and 35 weeks of parental leave that mothers and fathers can share. QPIP introduced the nation's first of it's kind 'daddy-only' quota by offering 5 weeks of leave on a 'use-it-or-lose-it' basis to fathers. Thus QPIP changed the allocation of leave to be more gender-specific by reducing the amount of gender-neutral parental leave to 32 weeks, and increasing the mother-only and father-only leave to 18 weeks and 5 weeks respectively. Notably, the total amount of potential parental leave that a mother had access to remained the same, but a father's potential leave increased by 5 weeks.

3.3.2 Expected Effects of QPIP on Breastfeeding

We expect QPIP's lowered eligibility criteria and improved compensation to have had an unambiguously positive effect on mothers' leave-taking. For married mothers, the effect of the fathers' quota on mothers' leave-taking is ex-ante ambiguous. On the one hand, increased paternal leave-taking could crowd out maternal leave-taking, if fathers now used more of the shared parental leave. However, in Chapter 1 of this dissertation I found that although fathers' leave participation increased dramatically under QPIP, the average father consumed exactly the 5 'daddy weeks' allocated to him. Therefore, the opposite is more likely, that fathers consuming their individual daddy quota freed up the shared parental leave for mothers to use. Moreover, fathers' leave-taking may be a complement rather than a substitute to mothers' leave-taking. On net, Patnaik (2015) found that on average QPIP increased mothers' leave participation of 16-25%. We would expect this increase in maternal leave-taking to have a positive impact on average breastfeeding initiation rates and duration.[4]

We also expect QPIP to have reduced some inequalities in mothers' leave utilization. First, the lowering of the eligibility criteria meant that many mothers from disadvantaged backgrounds that could not meet the EI program's eligibility criteria for benefits, could now qualify for compensation under QPIP. Although these mothers may have always had access to unpaid job-protected leave, these mothers were much less likely to have access to employer top-up payments, private parental insurance, or personal savings to support them during unpaid leave. Therefore giving them access to government benefits during parental leave should have particularly facilitated their leave-taking. Second, although the income replacement rates and earnings ceilings increased for everyone, the capping of benefits means that the lowest-earning parents would have experienced the larges marginal increase in benefits. Therefore, even among mothers who always qualified for paid leave, QPIP granted an especially generous increase in compensation for lower-earning parents. Overall, mothers from more vulnerable populations experienced the greatest decrease in opportunity costs under QPIP, so we expect QPIP to to have boosted their leave utilization and therefore decreased inequalities in leave-taking.

[4] QPIP also increased fathers' leave participation, which may also have impacted breastfeeding. By increasing the proportion of fathers on leave, QPIP may have increased breastfeeding by providing mothers with additional support and increasing the proportion of fathers proficient in bottle-feeding breastmilk. Since the changes to fathers' leave behavior occurred at the same time as the changes to mothers' leave behavior, it is not possible to isolate their separate contributions to breastfeeding behavior. However it is reasonable to assume that the changes in average breastfeeding outcomes were driven mainly by increases in the average mothers' leave duration than fathers' duration. Furthermore, any reduction in breastfeeding inequalities under QPIP would certainly be driven by reductions in mothers' leave-taking inequalities, as I find in supplementary regressions that QPIP actually increased inequalities in fathers' leave-taking across income groups.

3.4 Data & Methods

3.4.1 Data on Leave Behavior

To investigate QPIP's impact on mothers' leave behavior, I use data on parental benefit claims that is collected by Statistics Canada through the Employment Insurance Coverage Survey (EICS) every year (Statistics Canada, 2002-2010). The target population for this annual survey is a subset of the target population for the Labor Force Survey, and comprises individuals who, given their recent status in the labor market, could potentially be eligible for employment insurance. Mothers of infants less than one year old, who I will focus on in my sample of potential leave-takers, fall into this last category since they could potentially be eligible for benefits via maternity or parental leave. I restrict my sample to a nine-year window framing the QPIP reform, from 2002 to 2010. Data from 2002-2005 thus comprises the pre-reform period (roughly 50.7% of the observations), and 2007-2010 as the post-reform period.[5] The main sample comprises 8,536 observations of mothers who have a child under one year old. Approximately 19% of my main sample are from Quebec, while the other observations come from the control group, which comprises the provinces of Ontario, Alberta, British Columbia, Atlantic Region, and Manitoba and Saskatchewan, where the EI system remained in place over the entire period of the analysis.

I analyze outcomes regarding leave behavior that are constructed as follows. Mothers' participation in paid parental leave is measured by an indicator taking value 1 if the mother has claimed or plans to claim maternity/parental/paternity benefits through either the EI

[5] I exclude data from 2001 and earlier because there were nation-wide reforms to both job-protected and paid parental leave in late 2000, as studied by Baker and Milligan (2008). I exclude data from 2006 because the public-use files do not record year of birth and it is impossible to identify births which occurred in 2005 (pre-reform) from those that occurred in 2006 (post-reform).

or QPIP system. Mothers' leave duration is a continuous measure of the total weeks of actual or planned leave taken by mothers who report claiming benefits, and taking the value 0 for mothers who do not claim benefits. The measure of leave duration is therefore not conditional on participation, offering a summary measure that takes into account both changes in participation and changes in duration conditional on participation.

There are two important issues to note about the data on mothers' leave duration. The first issue is that mothers who are on leave at the time of survey can only offer responses about planned leave duration, which may not be the length of leave they actually end up taking. However, since the EICS only covers mothers who have an infant under a year old, limiting our sample to mothers who have completed their leave spells would lead to the systematic over-representation of mothers who took shorter leaves and skew the distribution of leave durations to the left. Thus, since there is no reason to believe that mothers systematically under- or over-estimate their planned leave, I simply treat duration of leave to be length of completed leave for those who have returned, and length of planned leave for mothers still on leave. The second issue is that the EICS survey asks new mothers about the duration of all leave (not specifically paid parental leave) taken by them in the last year, and could capture unpaid leave or paid sick or vacation leave. Since my aim is to study mothers' participation in a paid parental leave program, I address this problem as follows. I assign a leave duration of 0 to mothers who do not ever claim benefits, ensuring that we do not count any unpaid leave or sick leave etc taken by mothers who do not qualify for benefits. Mothers who do qualify for benefits could in theory use sick days or vacation days in lieu of parental leave, but given the generous paid parental leave available and the lack of stigma to maternal leave-taking, this is very unlikely to have been the case as long as parental leave is available to them. It is however possible that mothers who qualified for benefits used their sick leave and vacation days to supplement the paid parental leave once they had exhausted these benefits. Accordingly, for the small proportion of mothers who report planning/taking leave for longer than the total possible weeks of parental leave benefits (52 weeks), I right-censor these observations at 52 weeks, implicitly assuming that mothers will exhaust their paid parental leave before considering using alternative types of leave.

Table 3.1 presents summary statistics for the EICS data sample, as well as difference-in-differences in sample characteristics over provinces and time. The data provides details on family characteristics such as mother's age, spouse's age, immigrant status, household size, and number of children etc. It should be noted I find no statistically significant double-differences in these characteristics, suggesting the sample composition did not change uniquely in Quebec in the post-reform period. Table 3.1 also provides information on the mothers' legal marital status, whether the household falls into four different income categories (below 20,000CAD, 20,000CAD-40,000CAD, 40,000CAD-60,000CAD, 60,000CAD+), and the highest education level attained by the mother. I use this information to create sub-samples study the effects of QPIP on mothers' leave behavior across marital status, income strata, and education levels.

3.4.2 Data on Breastfeeding

To investigate the impact of QPIP on health outcomes I use restricted-access data from the Canadian Community Health Survey (CCHS) from the years 2005 and 2007 - 2013 (Statistics Canada, 2005-2010).[6] The CCHS is a sample survey with cross-sectional design that collects information related to health status, health care utilization and health determinants for the Canadian population. The survey used to be collected every 2 years, but beginning in 2007 became an annual survey. The full sample comprises 24,386 mothers aged 18-55 who

[6] I use the restricted-access master files because the public use microdata files do not provide the year of the child's birth and therefore cannot cleanly identify births as being pre- or post-treatment.

have experienced a birth in the last 5 years. I additionally restrict the sample to mothers who have experienced a birth on or after 1st January 2002. Approximately 20% of the respondents reside in the treatment province of Quebec, while the other 80% reside in the control group that comprises Ontario, Alberta, British Columbia, Atlantic Region, Manitoba and Saskatchewan, Newfoundland and Labrador, Nova Scotia, Nunavut and Territories, New Brunswick and Prince Edward Island., which I treat as a control group. If the mother reports the birth to have occurred in the year 2005 or earlier, the child was born under the old EI program and is not considered treated. A mother who reports a birth to have occurred on or after 1st January 2006 is considered an observation from the post-reform period. Mothers may have moved between provinces in the last 5 years, which would confuse treatment status, meaning my estimates are an undervaluation of the true program effect.[7]

The measures of breastfeeding outcomes are constructed as follows. Breastfeeding initiation is measured by an indicator variable taking the value 1 if the mothers reports that she breastfed or tried to breastfeed her last child and takes the value 0 otherwise. For my analyses of breastfeeding initiation rates, I use the full sample of mothers with children aged 1-5 since all mothers are asked this question in the CCHS. However, the CCHS survey questions surrounding duration of breastfeeding and exclusive breastfeeding is more complicated and my construction of the measures and my samples must take into account two main issues. The first issue is that when asked about the duration of their breastfeeding, respondents are offered a choice of several time-intervals, but these intervals are not equally spaced and are not consistent across survey years. To overcome this issue, I identify the intervals which can be consistently tracked across the survey years, and construct various indicator variables to describe whether a mother has breastfed for at least a certain period of time. In my analyses

[7]The proportion of people moving in and out of Quebec in any given year is small. For example, using inter-provincial migration rates for 2007 reported by Milan (2011), I calculate that the proportion of people migrating into and out of Quebec was 0.26% and 0.41% respectively of Quebec's population. Therefore, the number of parents with young children moving in and out of Quebec and therefore confusing treatment in my GSS sample is expected to be small and unlikely to bias the results significantly.

of breastfeeding duration I therefore use six indicator variables taking the value of 0 or 1 depending on whether the respondent breastfed for at least 3 months, at least 6 months, at least 1 year, and whether the respondent exclusively breastfed for at least 12 weeks, at least 20 weeks, and at least 28 weeks. It should be noted that these measures are not conditional on breastfeeding initiation.

The second issue with the CCHS is that it only asks mothers for the duration of completed breastfeeding and therefore excludes mothers who may still be breastfeeding at the time of the survey. To minimize bias, for my analyses of breastfeeding durations I restrict the sample to mothers whose youngest child is at least 1 year old. Including mothers with children aged 0-1 would only have reported durations for mothers who have completed breastfeeding and have missing values for mothers who are still breastfeeding, thus skewing the distribution of breastfeeding durations in my sample to the left.[8] It should also be noted that, since mothers often stop exclusively breastfeeding before they stop breastfeeding altogether, I have more non-missing observations for completed durations of exclusive breastfeeding than for completed durations of any breastfeeding.

It should be noted that since I cannot track mothers from their leave-taking behavior to their later breastfeeding outcomes, I essentially focus on changes in breastfeeding behavior among mothers who were exposed to the QPIP program rather than the EI Program. The relationships I identify between QPIP and breastfeeding behaviors should therefore be interpreted as intent-to-treat effects rather than the effects of treatment-on-the-treated. After presenting the ITT effects in Section 3.5, I will also present and discuss estimates for the upper bound of the TOT effects.

[8]Although some mothers do continue to breastfeed after a year, and thus would still have missing values for duration and not be represented in my sample, the proportion of these mothers in the population is under 7% and should introduce negligible bias.

In addition, the CCHS contains information on various personal and household characteristics such as the age of the mother, household size, number of children, number of children under the age of 5, immigrant status, and legal marital status, mothers' education, personal annual income, and household annual income. The mean characteristics for the control group and treatment group can be seen in Table 3.2. It is worth noting that I use the information on marital status, education level, personal income and household income to divide the sample into various sub-groups in order to analyze how the effects of QPIP may have differed across the population. To stratify the sample by marital status, I compare mothers who are legally married or cohabitating in a common-law marriage to mothers who have never been married.[9]

To stratify the sample by education, I divide mothers into three groups: those with at most a high school education, those who have some post-secondary education but not a college degree, and those who have a BA or more advanced degree. To stratify the sample based on personal income, I divide the mothers into four groups based on whether their reported personal income for the last year was below 10,000CAD, between 10,000CAD-20,000CAD, between 20,000CAD-30,000CAD, or above 30,000CAD. These income groups correspond roughly to: below the 25th percentile, between 25th-45th percentile, between the 45th-65th percentile, or above the 65th percentile of the distribution of personal incomes for the full sample of mothers. In a similar fashion, to stratify the sample based on household income, I divide the mothers into four groups based on whether their household income for the last year was below 30,000CAD, between 30,000CAD-50,000CAD, between 50,000CAD-80,000CAD, or above 80,000CAD. These income groups correspond roughly to: below the 15th percentile, between 15th-30th percentile, between the 30th-65th percentile, or above the 65th percentile of the distribution of household incomes for the full sample of mothers.[10]

[9] It is reasonable to combine married and common-law married into a single group as there is a very high incidence of long-term committed cohabitation in Quebec. As a check, in Appendix Table C.1 I provide the regression results for the samples of single mothers, cohabitating mothers, and married mothers separately.

[10] These numbers are approximate; since personal and household incomes are reported only as intervals

3.4.3 Empirical Methods

I utilize a difference-in-differences identification strategy which exploits the fact that all provinces in Canada utilized the same Employment Insurance Program for parental leave benefits from 2002-2005, and that only the province of Quebec reformed its benefit offerings in 2006. For every separate sample I explore QPIP's program effect on a particular outcome, estimating:

$$Y_{ijt} = \alpha + \beta I[j = Quebec] * I[t >= 2006] + \theta I[t >= 2006] + \phi Z_{ijt} + \lambda_j + \delta_t + \epsilon_{ijt} \tag{3.1}$$

where subscript i denotes the individual, subscript j denotes province and subscript t denotes the year of last birth. Y_{ijt} therefore represents the outcome of mother i observed in province j who gave birth in year t. $I[t >= 2006]$ is an indicator variable taking the value 1 if the birth-year t is 2006 or greater, i.e., if the observation is from the post-reform period. The coefficient θ represents the change in the value of the outcome that is shared by all provinces. The term $I[j = Quebec] * I[t >= 2006]$ takes the value of 1 if the individual lives in Quebec and gave birth in a post-reform year, and otherwise takes the value 0. The coefficient β therefore represents the DD estimate of primary interest as it captures the change in the value of the outcome post-reform that is unique to Quebec. Under the assumption that no other policy changes were enacted to affect it, β represents QPIP's average treatment effect. λ_j and δ_t denote the fixed province and year effects. It should be noted that I do not control for all province-year interactions, but instead collapse them into the term $I[j = Quebec] * I[t >= 2006]$. [11]

in the CCHS, it is impossible to create subsamples corresponding to exact percentiles of these income distributions

[11] I do not include controls for province-specific time trends as these would be highly collinear with the

All regression specifications include controls for personal characteristics, represented by the term Z_i. In the EICS data these characteristics include age, legal marital status, immigrant status, family size, and number of children. In the CCHS data these characteristics include mother's age, household size, number of children, and age of youngest child. Including these as regressors control for changes in group composition. ϵ_i is the error term. I calculate cluster-robust standard errors that generalize the White (1980) heteroskedastic-consistent estimates of OLS standard errors to the clustered setting in order to account for possible heteroskedasticity and within-province dependence of standard errors, which are particularly a concern in difference-in-difference estimations since the regressor of interest is highly correlated within clusters (Bertrand et al., 2004). All analyses are conducted using ordinary least squares regressions despite the binary nature of some of the indicators because they resulted in very similar estimates as those from logit estimates.

3.5 Results

3.5.1 QPIP's Effects on Inequalities in Leave-taking

Table 3.3 presents results from difference-in-difference regressions that estimate QPIP's treatment effects on mothers' leave behavior. Using the full sample of mothers, Column 1 reports that QPIP was associated with an increase in mothers' participation rates of 16.3 p.p. and leave duration of over 7 weeks. These are both economically and statistically significant effects, representing 22% and 23% of the pre-reform means in Quebec for mothers' participation and leave duration respectively. Notably, the average treatment effects for this sample

program effect of QPIP. As one would expect, in supplementary regressions I confirm that the inclusion of a Quebec-specific time trend absorbs some of the program effects, leading to smaller but consistent point estimates of QPIP's impact on leave behavior.

are larger than those reported for my analysis in Chapter 1 (where I used a more restricted sample of married or cohabitating mothers aged 18-40), suggesting that QPIP's effects may have been larger for never-married mothers and teenage mothers etc.

In Columns 2-3, I consider whether QPIP had heterogeneous effects across women of different marital statuses. The pre-reform means for Quebec show that single mothers faced considerable disadvantage in their ability to take parental leave: compared to mothers who are married, single mothers not only exhibited lower participation rates (39.5% versus 75.2%), but also reported leaves of shorter duration (18.6 weeks versus 33.6 weeks). Table 3.3 shows that the gap was narrowed by the introduction of QPIP. Under QPIP, single mothers' participation rates increased by 22.1 p.p., while those of married mothers increased by 15.8 p.p.. Similarly, in terms of leave duration, single mothers' leave duration rose by 6.4 weeks and that of married mothers only by 7.4 weeks. The overall pattern shows that QPIP's program effects helped reduce the gap in leave utilization rates between married and unmarried mothers.

Next, I consider whether the QPIP program affected inequalities in leave-behavior across families from different income strata. To begin with, the pre-reform mean participation rates display a steep income gradient that favors wealthier families. For example, from 2002-2005 in Quebec, only 44% of mothers with annual household income blow 20,000CAD reported taking leave, compared to 88% of mothers with household incomes above 80,000CAD. Similarly, mothers' leave duration also was shortest for the lowest-income families, increasing as the annual household income rose. Columns 4-7 of Table 3.3 show that QPIP had heterogeneous effects across income strata that actually favored lower-income households. For example, though QPIP increased mothers' participation rates for all families, QPIP was associated with a 20.6 p.p. increase in participation for mothers with household income below 20,000CAD, compared to 18.5 p.p. for those with income of 20,000-40,000CAD, 15.9

p.p. for those with income of 40,000-60,000CAD, and only 6.6 p.p. for mothers from the highest-income households, earning above 60,000CAD annually. Mothers' leave duration also exhibits a similar pattern: QPIP increased mothers' leave duration by 7.8 weeks in households with income below 20,000CAD, by 9 weeks for households earning 20,000-40,000CAD, and 7.4 weeks for households earning 40,000-60,000CAD, while the highest income households only increased mothers' leave duration by 2.9 weeks. Thus, the introduction of QPIP was associated with a decrease in inequalities in mothers' leave taking across households of different income levels.

And lastly, I consider heterogeneous program effects across educational levels. The pre-reform means for mothers of different education levels in Quebec show that mothers who are not educated beyond high school take less leave compared to mothers who have some post-secondary education or a college degree. Unfortunately, QPIP's program effects may have exacerbated these differences. QPIP was associated with an increase in participation rates of 12.5 p.p. for high-school educated mothers, significantly less than the increase of 16.6 p.p. for mothers with some post-secondary education, and the statistically largest increase of 20.1 p.p. for college-educated mothers. Similarly, QPIP increased leave duration by only 3.9 weeks for high-school educated mothers, compared to the increase of 7.8 weeks for mothers with some secondary education and 9.6 weeks for college-educated mothers. When looking across mothers' education levels, QPIP did not reduce and in fact may have exacerbated inequalities as its program effects were largest for college-educated mothers.

To summarize, QPIP succeeded in improving mothers' leave utilization on average, but the magnitude of effects differed considerably across different groups of mothers. QPIP succeeded in reducing inequalities in leave-taking between single mothers and married mothers, and households across different income strata, but did not help mothers with only a high school degree bridge the gap with mothers who are more educated.

3.5.2 QPIP's Effects on Inequalities in Breastfeeding

Table 3.4 presents estimates of the average treatment effects of QPIP on mothers' breastfeeding behaviors. Column 1 shows that on average, mothers who were exposed to QPIP were 4.6 p.p more likely to initiate breastfeeding, compared to mothers who are not exposed. This represents a significant increase of 6% from the pre-reform average for Quebec. Columns 2-7 also show that QPIP significantly increased the length of breastfeeding among mothers, especially at longer breastfeeding durations. Compared to mothers who gave birth under the EI program, mothers who gave birth under QPIP were 5.6 p.p. more likely to breastfeed at least 3 months, 4.1 p.p. more likely to breastfeed at least 6 months, and 2.2 p.p. more likely to breastfeed for a full year. Additionally, mothers exposed to QPIP were 4.8 p.p. more likely to exclusively breastfeed for at least 12 weeks and 5.9 p.p. more likely to exclusively breastfeed for at least 20 weeks. There was no program effect on the probability of exclusively breastfeeding for 28 weeks or longer.

Table 3.5 presents regression results for sub-samples according to mothers' marital status. Unlike the results for leave-taking, I find QPIP's program effects on breastfeeding favored married mother rather than never-married mothers. Under QPIP, married mothers increased breastfeeding initiation rates by 4.4 p.p. Married mothers' likelihood of breastfeeding for at least 3 months increased by 6.4 p.p, and at least 6 months by 5.7 p.p, and at least 1 year by 2.4 p.p. In comparison, I cannot detect any statistically significant effects on the breastfeeding behaviors of never-married mothers. Therefore, although QPIP did boost leave utilization particularly among never-married mothers, this did not translate into improved breastfeeding outcomes among never-married women.

Table 3.6 presents regression results for sub-samples according to the household's reported annual income. Under QPIP, households earning below 30,000CAD increased breastfeeding initiations by 4.9 p.p., while in households from the three higher income groups QPIP increased initiation rates by a little over 3 p.p. For measures of breastfeeding duration (Columns 2-4), I find that QPIP had the most significant program effects on the highest-income households, followed by the lowest- income households - with small or no effects for middle-income households. The probability of breastfeeding at least 3 months increased by 5.9 p.p. in households earning below 30,000CAD, and by 9.6 p.p. in those earning more than 80,000CAD. The probability of breastfeeding at least 6 months increased by 5.4 p.p. in households earning less than 30,000CAD, by 5.6 p.p. in those earning 50,000-80,000CAD, and by 8.7 p.p. in the highest-income households. At the longest measure of breastfeeding duration, i.e. a full year, QPIP only has a program effect for mothers from the highest-income households. Prior to the reform, these mothers had the lowest likelihood of breastfeeding a full year, but under QPIP that likelihood rose by 5.3 p.p.. For measures of exclusive breastfeeding, the poorest households experience larger gains. In the likelihood of exclusively breastfeeding at least 12 weeks, QPIP leads to an increase of 12.6 p.p. for households earning less than 30,000CAD and of 6.4 p.p. for those earning over 80,000CAD. In the likelihood of exclusively breastfeeding at least 20 weeks, QPIP leads to an increase of 8.5 p.p. for households earning less than 30,000CAD, 7.9 p.p. for households earning 50,000-70,000CAD, and 4.4 p.p. for those earning over 70,000CAD. The overall pattern indicated improvements in breastfeeding durations for the poorest and richest households. The results for inequality are therefore mixed, and do not uniformly mirror the pattern of QPIP's program effects on leave-taking, which were largest for the poorest households and gradually decreased along the income gradient. QPIP's program effects on exclusive breastfeeding favored the poorest households, but the effects on breastfeeding durations favored the richest households. It is good news that QPIP shifted the overall distribution of breastfeeding to the right, but unfortunately it is not possible to say that QPIP narrowed the distribution per se.

Table 3.7 presents regression results for different sub-samples according to the mothers' education level. The results are disappointing. Under QPIP, breastfeeding initiation rates rose by 8 p.p. for mothers with no more than a high school diploma, compared to 2.7 p.p.for those with some post-secondary education and 7.6 p.p. for those with at least a Bachelor's degree. However, for every measure of breastfeeding duration and exclusive breastfeeding duration, QPIP's program effects were much smaller for mothers with only a high school degree, compared to more educated mothers. Mothers with only a high-school education experienced no statistically significant change in breastfeeding duration under QPIP, while those with some post-secondary education became 3.3 p.p. more likely to breastfeed 3 months and 7.5 p.p. more likely to breastfeed 6 months. Mothers with college-degrees experienced even bigger increases - under QPIP, they became 13.5 p.p. more likely to breastfeed 3 months, 9.9 p.p. more likely to breastfeed 6 months, and 5.4 p.p. more likely to breastfeed a full year. Similarly, in the length of exclusive breastfeeding, high-school educated mothers became 3.2 p.p. more likely to achieve the threshold of 12 weeks under QPIP, but that effect is small in comparison to the gains experienced by more educated mothers. Mothers with some post-secondary education are 8.4 p.p. more likely to exclusively breastfeed 20 weeks and 2.1 p.p. more likely to exclusively breastfeed 28 weeks. College-educated mothers are 12.8 p.p. more likely to exclusively breastfeed 12 weeks and 9.7 p.p. more likely to exclusively breastfeed 20 weeks. Overall, for every measure of breastfeeding duration QPIP's program effects increased educational inequalities by favoring higher-educated mothers.

The regression results reported in Tables 3.4-3.7 should be interpreted as QPIP's ITT (Intent To Treat) effects on breastfeeding, as they do not connect individual mothers' leave behavior to their subsequent breastfeeding outcomes. However, it is possible to derive estimates for the TOT (Treatment Effects on the Treated) effects by dividing the ITT effects on breastfeeding by the 1st stage effects on leave take-up. Accordingly, Table 3.8 provides estimates for the upper bound of the TOT effects of QPIP.[12] Table 3.8 clearly shows the same pattern of breastfeeding behaviors that we saw in Tables 3.4-3.7. QPIP had a large TOT effect on the average probability of breastfeeding, increasing it by 12.7 percentage points, and increased the probability of attaining many critical breastfeeding duration thresholds. However, QPIP's TOT reveal a clear pattern of increasing inequality. First, the TOT effects were large and significant for married/cohabitating mothers, and statistically insignificant for never-married mothers. Second, the TOT effects were larger for the highest-income households than lower-income households. And lastly, for any measure of breastfeeding duration, QPIP's TOT effects were larger for mothers who have some post-secondary education, while those who only attended high school experienced the smallest TOT effects.

3.6 Conclusion

This paper finds that a lowering of eligibility requirements and an increase in financial compensation for parental leave, without any change in the duration of job protection or benefits, not only raised the average mother's leave utilization but also reduced the income gradient to leave participation and narrowed the gap between single and married mothers. I find that the QPIP reform raised average breastfeeding rates, confirming the idea that lowering the opportunity cost of time at home can facilitate more breastfeeding. However, I also find that although QPIP reduced some inequalities in leave-taking, it was not successful in reducing inequalities in breastfeeding behaviors. This suggests that policies which lower the opportunity costs of breastfeeding are less effective for families of low socio-economic status, who have lower opportunity costs to begin with, and likely also have lower latent

[12] The estimates derived by dividing the ITT effects by the 1st stage effects on leave take-up should be considered upper bounds for the TOT effects because QPIP may have had feedback effects on breastfeeding. That is, in addition to the direct effects on breastfeeding of mothers who took more parental leave under QPIP, the program may have had feedback effects on all Quebecois mothers' breastfeeding behavior by changing social norms and/or institutions.

demand for breastfeeding.

These findings have important implications for the way we interpret some of the prior research on the subject. When we consider previous studies that found positive effects of parental leave programs on infant health only for college-educated, high-income populations (e.g. Rossin (2011)), we should be cautious in laying the blame on the policy design for being exclusionary. My findings suggest that even if these leave programs had been more accessible and inclusive, they are unlikely to have eliminated inequalities in health outcomes, as more advantaged families seem to respond more strongly to reductions in opportunity costs.

The results of this study also have important implications for the design of future policy. They confirm that a good way to raise the average breastfeeding rate in a nation is to offer paid parental leave and facilitate mothers' time at home. However, they also suggest that policy-makers who are concerned with health inequalities across society should not think of parental leave as a panacea. Programs that make leave accessible for more families may be necessary, but not sufficient, for bringing equity to the issue of mothers' milk. Policy-makers must pay increased attention on policies that will increase the latent demand for breastfeeding among vulnerable populations, .e.g., through informational campaigns about the benefits of breastfeeding, pre- and post-natal training with breastfeeding techniques, or efforts to change attitudes within mothers' support networks.

Table 3.1: Sample Means for Mothers in the EICS Data

	(1) Control Provinces 2002-2005	(2) Control Provinces 2006-2013	(3) Treatment Province 2002-2005	(4) Treatment Province 2006-2013	(5) Difference in Difference
Proportion of Mothers aged 18-24	0.160	0.129	0.158	0.121	-0.028
Proportion of Mothers aged 25-44	0.839	0.871	0.842	0.879	0.028
Family Size	3.707	3.802	3.707	3.755	-0.005
Proportion of first-time mothers	0.463	0.409	0.444	0.429	0.009
Immigrant	0.211	0.203	0.112	0.176	0.035
Proportion of married/cohabiting mothers	0.916	0.918	0.937	0.953	0.017
Household Income: below 20,000CAD	0.145	0.123	0.168	0.134	-0.016
Household Income: 20,000CAD-40,000CAD	0.389	0.316	0.431	0.357	-0.007
Household Income: 40,000CAD-60,000CAD	0.245	0.216	0.204	0.229	**0.051**
Household Income: 60,000CAD+	0.221	0.344	0.197	0.280	-0.028
Educ: high school diploma or less	0.257	0.224	0.171	0.173	-0.001
Educ: some post-secondary schooling	0.417	0.416	0.539	0.462	**-0.034**
Educ: Bachelors' or more advanced degree	0.322	0.356	0.288	0.362	**0.037**

Notes: Sample spans 2002-2010 of the EICS data and comprises mothers aged 15-55 who have experienced a birth in the last year. Treatment province refers to Quebec, while control group comprises the other regions of Ontario, Alberta, British Columbia, Atlantic Region, and Manitoba and Saskatchewan. Column 5 shows difference-in-differences across provinces and time while controlling for province and year-fixed effects. Double differences in bold are significant at the 1% level

Table 3.2: Sample Means for Mothers in the CCHS Data

	(1) Control Provinces 2002-2005	(2) Control Provinces 2006-2013	(3) Treatment Province 2002-2005	(4) Treatment Province 2006-2013	(5) Difference in Difference
Age of Mother	32.82	31.946	32.482	31.924	0.034
Household Size	3.931	3.911	3.682	3.763	**0.105**
Number of children	1.841	1.807	1.742	1.777	**0.065**
Number of children aged 1-5	1.248	1.338	3.117	3.343	0.017
Married or Cohabitating	0.830	.834	0.842	0.867	0.026
Household Income: below 30,000CAD	0.161	0.173	0.168	0.155	-0.026
Household Income: 30,000CAD-50,000CAD	0.162	0.227	0.206	0.139	**-0.051**
Household Income: 50,000CAD-80,000CAD	0.269	0.230	0.310	0.300	0.029
Household Income: 80,000CAD+	0.408	0.443	0.316	0.400	**0.051**
Educ: high school diploma or less	0.234	0.257	0.161	0.168	-0.013
Educ: Some post-secondary schooling	0.474	0.432	0.530	0.489	0.001
Educ: Bachelors' or more advanced degree	0.292	0.310	0.307	0.341	0.012

Notes: Sample spans births from 2002-2013 recorded in the 2005 and 2007-2013 waves of the CCHS survey, including mothers aged 18-55 who have experienced a birth in the last 5 years. Treatment province refers to Quebec, while control group comprises the other regions of Ontario, Alberta, British Columbia, Atlantic Region, Manitoba and Saskatchewan, Newfoundland and Labrador, Nova Scotia, Nunavut and Territories, New Brunswick and Prince Edward Island. Column 5 shows difference-in-differences across provinces and time while controlling for province and year-fixed effects. Double differences in bold are significant at the 1% level

Table 3.3: QPIP's Average Treatment Effects on Mothers' Leave Behavior

Sample Based on:		Marital status		Annual Household Income Reported for Last year				Mothers' Reported Education Level		
	(1) Full Sample	(2) Single	(3) Married/ Cohabiting	(4) Income below 20K	(5) Income 20K-40K	(6) Income 40K-60K	(7) Income 80K+	(8) High school or less educ	(9) Some post-secondary ed	(10) BA or more Adv. Degree
I. Outcome: Probability of mother claiming Parental Leave Benefits										
Average Treatment	0.163***	0.221***	0.158***	0.206**	0.185***	0.159***	0.066_s***	0.125_s***	0.166***	0.201_b***
Effect of QPIP	(0.00)	(0.00)	(0.00)	(0.02)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
No. of Observations	8536	681	7855	1340	3135	1964	2096	2049	3924	2536
Pre-reform Mean	0.733	0.395	0.752	0.438	0.733	0.831	0.884	0.605	0.761	0.758
II. Outcome: Actual/Planned Duration of all leave taken by mother after childbirth										
Average Treatment	7.424***	6.416**	7.414***	7.850**	9.018***	7.358***	2.904_s**	3.962_s***	7.798***	9.662_b***
Effect of QPIP	(0.00)	(0.03)	(0.00)	(0.04)	(0.00)	(0.00)	(0.02)	(0.00)	(0.00)	(0.00)
No. of Observations	8536	681	7855	1340	3135	1964	2096	2049	3924	2536
Pre-reform Mean	32.781	18.672	33.563	19.884	32.598	37.042	39.797	27.611	34.013	33.639

Robust province-clustered p-values in parentheses, * p<0.10 ** p<0.05 *** p<0.01

Subscripts b & s denote coefficients significantly bigger or smaller than coefficients for relevant comparison groups

Notes: Table presents results from regressions exploring the program effects of QPIP by exploiting variation in time and province in difference-in-difference regressions. Regressions control for province- and year- fixed effects as well as personal covariates such as age, household size, number of children in various age groups. Sample spans 2002-2010 of the Employment Insurance Coverage Survey data and comprises mothers aged 15-55 who have a child aged 1-5 years old (or in the case of breastfeeding initiation, 0-5 years old). To compare whether regression coefficients for different subgroups are statistically different from each other, please see Z-statistics in Appendix Table C.2.

Table 3.4: QPIP's Average Treatment Effects on Mothers' Breastfeeding Behavior

OUTCOMES:	(1) Breastfed / Tried Breastfeeding	(2) Breastfed at least 3 months	(3) Breastfed at least 6 months	(4) Breastfed at least 1 year	(5) Breastfed exclusively at least 12 weeks	(6) Breastfed exclusively at least 20 weeks	(7) Breastfed exclusively at least 28 weeks
Average Treatment	0.046***	0.056***	0.041***	0.022**	0.048***	0.059***	0.005
Effect of QPIP	(0.00)	(0.00)	(0.00)	(0.04)	(0.00)	(0.00)	(0.29)
No. of Observations	24386	20048	20048	20048	21515	21515	21515
Pre-reform Mean	0.811	0.522	0.379	0.096	0.487	0.235	0.067

Robust province-clustered p-values in parentheses, * p<0.10 ** p<0.05 *** p<0.01

Notes: Table presents results from regressions exploring the the program effects of QPIP by exploiting variation in time and province in difference-in-difference regressions. Regressions control for Personal Covariates such as mothers' age, household size, number of children in various age groups, age of youngest child, and Province & Year- Fixed Effects. The data is from the 2005 and 2007-2013 waves of the Canadian Community Survey, and the sample comprises mothers aged 18-55 who have a child aged 1-5 years old (or in the case of breastfeeding initiation, 0-5 years old) who was born in or after 2002.

Table 3.5: QPIP's Effects on Breastfeeding Behavior Across Marital Statuses

OUTCOMES:	(1) Breastfed / Tried Breastfeeding	(2) Breastfed at least 3 months	(3) Breastfed at least 6 months	(4) Breastfed at least 1 year	(5) Breastfed exclusively at least 12 weeks	(6) Breastfed exclusively at least 20 weeks	(7) Breastfed exclusively at least 28 weeks
Sample A: Mothers who have never been married							
Average Treatment	0.029	-0.012$_s$	0.022$_s$	0.058	-0.041	-0.011$_s$	-0.022
Effect of QPIP	(0.40)	(0.77)	(0.43)	(0.11)	(0.33)	(0.77)	(0.43)
No. of Observations	3176	2741	2741	2741	2805	2805	2805
Pre-reform Mean	0.698	0.362	0.267	0.067	0.371	0.195	0.062
Sample A: Mothers who are married or in a common-law marriage							
Average Treatment	0.044***	0.064$_b$***	0.057$_b$***	0.024***	0.051***	0.059$_b$***	0.003
Effect of QPIP	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.44)
No. of Observations	19821	16027	16027	16027	17400	17400	17400
Pre-reform Mean	0.824	0.542	0.395	0.092	0.504	0.241	0.070

Robust province-clustered p-values in parentheses, * p<0.10 ** p<0.05 *** p<0.01
Subscripts b & s denote coefficients significantly bigger or smaller than coefficients for all other comparison groups

Notes: Table presents results from difference-in-difference regressions exploring the the program effects of QPIP by exploiting variation in time and province for two separate samples: mothers who have never been married, and mothers who are legally married or in a common law marriage. Regressions control for Personal Covariates such as mothers' age, household size, number of children in various age groups, age of youngest child, and Province & Year- Fixed Effects. The data is from the 2005 and 2007-2013 waves of the Canadian Community Survey, and the sample comprises mothers aged 18-55 who have a child aged 1-5 years old (or in the case of breastfeeding initiation, 0-5 years old) who was born in or after 2002. To compare whether regression coefficients for different subgroups are statistically different from each other, please see Z-statistics in Appendix Table C.2.

Table 3.6: QPIP's Effects on Breastfeeding Behavior across Household Income Groups

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
OUTCOMES:	Breastfed / Tried Breastfeeding	Breastfed at least 3 months	Breastfed at least 6 months	Breastfed at least 1 year	Breastfed exclusively at least 12 weeks	Breastfed exclusively at least 20 weeks	Breastfed exclusively at least 28 weeks
Sample A: Mothers with Household income $\leq$ 30,000CAD							
Average Treatment Effect of QPIP	0.049**	0.059***	0.054*	0.012	0.126_b***	0.086_b***	0.017
	(0.02)	(0.00)	(0.06)	(0.33)	(0.00)	(0.00)	(0.25)
No. of Observations	5792	4439	4439	4439	5025	5025	5025
Pre-reform Mean	0.748	0.438	0.309	0.115	0.391	0.186	0.062
Sample B: Mothers with household income of 30,000CAD-50,000CAD							
Average Treatment Effect of QPIP	0.038***	0.017	0.006	0.049	-0.009	0.036	-0.002
	(0.00)	(0.63)	(0.90)	(0.13)	(0.69)	(0.28)	(0.78)
No. of Observations	3789	3105	3105	3105	3339	3339	3339
Pre-reform Mean	0.727	0.473	0.368	0.093	0.451	0.219	0.077
Sample C: Mothers with household income of 50,000CAD-80,000CAD							
Average Treatment Effect of QPIP	0.031***	0.007	0.056***	-0.032	0.000	0.079***	0.003
	(0.00)	(0.45)	(0.00)	(0.15)	(0.94)	(0.00)	(0.88)
No. of Observations	6011	4889	4889	4889	5321	5321	5321
Pre-reform Mean	0.853	0.552	0.355	0.107	0.526	0.232	0.050
Sample D: Mothers with household income $\geq$ 80,000CAD							
Average Treatment Effect of QPIP	0.037***	0.096_b***	0.087***	0.053***	0.064**	0.044***	0.001
	(0.00)	(0.00)	(0.00)	(0.00)	(0.02)	(0.00)	(0.84)
No. of Observations	8794	7315	7315	7315	7830	7830	
Pre-reform Mean	0.873	0.595	0.441	0.084	0.524	0.244	0.076

Robust province-clustered p-values in parentheses, * p<0.10 ** p<0.05 *** p<0.01
Subscripts b & s denote coefficients significantly bigger or smaller than coefficients for comparison groups

Notes: Table presents results from difference-in-difference regressions exploiting variation in time and province for four separate samples based on household's reported annual income for the previous year. Regressions control for Personal Covariates such as mothers' age, household size, number of children, age of youngest child, and Province & Year- Fixed Effects. The data is from the 2005 and 2007-2013 waves of the Canadian Community Survey, and the sample comprises mothers aged 18-55 who have a child aged 1-5 years old (or in the case of breastfeeding initiation, 0-5 years old) who was born in or after 2002. To compare whether regression coefficients for different subgroups are statistically different from each other, please see Z-statistics in Appendix Table C.2.

Table 3.7: QPIP's Effects on Breastfeeding Behavior across Education Groups

OUTCOMES:	(1) Breastfed / Tried Breastfeeding	(2) Breastfed at least 3 months	(3) Breastfed at least 6 months	(4) Breastfed at least 1 year	(5) Breastfed exclusively at least 12 weeks	(6) Breastfed exclusively at least 20 weeks	(7) Breastfed exclusively at least 28 weeks
Sample A: Mothers with at most high school education							
Average Treatment Effect of QPIP	0.080***	0.022	-0.005_s	-0.023_s	0.032**	-0.015_s	-0.028*
	(0.00)	(0.41)	(0.84)	(0.40)	(0.02)	(0.30)	(0.06)
No. of Observations	6162	5143	5143	5143	5403	5403	5403
Pre-reform Mean	0.605	0.304	0.193	0.056	0.300	0.142	0.064
Sample B: Mothers with some post-Secondary education							
Average Treatment Effect of QPIP	0.027***	0.033**	0.075***	0.032	0.032	0.084***	0.021*
	(0.00)	(0.02)	(0.00)	(0.12)	(0.12)	(0.00)	(0.09)
No. of Observations	11767	9799	9799	9799	10383	10383	10383
Pre-reform Mean	0.829	0.532	0.355	0.084	0.491	0.221	0.064
Sample C: Mothers with BA or more advanced degrees							
Average Treatment Effect of QPIP	0.076***	0.135_b***	0.099***	0.054***	0.128_b***	0.097***	0.004
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.01)	(0.55)
No. of Observations	7165	5674	5674	5674	6305	6305	6305
Pre-reform Mean	0.883	0.641	0.534	0.142	0.569	0.294	0.073

Robust province-clustered p-values in parentheses, * p<0.10 ** p<0.05 *** p<0.01
Subscripts b & s denote coefficients significantly bigger or smaller than coefficients for all other comparison groups

Notes: Table presents results from difference-in-difference regressions exploring the the program effects of QPIP by exploiting variation in time and province for three separate samples based on the reported education level attained by the mother. Regressions control for Personal Covariates such as mothers' age, household size, number of children in various age groups, age of youngest child, and Province & Year- Fixed Effects. The data is from the 2005 and 2007-2013 waves of the Canadian Community Survey, and the sample comprises mothers aged 18-55 who have a child aged 1-5 years old (or in the case of breastfeeding initiation, 0-5 years old) who was born in or after 2002. To compare whether regression coefficients for different subgroups are statistically different from each other, please see Z-statistics in Appendix Table C.2.

Table 3.8: Upper Bounds for QPIP's 'Treatment Effects on the Treated'

OUTCOMES:	(1) Breastfed / Tried Breastfeeding	(2) Breastfed at least 3 months	(3) Breastfed at least 6 months	(4) Breastfed at least 1 year	(5) Breastfed exclusively at least 12 weeks	(6) Breastfed exclusively at least 20 weeks	(7) Breastfed exclusively at least 28 weeks
Full Sample							
All Mothers	0.127***	0.343***	0.251***	0.136**	0.294***	0.362***	0.031
	(0.00)	(0.00)	(0.00)	(0.04)	(0.00)	(0.00)	(0.29)
Samples Stratified by Relationship Status							
Never-married	0.131	-0.049	0.099	0.262	0.185	0.050	0.099
Mothers	(0.40)	(0.77)	(0.43)	(0.11)	(0.33)	(0.77)	(0.43)
Married/Cohabitating	0.279***	0.405***	0.361**	0.152**	0.323***	0.373***	0.019
Mothers	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.44)
Samples Stratified by Annual Household Income							
Below 30,000CAD	0.233**	0.285***	0.257*	0.057	0.60***	0.405***	0.081
Household Income	(0.02)	(0.00)	(0.06)	(0.33)	(0.00)	(0.00)	(0.25)
30,000-50,000CAD	0.205***	0.092	0.032	0.265	0.048	0.195	0.011
Household Income	(0.00)	(.63)	(0.90)	(0.13)	(0.69)	(0.28)	(0.78)
50,000-80,000CAD	0.195***	0.044	0.352	-0.201	0.000	0.497	0.018
Household Income	(0.00)	(.45)	(0.00)	(0.15)	(0.94)	(0.00)	(0.88)
80,000CAD+	0.560***	1.454***	1.318***	0.803***	0.969**	0.666***	0.015
Household Income	(0.00)	(.00)	(0.00)	(0.00)	(0.02)	(0.00)	(0.84)
Samples Stratified by Mothers' Education							
High School Degree	0.640***	0.176	0.04	0.184	0.254**	0.120	-0.224*
Or less	(0.00)	(0.41)	(0.84)	(0.40)	(0.02)	(0.30)	(0.06)
Some Post-Secondary	0.163***	0.199**	0.452***	0.193	0.193	0.506***	0.124*
Education	(0.00)	(0.02)	(0.00)	(0.12)	(0.12)	(0.00)	(0.09)
BA or more	0.378***	0.675***	0.493***	0.269***	0.637***	0.483***	0.019
Advanced Degree	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.01)	(0.55)

Robust province-clustered p-values in parentheses, * $p<0.10$ ** $p<0.05$ *** $p<0.01$

Notes: Table presents upper bound estimates of treatment effects on the treated, calculated by dividing the intent-to-treat effects reported in Tables 3.4-3.7 by the first stage effects on leave participation reported in Table 3.3.

APPENDIX A

APPENDIX OF CHAPTER 1

Table A.1: Parametric RD Analyses of Quebec

	Fathers' Participation Rates	Fathers' Leave Duration (Weeks)	Mothers' Participation Rates	Mothers' Leave Duration (Weeks)
I. Bandwidth = 24 months				
RD Estimate	0.521***	3.611**	-0.022	-0.321
	[0.00]	[0.03]	[0.78]	[0.95]
N	806	786	806	617
II. Bandwidth = 18 months				
RD Estimate	0.500***	3.339	0.072	1.757
	[0.00]	[0.11]	[0.36]	[0.72]
N	606	590	606	464
III. Bandwidth = 12 months				
RD Estimate	0.601***	4.701**	0.245	4.611
	[0.00]	[0.01]	[0.62]	[0.53]
N	396	391	396	305

*p<0.10 **p<0.05 ***p<0.01, Heteroskedasticity-robust p-values in brackets

Notes: Table present results from Parametric RD regressions on treatment province of Quebec to detect discontinuities in outcomes at point of cutoff. Sample comprises births in months surrounding the reform, with the window varying from 12 to 18 to 24 months on either side of 1st January 2006. For each outcome, I determined the appropriate functional form using the F-test method suggested by Lee and Lemieux (2010) to test different candidate models against the underlying data. In most cases, this method suggested the use of a quadratic model with interactions between the running variable and the indicator for the post-reform period. Regressions include controls for Personal covariates e.g. age and education of mother and spouse as well as household size, religion and number of children in various age groups. Errors are allowed to cluster around the month of birth. Robust p-values presented in square brackets.

Table A.2: Non-Parametric RD Analysis of Quebec, Trimming around cutoff

	(1) Fathers' Participation Rates	(2) Fathers' Leave Duration (Weeks)	(3) Mothers' Participation Rates	(4) Mothers' Leave Duration (Weeks)
Jump at Cutoff in Quebec	0.531***	1.855	0.194*	4.888
	[0.00]	[0.21]	[0.08]	[0.12]
Bandwidth (months)	9.104	19.899	8.884	24.408

*p<0.10 **p<0.05 ***p<0.01, Heteroskedasticity-robust clustered p-values in brackets

Notes: Table presents results from RD regressions to detect discontinuities in outcomes at the point of cutoff. Regressions conducted non-parametrically using local linear regression methods, and optimal bandwidth selected using the plug-in procedure suggested by Imbens and Kalyanaraman (2009). Births from January 2006 and December 2005 are excluded from this sample. Regressions include controls for Personal covariates e.g. age and education of mother and spouse as well as household size, religion and number of children in various age groups. Errors are allowed to cluster around the month of birth. Robust p-values presented in square brackets.

Table A.3: Summary Test for Sample Composition Bias

	(1) Predicted Fathers' Participation Rates	(2) Predicted Fathers' Leave Duration (Weeks)	(3) Predicted Mothers' Participation Rates	(4) Predicted Mothers' Leave Duration (Weeks)
Quebec * Post-Reform	-0.004	-.051	-0.025*	-0.755***
	[0.45]	[0.79]	[0.07]	[0.00]
N	8,096	6,236	8,096	5,789

*p<0.10 **p<0.05 ***p<0.01, Heteroskedasticity-robust province-clustered p-values in brackets

Notes: Table presents difference-in-difference estimates of parents' predicted leave behavior between Quebec and Other Provinces before and after the introduction of QPIP in 2006. The dependent variable is a predicted outcome obtained by regressing the participation rate or leave duration on a host of personal and household characteristics on the pre- treatment sample i.e. mothers in Quebec before 2006. Sample spans 2002-2010 of the EICS data and comprises non-immigrant mothers aged 18-40 in cohabitating or married relationships who have experienced a birth in the last year. Set of personal and household characteristics include age of respondent and spouse, legal marital status, indicator for immigrant, household size, number of children aged 0-1 and 1-5, and various education levels of respondent and spouse. Regressions control for personal and household characteristics.

Table A.4: Difference-in-Differences in Leave Behavior, Using Sub-samples of Parents

	(1) Fathers' Participation Rates	(2) Fathers' Leave Duration (Weeks)	(3) Mothers' Participation Rates	(4) Mothers' Leave Duration (Weeks)
I. Non-Immigrant Mothers				
Quebec * Post-Reform	0.548***	3.240**	0.125**	2.203
	[0.00]	[0.01]	[0.01]	[0.14]
N	7,753	6,236	7,753	5,537
II. Young Mothers aged 18-35				
Quebec * Post-Reform	0.524***	3.040	0.125**	2.488**
	[0.00]	[0.21]	[0.01]	[0.05]
N	7,790	6,245	7,790	5,383
III. Young Mothers aged 18-30				
Quebec * Post-Reform	0.534***	3.201	0.137*	2.297*
	[0.00]	[0.21]	[0.08]	[0.07]
N	4,888	3,876	4,888	3,274
IV. Mothers who are not legally married				
Quebec * Post-Reform	0.528***	2.581**	0.091	3.139
	[0.00]	[0.04]	[0.35]	[0.10]
N	2,387	1,932	2,387	1,726
V. Mothers who have more than high school education				
Quebec * Post-Reform	0.556***	3.289*	0.126***	2.663**
	[0.00]	[0.09]	[0.00]	[0.03]
N	2,048	1,597	2,048	1,035
VI. *Mothers with spouses who have high school education or less				
Quebec * Post-Reform	0.486***	3.95**	0.080	-2.048
	[0.00]	[0.03]	[0.97]	[0.81]
N	2,179	1,799	2,179	1,565

*p<0.10 **p<0.05 ***p<0.01, Heteroskedasticity-robust province-clustered p-values in brackets

Notes: Sample spans 2002-2010 of the EICS data and comprises mothers aged 18-40 in cohabitating or married relationships who have experienced a birth in the last year. Regressions control for personal and household characteristics and province and year- fixed effects. Subsamples were selected using information from Table 4 to determine the groups that appear to have become less prominent over time in the EICS Sample in Quebec over time.

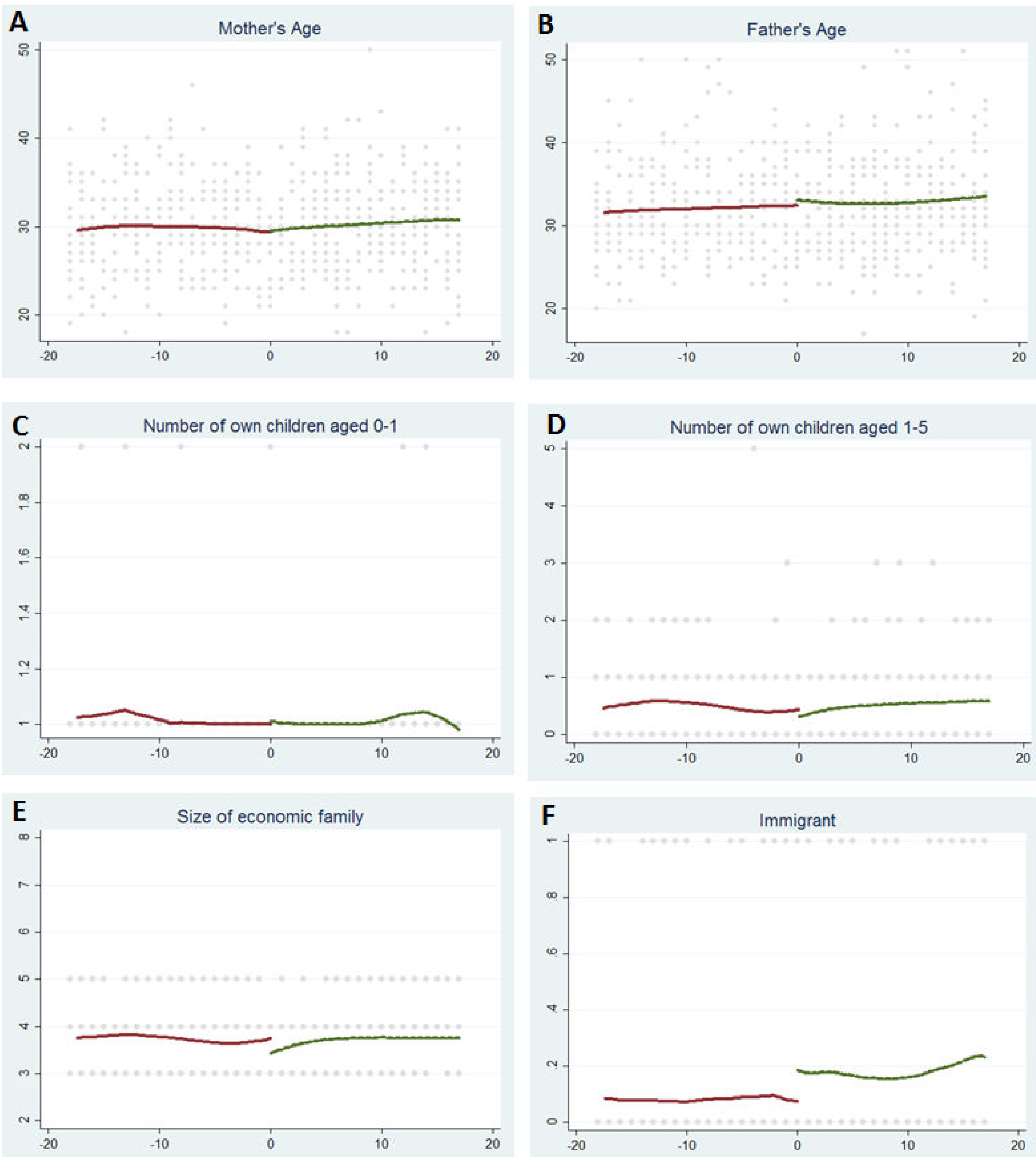
Table A.5: Difference-in-differences in Fathers' Leave Participation, 2004-2010

	(1) Fathers' Participation Rates	(2) Fathers' Participation Rates
Quebec * Post-Reform	0.521***	0.521***
	[0.00]	[0.00]
Includes Personal Covariates and Province and Year FEs	No	Yes
N	7,288	7,287

*p<0.10 **p<0.05 ***p<0.01, Heteroskedasticity-robust province-clustered p-values in brackets

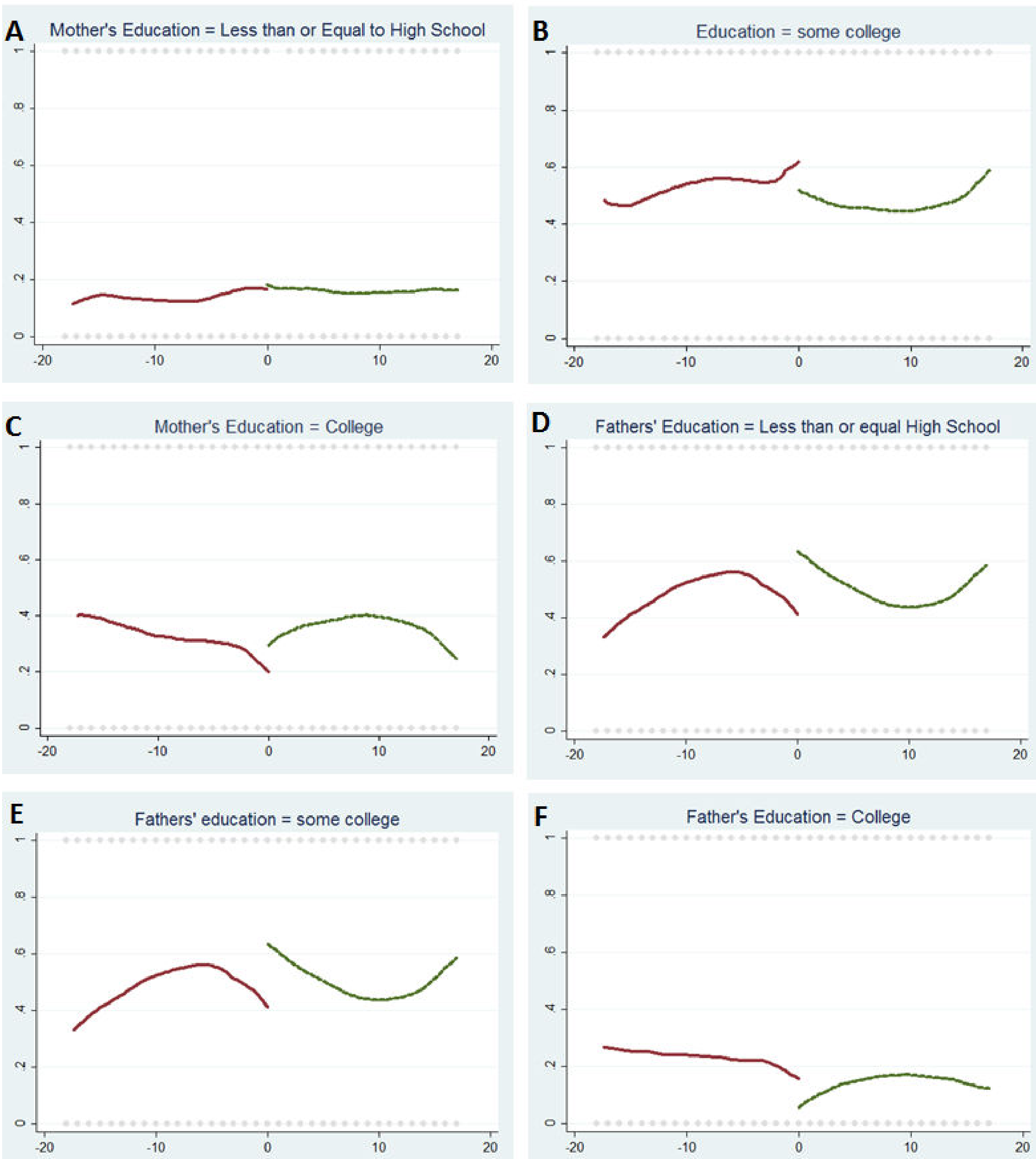
Notes: Sample is restricted to the years 2004-2010 of the EICS data to ensure comparability with data on fathers' leave duration which was only collected in those years. Sample comprises mothers aged 18-40 in cohabitating or married relationships who have experienced a birth in the last year.

Figure A.1: Discontinuities in household characteristics in EICS data



Source: Graphs created from non-parametric local linear regressions using EICS data for Quebec to identify jumps in mothers' leave duration at the cutoff of January 2006. Corresponding Regression results in Table 2.

Figure A.2: Discontinuities in educational characteristics in EICS data



Source: Graphs created from non-parametric local linear regressions using EICS data for Quebec to identify jumps in mothers' leave duration at the cutoff of January 2006. Corresponding Regression results in Table II.

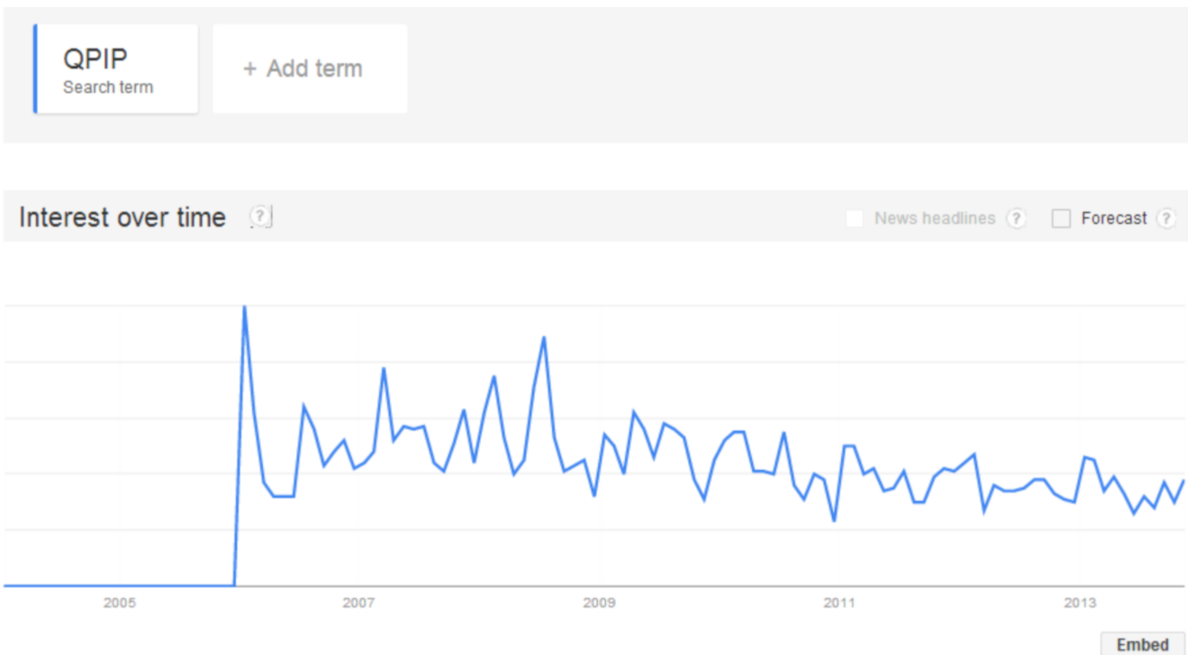


Figure A.3: Trends in Google Searches for the word 'QPIP'

Source: Graph collected by author from using Google Trends Search. Similar results were obtained when using terms such as "Quebec parental insurance program" or "Regime Quebecois de l'assurance parentale".

APPENDIX B

APPENDIX OF CHAPTER 2

Table B.1: Difference-in-Differences in Placebo Parents' Long-term Market Outcomes

OUTCOMES:	(1) Time in Paid Work	(2) Employed	(3) Time at Workplace (if employed)	(4) Full-time Worker (if employed)	(5) Usual Weekly Hours (if employed)	(6) Weeks worked Last Year ly (if employed)
Fathers:						
Quebec * Post-reform	144.370***	-0.016	133.622***	0.0131	-1.864	3.759***
	[0.00]	[0.48]	[0.00]	[0.38]	[0.20]	[0.06]
N	1115	1115	567	692	694	
Mothers:						
Quebec * Post-Reform	31.88	0.056**	-4.860	-0.116**	-2.545	-0.217
	[0.12]	[0.03]	[0.51]	[0.03]	[0.11]	[0.76]
N	824	824	590	586	583	584 1

*** p<0.01, ** p<0.05, * p<0.1, Robust province-clustered p-values in parentheses

Notes: Table presents results from double-difference regressions exploiting variation across provinces and time. Data is GSS sample of mothers and fathers aged 18-50 in married or cohabitating relationships whose youngest child is aged 5-8 years old.

Table B.2: Difference-in-Differences in Placebo Parents' Long-term Non-market outcomes

OUTCOMES:	(1) Total Time in Non-market Work	(2) Time in Domestic Work	(3) Time in Childcare	(4) Total Time Spent at Home	(5) Time with Family
Fathers:					
Quebec * Post-reform	-6.116	-31.587**	25.471***	-35.890	-63.756**
	[0.62]	[0.03]	[0.00]	[0.11]	[0.01]
N	1115	1115	1115	1115	1115
Mothers:					
Quebec * Post-Reform	-10.100	8.811	-18.911**	2.584	-5.95
	[0.31]	[0.26]	[0.04]	[0.88]	[0.65]
N	824	824	824	824	824

*** p<0.01, ** p<0.05, * p<0.1, Robust province-clustered p-values in parentheses

Notes: Table presents results from double-difference regressions exploiting variation across provinces and time. Data is GSS sample of mothers and fathers aged 18-50 in married or cohabitating relationships whose youngest child is aged 5-8 years old.

APPENDIX C

APPENDIX OF CHAPTER 3

Table C.1: QPIP's Effects on Breastfeeding Behavior Across Relationship Statuses

OUTCOMES:	(1) Breastfed / Tried Breastfeeding	(2) Breastfed at least 3 months	(3) Breastfed at least 6 months	(4) Breastfed at least 1 year	(5) Breastfed exclusively at least 12 weeks	(6) Breastfed exclusively at least 20 weeks	(7) Breastfed exclusively at least 28 weeks
Sample A: Mothers who have never been married							
Average Treatment Effect of QPIP	0.029	-0.012	0.022	0.058	-0.041	-0.007	-0.010
	(0.402)	(0.777)	(0.432)	(0.111)	(0.334)	(0.792)	(0.469)
No. of Observations	3176	2741	2741	2741	2805	2805	2805
Pre-reform Mean	0.698	0.404	0.295	0.085	0.380	0.181	0.071
Sample B: Mothers who are currently in a common law marriage							
Average Treatment Effect of QPIP	0.043***	0.045**	0.040	0.016	0.060***	0.097**	0.008
	(0.001)	(0.055)	(0.369)	(0.253)	(0.002)	(0.016)	(0.302)
No. of Observations	4932	3998	3998	3998	4273	4273	4273
Pre-reform Mean	0.805	0.484	0.348	0.062	0.452	0.188	0.047
Sample C: Mothers who are currently legally married							
Average Treatment Effect of QPIP	0.038***	0.097***	0.087***	0.036***	0.076***	0.092***	0.000
	(0.001)	(0.000)	(0.000)	(0.002)	(0.000)	(0.001)	(0.831)
No. of Observations	14925	12059	12059	12059	13159	13159	13159
Pre-reform Mean	0.837	0.598	0.439	0.121	0.555	0.291	0.091

Robust province-clustered p-values in parentheses, * p<0.10 ** p<0.05 *** p<0.01

Notes: Table presents results from regressions exploring the the program effects of QPIP by exploiting variation in time and province in difference-in-difference regressions. Regressions control for Personal Covariates such as mothers' age, household size, number of children in various age groups, age of youngest child, and Province & Year- Fixed Effects. The data is from the 2005 and 2007-2013 waves of the Canadian Community Survey, and the sample comprises mothers aged 18-40 who have a child aged 1-5 years old (or in the case of breastfeeding initiation, 0-5 years old) who was born in or after 2002.

Table C.2: Z Statistics Comparing Regression Coefficients Across Stratified Samples

OUTCOMES:	(1) Takes Leave	(2) Leave Duration (Weeks)	(3) Breastfed / Tried to Breastfeed	(4) Breastfed at least 3 months	(5) Breastfed at least 6 months	(6) Breastfed at least 1 year	(7) Breastfed exclusively at least 12 weeks	(8) Breastfed exclusively at least 20 weeks	(9) Breastfed exclusively at least 28 weeks
Single Mothers, compared to:									
Married Mothers	0.81	0.25	-0.45	-1.83**	-1.29**	-1.03	-0.25	-1.93**	-1.29
Households with income below 30,000CAD, compared to Households earning:									
30,000-50,000CAD	0.29	0.41	0.50	1.10	0.98	1.16	5.85***	1.52*	0.35
50,000CAD-80,000CAD	0.67	0.16	0.95	2.02**	-0.08	0.86	8.86***	3.64**	0.48
80,000CAD+	2.09**	1.71**	0.66	-1.71**	-1.14	-3.33***	3.47***	1.48***	1.01
Households with income 30,000-50,000CAD, compared to Households earning:									
50,000-80,000CAD	0.92	1.52*	0.49	0.33	-1.14***	2.00**	-0.38	-1.23	-0.03
80,000CAD+	5.22***	5.56***	0.07	-2.14**	-1.74**	-0.42	-1.51*	-0.23	-0.08
Households with income 50,000-80,000CAD, compared to Households earning:									
80,000CAD+	3.57***	3.73***	-0.55	-5.86***	-1.87**	-0.89	-2.65***	1.41*	-0.80
Mothers with High School Degrees or Less, compared to mothers with									
Some College	-1.41***	-4.73***	1.26	0.21	-1.48*	-1.77**	0.00	-4.71***	-1.36*
BA+ Degree	-2.92***	-4.83***	0.11	-1.68**	-1.79**	-2.56***	-2.82***	-2.41***	0.69
Mothers with some post-secondary education, compared to mothers with									
BA+ Degree	-1.94***	-1.96***	-2.23**	-2.13**	-0.97	-0.94	-4.17***	-0.38	1.21

* $p<0.10$ ** $p<0.05$ *** $p<0.01$

Notes: Table constructed by author using z-statistics for comparing regression coefficients across regressions that use different samples, using the method developed by Clogg et al. (1995).

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