RESERVING TIME FOR DADDY: 
THE CONSEQUENCES OF FATHERS’ QUOTAS

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Abstract

‘Daddy quotas’ that reserve some parental leave for fathers are increasingly common in
developed nations, but it is unclear whether fathers respond to the binding constraints or
the labeling effects they produce. Further, little is known about their long-term effects on
household behavior. I examine the Quebec Parental Insurance Program, which improved
compensation and reserved 5 weeks of leave for fathers. I find that fathers’ participation
increased by 250%, driven by a combination of higher benefits and the framing effect of
labeling some weeks as ‘daddy’-only. I also present causal evidence that paternity leave
reduces sex specialization long after the leave period.

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2010) restricted-access master files. All computations, use and interpretation of these data are entirely that
of the author.
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 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 more unpaid work in the home than do fathers (Hochschild and Machung, 1989; Blair and Lichter, 1991; Bianchi, 2011, 2012). 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.\(^2\) Consequently, men’s participation in care work has attracted growing public attention. Several nations have enacted policies to promote fathers’ participation in parental leave, e.g., Sweden, Norway, Germany and Iceland have instituted ‘daddy quotas’ that reserve some paid parental leave for fathers.

These policies are motivated by the idea that active fathering not only benefits child development (Tamis-LeMonda and Cabrera, 2002; Lamb, 2010; El Nokali et al., 2010) and improves men’s satisfaction with contact with their children (Haas and Hwang, 2008), but also may promote gender equality. Fathers’ participation in parental leave strengthens women’s labor market positions by enabling mothers to return to work earlier and by eroding the rationale behind statistical discrimination against women. Further, paternity leave makes the initial parenting experience more similar across genders, which can have long-term consequences for household behavior. In the absence of such schemes, mothers often leave the labor market for varying amounts of time and increase investments in childcare and housework, while fathers maintain or strengthen their ties to the workforce (Sanchez and Thomson, 1997), leading to differential investments in non-market and market skills. This establishes a gendered division

\(^1\)Several studies note that the provision of extended maternity leave delays women’s return to work (Schonberg and Ludsteck, 2014; Lalive and Zweimuller, 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 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.
of household labor, which persists after the leave period ends. So it can be argued that if mothers and fathers experience more similar transitions to parenthood, it might reduce sex specialization in the long run. Keeping these arguments in favor of paternity leave in mind, we are faced with two broad questions. First, what kinds of incentives are successful in getting parents to share leave more equally and what are the mechanisms behind that success? Second, if a policy succeeds in getting fathers to take leave, does this lead to more egalitarian households in the long run, or are social norms so sticky that parents revert to traditional gender roles once the leave period ends?

In this paper, I explore these questions while providing the first comprehensive causal analysis of the short- and medium-term consequences of a policy aiming to promote paternity leave. I study the Quebec Parental Insurance Program (QPIP), a system of parental leave benefits introduced in Quebec in 2006 that sought to boost fathers’ participation in 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. 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. 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 rates in Quebec never exceeded 22% prior to QPIP. Further, the majority of families did not exhaust their leave prior to the reform, such that families were leaving benefits ‘on the table’ even as fathers declined to participate.

This study makes several important contributions to the literature. First, it is the first study to explore the mechanisms behind why daddy quotas may be effective. I not only analyze the extent of QPIP’s impact on parents’ leave behavior, but also investigate whether daddy quotas can succeed in getting fathers involved by forcing their hand or by eliciting a behavioral response to the ‘daddy-only’ label. Second, this paper offers the first comprehensive examination of the longer-term causal effects of paternity leave. By simultaneously investigating multiple outcomes related to the division of household labor, it examines QPIP’s impact on overall household dynamics. Third, 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. Fourth, this is the first causal study of paternity leave to use data from time-diaries, which are considered a highly reliable method for gathering information about contributions to non-market work. This analysis explores detailed measures of parent’s daily behavior, and is able to glean a more accurate insight into the effects of paternity leave than other studies have been able to. Lastly, this is the first study of how this
Canadian policy episode affected gendered patterns in leave behavior and 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 providing a novel examination of the causal channels behind why daddy quotas may be effective and the bigger picture of sex specialization across the breadth of parental responsibilities.

To investigate the short-run effects of QPIP on parents’ leave behavior, I analyze 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 boosting fathers’ involvement: it increased fathers’ claim rates by 53 percentage points and fathers’ leave duration by 3 weeks. There is some evidence that QPIP also increased mothers’ participation, but the effect is considerably smaller in both absolute and relative magnitude. To examine the causal mechanisms behind QPIP’s impacts, I study how the program effects differed across several groups, and also the program impact on the joint distribution of parents’ leave duration. The results strongly suggest that fathers responded to not only the higher benefits under QPIP but also the ‘daddy-only’ label attached to the quota - a combination of these two factors caused a dramatic rise in men’s participation in parental leave. 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 leave still available. This paper thus provides novel evidence that daddy quotas can have a powerful labeling effect when they are accompanied with increased benefits. This is an important finding in terms of policy design, as it suggests that labeling can play an important and complementary role to compensation in influencing program participation.

Next, I investigate whether the increase in fathers’ leave-taking under QPIP had an impact on household division of labor 1-3 years after the leave period ended. I use time-diary data from the 2005 and 2010 rounds of the General Social Survey, exploiting 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 and 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. 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 shed light on how leave schemes can be designed to induce
fathers to participate, and they suggest that small changes in the initial parenting experience
can have lasting effects on parents’ behavior. In particular, they suggest that people respond
to a combination of labels and financial incentives - and that behavior learned at the begin-
ing of the parenthood experience tends to stick in later years. 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 paper proceeds as follows. Section 2 provides background on the parental leave
programs in Canada and discusses the expected effects of the reform in Quebec. Section
3 describes the existing literature. Section 4 discusses the data, methodology, results and
robustness checks for my analysis of short-run effects, and Section 5 does the same for my
analysis of longer-term effects. Section 6 concludes.

2 Background

2.1 Parental Leave Programs in Canada and the QPIP Reform

In Canada, 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. Eligible parents 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 after the birth as well as parental benefits that mothers and
fathers must decide how to share between them. All provinces continue to use the EI Program,
with the exception of Quebec. On the 1st of January 2006, Quebec launched the Quebec
Parental Insurance Plan (QPIP), to which employees now contribute and claim benefits from
instead of the EI system. Both programs are financed through payroll taxes. The details of
the programs are shown in Table I.

QPIP was designed to ease the barriers parents face to taking parental leave, namely,
inflexibility, ineligibility, financial feasibility, and gendered attitudes. First, the new system
is 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

3 The length of job-protected leave in Canadian provinces does not change over the period of my analysis.
a claimant to have worked 600 hours of insurable employment, which 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. In comparison, QPIP uses a lower earnings-based threshold, 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 ‘daddy quota’, whereby 5 weeks of leave were set aside for the father and could not be transferred to the mother. This important feature of QPIP 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 leave to be shared between parents was reduced and some weeks were reallocated to individual non-transferable leave for each parent. Mothers retained access to the same amount of potential leave (50 weeks of paid leave), while fathers gained access to more potential leave than they had earlier: 37 potential weeks under QPIP (5 of which are ‘daddy-only’) versus 35 weeks under the EI Program. Overall, QPIP increased the amount of paid leave available to a family from 50 to 55 weeks, such that total leave increased by the amount equivalent to the ‘daddy-only’ weeks.\(^4\)

### 2.2 Expected Impact of QPIP on Parents’ Leave Behavior

QPIP’s easier eligibility criteria are not expected to impact fathers significantly. Even under the stricter eligibility criteria of the EI Program, a higher proportion of men than women are eligible for benefits due to differences in employment characteristics between the two groups (ESDC, 2012). 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.\(^5\) The choice of two benefit structures under QPIP is also unlikely to be important - since the whole family had to use 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 compensation and the daddy quota.

By increasing benefits, QPIP reduced parents’ opportunity cost of taking leave by reducing the difference between benefits and foregone wages (wages lost while on leave and any change in wage trajectory upon returning from leave due to human capital depreciation). Assuming

\(^4\)Under both the EI and the 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.

\(^5\)Consistent with this, in the EICS data on benefit claims, I find no statistically significant change in the proportion of fathers reported ineligible for parental leave benefits between 2005 and 2007.
leave is a normal good, this should result in positive income and substitution effects, leading to an unambiguous increase in the amount of leave consumed. QPIP improved compensation for both mothers and fathers by increasing the replacement rate and raising the cap on earnings. If a mother and father earned the same amount, they would experience the same increase in benefits, and we might expect the mother to respond more strongly given the evidence showing that married women have more elastic labor supplies than men (Juhn and Murphy, 1997; Blau and Kahn, 2007). However, a higher-earning parent would experience a bigger increase in benefits under QPIP than a lower-earning parent, since a given percentage of their earnings would translate to a bigger absolute amount and the raised cap would mean that more of their earnings were eligible for compensation. Accordingly, whether QPIP’s improved compensation provided a family with greater incentive to have the mother or father take more leave depended on who was the higher-earning parent. Therefore, we can expect QPIP’s improved compensation to have a greater impact on fathers’ leave-taking in families where the father is the higher-earning spouse, compared to families where the father earns the same amount or less than the mother. Since fathers in Quebec earn more on average than mothers, we expect QPIP’s improved benefits to have provided the average family with increased incentive for the father to take some leave.

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, in Quebec prior to 2006, most families did not use all of their leave. Figure IA presents a histogram showing the distribution of maternal leave duration in Quebec in the period 2002 through 2005, and shows that a significant portion of mothers were not consuming the 12 months of paid leave available to the family. Figure IB shows the cumulative distribution function of mothers’ leave duration in Quebec in the period 2002 through 2005, and shows that nearly 60% of mothers in Quebec took at most 11 months of leave. Further, since the EICS survey asks mothers about all leave and not specifically paid parental leave, their answers include the 2 weeks of unpaid ‘waiting period’ under the EI program, and may include other kinds of paid leave such as vacation or sick days. In Quebec in 2002-2005, the average mother reported taking 10 months of leave - which means at most taking 9.5 months of paid parental leave. Therefore, for the average family in Quebec pre-reform, 2.5 months of leave was available to fathers but they chose not to use it. Overall, nearly 60% of families in Quebec did not consume all the leave available to them prior to the reform and therefore the newly imposed constraint of the daddy quota was not binding for them. Any increase in total family leave

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7In Canada, the employment rate of women whose youngest child is under the age of 6 is considerably lower than their male counterparts, and in 2015 full-time full-year female workers earned on average 87 cents for every dollar earned by their male counterparts (Moyser, 2017).
under QPIP should have been considered an ordinary extension of family leave since the additional five weeks were essentially fungible between parents.

Since the constraint did not bind for most families, could the daddy quota have altered the parents’ maximization problem in a different way? Let us consider the possibility of a non-monetary cost of leave. In addition to the opportunity cost of taking leave, parents may face a non-monetary ‘stigma’ cost that causes them to discount benefit income compared to wage income (Moffitt, 1983). 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 costs may be higher for men than for women. A daddy quota could have an impact even when the constraint does not bind if the ‘daddy-only’ label reduces this stigma cost for men. The 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 designated for him. Moreover, the quota sends a public message that promotes fathers’ involvement, which may reduce social stigma against men 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 may 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. More generally, several studies have found evidence of framing effects in income and cash transfers, wherein people do not treat labeled money as fungible and instead put most of the money toward the labeled purpose (Kooreman, 2000; Hastings and Shapiro, 2013; Beatty et al., 2014; Lange et al., 2014; Yinger and Nguyen-Hoang, 2015).

We thus can consider two alternate hypotheses, with several testable implications.

$H_0$: No stigma cost exists, or it exists but is not affected by the daddy-only label.

In this case, QPIP only lowers the opportunity cost of taking leave. When parents respond to the change in financial incentives, we would expect the following. (i) Leave consumption should increase for both men and women since they both experience a change in opportunity costs. (ii) We expect a bigger effect on fathers’ leave-taking when the father is the higher-earning spouse, since he experiences a bigger drop in opportunity cost. (iii) We expect a similar effect on leave-taking among fathers who experience a first birth compared to fathers who already have older children.

$H_1$: A stigma cost exists, and the daddy-only label can reduce it for men.

In this case, we would expect QPIP to cause an increase in fathers’ leave consumption that is (i) larger than the increase in leave consumption of mothers, (ii) not necessarily larger when the father out-earns the mother and (iii) larger for first-time fathers, who we expect to be more flexible in their ideas of fatherhood and caregiving.

In Section 4 I test these two hypotheses in the context of QPIP. It is important to note that since QPIP offered a bundle of reforms, we can only observe what happened when the
reform reserved some leave for fathers but also improved parents’ financial compensation. It is not possible to evaluate how QPIP’s daddy quota may have affected fathers’ leave-taking in a context where the fathers did not also experience a change in financial incentives. Therefore, it is not possible to determine whether daddy quotas can be effective by themselves - but I can evaluate whether daddy quotas can cause more men to participate in parental leave when there is also a general improvement in parents’ financial compensation.

2.3 Expected Longer-Run Effects of Fathers’ Leave Participation

Paternity leave promotes gender equality by intervening at a crucial time for renegotiating household work (Hook, 2010). It makes fathers available for time-inflexible housework and childcare, enabling mothers to return to work 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. This study, however, seeks 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 leave may lead to a persistent reduction in sex specialization, as I argue below.

The first explanation builds on the theory of Becker (1981) that a household uses productivity differentials across its members to determine an efficient allocation of resources. Since men earn higher market wages on average and women have some biological advantages in childcare, comparative advantage suggests that men allocate more time to market work while women take on more domestic responsibilities because it is efficient. However, 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. In theory, women whose husbands take paternity leave can return to work earlier and mitigate human capital depreciation - though realistically, returning 5 weeks earlier is likely too small to impact mothers’ long-term wages. As the ratio of the returns to market versus non-market

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8If 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 domestic work, while their wives who have returned to work would have higher earnings and therefore bargaining power.

9Over 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.

10Research on how paternity leave affects women’s long-term employment and wages is limited and inconsistent in findings (Johansson, 2010; Kotsadam et al., 2010; Karimi et al., 2012; Ekberg et al., 2013). Expansions in maternity leave were found to delay mother’s return to work and increase job continuity but have little effect on employment in the long run, both in Canada (Baker and Milligan, 2008a,b; Harratty and Trzcinski, 2009; Baker and Milligan, 2010, 2015) and in Europe (Liu and Skans, 2010; Dustmann and Schonberg, 2012; Schonberg and Ludsteck, 2014; Dahl et al., 2016). This suggests that paternity leave is unlikely to have large long-run effects solely through the channel of mothers returning to work a few weeks earlier.
work for mothers and fathers converge, this should lead to a within-family time allocation that is less sex-specialized.

Paternity leave may influence behavior even after the leave period ends through 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 $\alpha$ denotes the intensity of habit formation. Due to the concavity of the utility function $u(.)$, $\frac{\partial u(c_t)}{\partial c_t} \leq 0$ and $\frac{\partial u(c_{t+1})}{\partial c_t} \geq 0$. 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.

Paternity leave may create a pattern of household behavior during the period of leave that is costly to reverse later. One potential cost of changing behavior 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 experience dis-utility from returning to traditional gender roles.\(^\text{11}\)

Paternity leave should also limit the possibility of strategic shirking, since fathers cannot credibly claim to lack skills 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 have a lasting influence on the identity of fathers and mothers (Akerlof and Kranton, 2010).

3 Prior Literature

3.1 Research on Leave Policies and Fathers’ Participation

Despite the evidence that fathers’ involvement in childcare is positively associated with children’s development (Allen and Daly (2007) provide a useful summary), leave participation rates of fathers worldwide remain much lower than those of mothers. Loss of earnings is an important factor in fathers’ decisions to not take parental leave (Zhelyazkova, 2013). Fathers also cite workplace attitudes as an obstacle to utilizing leave even when they are entitled\(^\text{11}\)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.
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 prioritize paid work, or that they are rarely exposed to male role models who are caregivers.

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 (Moss and O’Brien, 2006) and is especially low in countries like the United States where leave is unpaid (Han et al., 2009). 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 - 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 leave policies and fathers’ behavior, they suffer from endogeneity issues since the assignment of a country to a policy is non-random. That is, a country may offer generous paternity leave precisely because parents are highly motivated or concerned about parental leave.

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 fathers’ parental leave use resulting from the reservation of the first ‘daddy month’. These studies present causal evidence that daddy quotas are effective. However, the specific nature of these reforms present limitations on how we can interpret the program effects. In Sweden, the reform involved a transfer of one month from family leave to ‘daddy-only’ leave, rather than the addition of one month to the family’s total (Karimi et al., 2012). Therefore, 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 daddy quota did represent additional time for the family - however, since most mothers took the entire amount of family leave previously (Dahl et al. (2014), pp.2), after the reform the family could only exhaust their total leave leave if the father utilized his month. Therefore, in both these countries, the introduction of the quota altered a binding constraint. We therefore cannot be sure about the mechanisms behind the reforms’ success. 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?
3.2 Research on the Long-Run Effects of Paternity Leave

Much of the extant research on the longer-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 fathers’ 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). Studies find an association between paternity leave and fathers’ average time 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 and longer work hours 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 household labor. 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 variation in actual leave-taking among fathers cannot control for unobserved heterogeneity in preferences, beliefs, motivation, and workplace constraints. Thus, their findings should be interpreted as informative associations rather than as causal estimates.

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 quasi-experimental studies 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 and Elseth, 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, 2011). 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., 2015). Similarly, some studies find no causal effect on mothers’ labor supply or earnings (Rege and Solli, 2013; Kotsadam et al., 2010), while others report that paternity leave leads to higher or lower maternal earnings (Johansson (2010) and Cools et al. (2015) 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 these quasi-experimental studies to reach a conclusive result may be explained by issues with the data and methodologies used. First, the data on non-market
production used in these studies is not 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. Kotsadam and Finseraas (2011) 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. Moreover, 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 and expectations 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. Ugreninov (2013)) and 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 (2011)) necessarily compare fathers of older children to those of younger children, whose behaviors may differ inherently - further, the differences may simply reflect cohort trends in parent’s behavior. Keeping these issues 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 and children’s ages in order to control for trends and provide clean identification of causal links.

4 Short-Run Effects of QPIP on Leave Utilization

4.1 Data on Benefit Claims

To analyze the immediate impact of QPIP on parents’ leave behavior, I use data on benefit claims collected annually through the Employment Insurance Coverage Survey (EICS) (Canada, 2010a). The target population for this annual survey comprises individuals who, given their recent status in the labor market, could be eligible for employment insurance. Mothers of infants less than one year old are surveyed since they could potentially be eligible for benefits via maternity or parental leave. I focus on mothers in a nine-year window framing the QPIP reform - I study data from 2002-2005 as the pre-reform period (roughly 42% of the observations) and 2006-2010 as the post-reform period.12 I use restricted-access versions of this data which can only be accessed on-site at a Statistics Canada Remote Data Center.

12There 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 survey data for births prior to 2002.
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 under one year old and identify as part of a married or cohabitating couple.\textsuperscript{13} Approximately one-fifth of the observations are from Quebec. The control group comprises observations from Ontario, Alberta, British Columbia, Manitoba and Saskatchewan and the Atlantic Region, where the EI system remained in place.

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, 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.\textsuperscript{14} 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.\textsuperscript{15} Second, the EICS survey asks new mothers about the duration of all leave (not specifically paid parental leave) taken, and could capture unpaid leave or paid sick or vacation leave that mothers may take in lieu of paid parental leave. Therefore, my measure of mothers’ leave duration represents the higher bound for the duration of paid parental leave taken by mothers. However, given the generous parental leave benefits available 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. My measure of fathers’ leave duration refers specifically to the number of weeks of paid parental leave that the mother reports that her spouse has claimed or plans to claim.

\textsuperscript{13}I exclude single parents because they do not face the decision of how to share leave or caregiving responsibilities with a partner, which is a central component of this analysis. Further, 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 targeted lower-income single parents. Due to relatively low rates of single parenthood in Canada, prior studies of Canadian leave reforms have also excluded single parents (Baker and Milligan, 2008b, 2010).

\textsuperscript{14}I confirm via regression analysis that among mothers who take maternity leave, the probability of reporting actual leave as opposed to planned leave does not change at the time of the reform.

\textsuperscript{15}There is some 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 and produce bias.
4.2 Immediate Impact of QPIP

4.2.1 Regression Discontinuity Method

To evaluate the immediate impact of QPIP at the point when 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. The final details of QPIP, 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 in December 2005 rather than January 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 time in months from the cutoff date. The model for each outcome is given by

\[ Y_i = f(m_i) + \beta(m_i \geq \text{Jan} 2006), \]

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. \( \beta \), 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) using local linear regressions (LLR) as Hahn et al. (2001) and Porter (2003) 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-in method proposed by Imbens and Kalyanaraman (2012). 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 II presents results from regression discontinuity analyses on the personal and household characteristics of the EICS Sample, showing no statistically significant

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16 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.

17 My results are robust to the use of a parametric regression discontinuity analysis instead of a non-parametric approach (See Appendix A.I).
discontinuities at the cutoff. Related RD graphs can be found in the Appendix (Figures A.I and A.II). Although the discontinuities in sample characteristics are not statistically significant, a few are large enough to warrant concern and I therefore present results from RD specifications with and without controls for these characteristics.

4.2.2 Regression Discontinuity Results

Table III presents results from regression discontinuity analyses to identify the immediate effect of QPIP on parents’ claim rates and duration. 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 an increase of more than 250% from the pre-reform participation rate of 21.3%. Column 2 indicates that QPIP resulted in a jump of 3.1 weeks in fathers’ leave duration. This is also an economically and statistically significant effect, representing an increase of 150% from the pre-reform average of 2.011 weeks. For mothers’ leave participation, the RD detects a jump of 15.5 percentage points, but the estimate is not statistically significant. There is no 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 biased by sharp changes in sample composition. Panel III presents RD analyses where the standard errors have been clustered at the level of the month 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 confirms that there was no increase in parents’ leave participation rates at the time of the reform in provinces that were not treated. Figures IIA-D and IIIA-D provide visual support for these results, graphing the local polynomial for each of the simple RD analyses (Panels I and IV of Table III). In Figures IIA-D we see large jumps in both fathers’ participation rates and leave duration at the cutoff for Quebec and in Figure IIIA, we see that in Quenec mothers’ leave participation does jump at the cutoff but falls back down again.

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 suggests QPIP induced participation from fathers even in families where the new ‘daddy weeks’ were essentially fungible.

4.2.3 Threats to Identification

Causal identification in a sharp RD 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 births is exogenous to the
introduction of QPIP. Several studies have shown that financial incentives can influence decisions surrounding fertility (Milligan, 2005; Lalive and Zweimuller, 2009; Cohen et al., 2013) and the timing of birth (Dickert-Conlin and Chandra, 1999; Tamm, 2013; Lalumia et al., 2015). For example, Gans and Leigh (2009) estimated that, when the Australian government announced a "baby bonus" of $3000 for babies born after 1 July 2004, over 1000 births were delayed to take advantage of the bonus, mostly via changes in the timing of induction and cesarean section procedures. We must therefore consider the possibility that, if people knew about QPIP sufficiently in advance, couples strategically timed 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.\textsuperscript{18} Figure III presents a ‘Google Trends’ graph tracking searches for the word ‘QPIP’ around the time of the reform.\textsuperscript{19} 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, I 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 it is infeasible to delay a birth by more than a few weeks, 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 until January in order to qualify for QPIP instead of EI. I drop all observations in the month surrounding the reform, and re-estimate the RD on this trimmed window to check how sensitive my 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.II) provide consistent estimates, with the exception that the effect on fathers’ leave duration appears smaller.

\textsuperscript{18}The idea of QPIP was discussed several years before the program came into place, but there were bottlenecks in the policy process. 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 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 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.

\textsuperscript{19}Google Trends Searches using the full English and French names of the program present similar patterns.
4.3 Average Treatment Effect of QPIP

4.3.1 Difference-in-differences Method

While the RD estimates 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 (2002 to 2010), and employ a difference-in-differences (DD) method which exploits variation over provinces and time. 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:

\[ Y_{ijt} = \alpha + \beta I[j = Quebec] \times I[t \geq 2006] + \theta I[t > 2006] + \phi Z_{ijt} + \lambda_j + \delta_s + \sigma_m + \epsilon_{ijt} \] (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. \(I[t \geq 2006]\) 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 \(\theta\) represents the change in the value of the outcome that is shared by all provinces. The term \(I[j = Quebec] \times I[t > 2006]\) 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. I do not control for all province-year interactions, but instead collapse them into the term \(I[j = Quebec] \times I[t > 2006]\). \(^{20}\) The coefficient \(\beta\) 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, \(\beta\) represents QPIP’s average treatment effect. \(\lambda_j\) denotes the set of fixed effects to capture time-invariant province-specific effects, while \(\delta_s\) denotes the set of time-specific fixed effects shared by all provinces in the year the survey was fielded.\(^{21}\) I include a set of birth-month fixed effects, denoted by \(\sigma_m\), to account for seasonality in patterns of births and leave-taking.

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, and number of children aged 0-1 and 1-5 and 6-17. Including these as regressors controls for changes in group composition. \(\epsilon_{ijt}\) 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

\(^{20}\)I do not control for province-specific time trends as these would be highly collinear with the program effect of QPIP. 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.

\(^{21}\)Survey year is not the year of last birth if the mother gave birth in the year prior but within twelve months of the survey date. Fixed effects for the survey year capture effects specific to the time that the survey is answered. A fixed effect for survey year 2002 is omitted to prevent collinearity.
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. I use the wild bootstrap-t procedures suggested by Cameron et al. (2008) to provide asymptotic refinement of standard errors. 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.

It is important to discuss the assumptions under which this difference-in-differences identification strategy is valid. 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 $\times$ 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 in January 2006 that significantly affected my sample. The second necessary assumption of the difference-in-differences identification strategy is that of ‘parallel trends’ between the treatment and control group, i.e., 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 IV.

Table IV 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 that merit mention. First, the average age of new mothers grew more in Quebec than in other provinces, a difference of 0.909 years. Second, the proportion of mothers that are married 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 de-

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22 It is possible 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.

23 In April 2004, Quebec announced a Government Action Plan to Combat Poverty and Social Exclusion which involved reforming social assistance programs to bolster low income families with children. In January 2005, three existing programs were merged into a Universal Child Assistance Benefit in the form of a refundable tax credit for families with children. In addition, a ‘Work Premium’ was introduced as a refundable tax credit that provided incentives to work. While these reforms were relevant to our sample population and not insignificant in size, there is little concern that they may be driving the results of the QPIP analyses. First, these two reforms were enacted a full year prior to the introduction of QPIP. Second, as per Milligan and Stabile (2007), the reforms’ largest impacts were felt by single parents (excluded from the sample). Third, these reforms were designed to increase incentives for parents to work, so they would, if anything, bias me against finding that parents took more leave under QPIP.
crease 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. 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, and should bias me against finding more equal sharing of parental leave responsibilities. Nevertheless, I present results from DD estimations with and without controlling for such personal characteristics, and in Section 4.3.3 I discuss whether these changes in sample composition could threaten identification.

4.3.2 Difference-in-differences Results

Table V presents results from difference-in-difference estimations of QPIP’s average impact on leave behavior. For fathers’ leave outcomes, the DD estimates of QPIP’s impact on leave behavior are very close to the estimates from the RD analysis. 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. The DD estimations of mothers’ leave duration result in a point estimate of 1.96 weeks, but this is not statistically significant. Panel II shows that controlling for personal covariates and province and year and birthmonth fixed effects does not affect the point estimates, though it affects inference in the case of fathers’ leave duration.

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. The contrast is especially stark when we consider the effects in relative terms: for mothers, a program effect of 2 weeks would represent an increase of less than 5% from their pre-reform baseline average, whereas for fathers, a program effect of 3 weeks represents an increase of 150% from their pre-reform average. The fact that all parents responded to QPIP but that fathers responded more strongly is consistent with the hypothesis that QPIP changed a stigma cost uniquely for fathers, in addition to changing financial costs for all parents.

An alternative explanation is that the response we see for fathers is driven entirely by the increased generosity of QPIP’s compensation that particularly benefited higher-earning workers. As discussed in Section 2.2, QPIP’s improved compensation would have had a larger impact on fathers’ leave taking among families where the father earned more than the mother, compared to families where the mother earned more or the same as the father. The EICS does not collect data on parents’ pre-birth earnings but does have information on parents’ education, which is a close proxy for earnings - in fact, education is a better

24 Mothers were asked about spouse’s leave participation in all survey years, but about spouse’s leave duration only in survey years 2004 and later. I examine difference-in-differences in fathers’ leave participation using the sample from 2004-2010 and find consistent results (See Appendix Table A.III).
measure of opportunity cost of leave since pre-birth earnings may be endogenous to leave decisions. Table VI Panel A presents results from DD regressions for different samples of parents depending on differences in the education level between the mother and father in a family. If QPIP’s financial incentives were all that mattered, we would expect to see the largest program impact on fathers’ leave taking among families where fathers were more educated than mothers - but we actually see the opposite. QPIP caused the largest boost in fathers’ claim rates among families where the father is less educated than the mother, and is therefore likely to earn less than her. On average, families where the father was more educated than the mother experienced a significantly smaller increase in fathers’ leave participation, compared to families where the father was similarly or less educated than the mother.\(^{25}\)

Interestingly, when I examine whether QPIP had heterogeneous effects according to the education level of each parent, I find patterns consistent with the expectation that QPIP’s improved financial incentives would cause higher-earning parents to respond more strongly than lower-earning parents (See Appendix Table A.IV). In the case of mothers, I find a significantly larger increase in the leave participation rates of those with a university degree compared to less-educated mothers. In the case of fathers, those with a university degree or with some post-secondary education experienced a significantly larger increase in their leave participation rates, compared to fathers who had at most a high school education. However, when I examine whether QPIP had heterogeneous effects according to the differences in education between spouses, as in Table VI Panel A, I find that fathers who were more educated than their wives did not respond most strongly to QPIP. One possible explanation that would reconcile these findings is that, while increased compensation was particularly effective at increasing the leave participation of higher-earning workers, the daddy quota was more effective at inducing men’s participation when parents were open to the idea of less traditional gender roles - which is more likely to be the case in families where the woman is at least as educated as the man.

Next, I examine whether QPIP had heterogeneous effects on leave-taking of parents depending on whether they had their first child or had prior children. As can be seen in Panel B of Table VI, QPIP led to a significantly larger increase in claim rates (60 percentage points) for first-time fathers, compared to fathers experiencing a birth of higher parity (47 percentage points).\(^{26}\) There were no heterogeneous program impacts by birth parity on mothers’ leave participation or duration. First-time fathers responded more strongly to QPIP even though

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\(^{25}\)The DD coefficients from the regressions of the 3 subsamples can be compared using the Z-statistic method (Clogg et al., 1995). When the outcome is fathers’ participation in parental leave, the DD coefficient for the subsample of fathers with more education than their spouses is statistically significantly smaller (at the 1% level) than the DD coefficients for the subsamples of fathers with similar or less education than their spouses. The DD coefficient for the subsample of fathers with less education than their spouses is statistically significantly larger (at the 5% level) than the DD coefficients for the subsamples of fathers with similar or more education than their spouses.

\(^{26}\)Using the Z-statistic method (Clogg et al., 1995) I find that, when the outcome is fathers’ participation in parental leave, the DD coefficient for the subsample of first-time parents is statistically significantly larger at the 1% level than the DD coefficient for the subsample of parents experiencing higher-parity births.
they faced identical financial incentives as fathers who have prior children, which strongly suggests that financial compensation alone cannot explain QPIP’s success in boosting fathers’ leave participation. A larger jump in claim rates for first-time fathers is consistent with a ‘daddy-only’ label reducing stigma - first-time fathers would be more susceptible to these effects than fathers who already have children, as the latter are more likely to be entrenched in their cultural views or personal habits around child-rearing.

Further evidence that QPIP’s daddy quota played an important role can be found in Table VII, which presents the impact of QPIP on the joint distribution of parental leave. Each cell represents a combination of mothers’ and fathers’ leave duration, and the reported coefficients are from difference-in-difference regressions where the outcome is an indicator for a family choosing that particular combination. For example, the coefficient in column 1 Row A reports QPIP’s program impact on the joint probability of a mother consuming 0 months of leave and the father consuming 0 weeks of leave. The negative coefficients across Row A confirm that QPIP reduced the likelihood of any combination where the father took 0 weeks of leave. The positive coefficients in Row B confirm that QPIP increased the likelihood that the average father took between 1 and 5 weeks of leave, i.e. consumes from his quota. An interesting pattern also becomes apparent when comparing coefficients across the columns of Table VII. The coefficients in Column 5 are consistent with families responding to the relaxation of the total family leave constraint: when the mother was consuming a full year of leave, QPIP made it less likely that the father consumed 0 weeks, and made it more likely that he consumed 1 to 5 of the newly available weeks of leave. However, the coefficients in Columns 3 and 4 find increases in fathers’ leave consumption even in families not constrained by the cap. That is, even in families where the father would always have had some leave available to him, QPIP made 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 the daddy quota), but not into Row C (consuming more than the 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. QPIP, which adds 5 ‘daddy weeks’ to the family’s budget, makes it more likely that he participates - but only consumes 1-5 weeks, even though he now has 2-3 months available. The overall pattern is that QPIP made fathers more willing to participate in parental leave, but only to consume from their ‘daddy quota’. This is highly suggestive that labeling 5 weeks of leave to be ‘daddy-only’ makes these weeks “stick” to fathers - a flypaper effect.

In summary, the difference-in-difference analyses confirm that QPIP had a large impact on fathers’ leave participation, and strongly suggest that it was driven by not only increased
financial incentives but also the daddy quota. Given QPIP’s bundling of reforms, it is not possible to infer how fathers would have responded to the daddy quota in the absence of increased financial support. However, these analyses present strong evidence that in the context where financial support improved for all parents, QPIP’s daddy quota induced more men to participate in parental leave - and it did so by leveraging a labeling effect rather than a binding constraint.

Notably, QPIP 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 5 weeks of leave, in stark comparison to mothers’ average 46 weeks of leave. QPIP was able to shrink but not close the gap in parental leave participation between men and women.

### 4.3.3 Threats to Identification

In difference-in-differences frameworks, it is common to challenge the exogeneity of reforms by questioning whether policies are in fact implemented as a response to trends in the outcome in the first place. For example, did policymakers in Quebec instituted a daddy quota because they were concerned by falling participation rates among fathers? Figure IV 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 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 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, 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 V 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 in my DD setup 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 composition 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.V). 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.VI). The results for each subsample are consistent with the results for the main sample, offering little reason to believe that QPIP’s average program effects are driven by heterogeneous effects on the sub-groups that may have increased or decreased in prominence in the sample over time.

5 Longer-Run Effects of QPIP on Sex Specialization

5.1 Data on Parents’ Time Use

To analyze the longer-run effects of QPIP on the division of household labor, I use time-diary data from Canada’s General Social Survey (GSS) (Canada, 2010b). In this time-diary survey, each 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. The 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. Notably, the GSS data are not collected at the couple-level. I therefore do not track changes in spouses’ behavior within the same household, but instead examine how mothers’ and fathers’ behavior changed
on average across households.

I 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 pre-reform while observations from 2010 are considered post-reform. The target population of the GSS is persons 15 years of age and older excluding full-time residents of institutions. Approximately one-fifth of the observations are from Quebec, while the rest come from the control group. 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 examine 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. The measure of full-time employment is constructed 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 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 examine “time spent at home,” which is the sum of all time that the respondent identifies his location as his residence, and “time spent with family,” which is the sum of all time where the respondent notes a family member was also present. Note that time spent sleeping could be recorded as “at home” or “in the presence of family” without necessarily indicating any contribution to home production or ‘quality’ time with family.

27The control group comprises observations of residents in Newfoundland and Labrador, Prince Edward Island, Nova Scotia, New Brunswick, Ontario, Manitoba, Saskatchewan, Alberta, British Columbia, Yukon, Nunavut and the Northwest Territories.
5.2 Double- and Triple-Differencing Methods

To identify the longer-term causal effects of QPIP, I apply the same difference-in-differences method described in equation (2) of section 4.3.1 to the GSS sample of parents whose youngest child is aged 1-3. This identification strategy exploits variation in exposure to QPIP versus provinces and time, and I focus on the coefficient on “Quebec*Post” as the DD effect of exposure to the QPIP program.

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 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. These parents form a convenient placebo group because even if they are observed in the treated province in the post-treatment year (2010), their youngest 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 DDD 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.

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

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

(3)

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 \( \gamma \) 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 \( \delta \), the coefficient on \( I[j = Quebec] * I[t >= 2006] * I[a <= 3] \), which captures the effect of being in the treated

28 I vary the age restrictions on the placebo group, using the same minimum age of 5 but varying the maximum from 8 to 14 and I get similar results.
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.\(^{29}\) I control for Province fixed effects through the term \(\lambda_j\). \(\epsilon_{ijta}\) 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 a placebo group of parents who were unlikely to be affected by the reform is similar to the approach used by Baker and Milligan (2008a) and Baker and Milligan (2008b). Including parents who have older children offers a good robustness check as it differences out changes that have occurred in Quebec over time that are 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. 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. 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 are of a similar or larger magnitude. 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, I focus only on the changes in parents’ behavior which are supported by both the double-difference and triple-difference results.

Table VIII 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 significant differences over time. Nevertheless, I include controls for these parental and household characteristics in each regression.

### 5.3 Difference-in-differences Results

Before beginning my discussion of the longer-term effects of QPIP, it is worth establishing a baseline for how households behaved prior to the existence of QPIP. Table IX 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

\(^{29}\)I do not control for educational characteristics for two reasons. First, the education questions on the GSS Time Use survey have especially low 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.
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 is evident, providing us with a baseline against which to evaluate the magnitude of any program effects.

Table X presents results exploring parents’ involvement in market production. Panel I shows results from both double-difference regressions using the sample of parents whose youngest child is aged 1-3, and to provide a robustness 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 conservatively focus my discussion on only those estimates from the double-difference regressions that are supported by the results of the triple-difference regressions. The DD cannot detect statistically significant changes in fathers’ time in market work, though Column 1 does show 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 spend a full hour more in paid work per day. The DD analysis finds that, conditional on being employed, exposed mothers spend 80 minutes more physically at the workplace (increase of 44% from baseline), and are 5.4 percentage points more likely to be full-time employed (increase of 7% 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 X presents evidence of increased female investments in market work.

In Table XI, 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 (increase of 23% from baseline). In housework, the DD finds exposed fathers spend 15 minutes longer per day than their counterparts (increase of 21% from baseline). In addition, exposed fathers spend 36 minutes more physically at home. Although 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 QPIP. Note that it is possible that this pattern is an indirect effect of QPIP, i.e., that the public message of the daddy quota and changing social norms motivated all fathers in Quebec to spend more time in childcare. However, the results show that fathers actually eligible for QPIP (i.e. Quebec residents who had a child after the program was introduced) were not particularly induced to spend more time in childcare.

Interestingly, Table XI shows that exposure to QPIP is 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 Exposed mothers reduced 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 (increase of 10% 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 XI does suggest that female specialization in home production is reduced overall. Nevertheless, it is interesting that I detect an increase 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 prefer childcare but exposed mothers have gained bargaining power that they use to negotiate away from less-preferred housework and towards more-preferred childcare. Lastly, although parents’ relative productivities in market versus non-market work may have changed, mothers may still have a comparative advantage in childcare versus housework.

Since I exploit variation in exposure to QPIP rather than actual participation, I provide evidence of ‘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 eligible for 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.

5.4 Threats to Identification

One threat to my identification strategy is that of migration. Since families can move between provinces, it is possible that families observed in Quebec experienced their most recent birth in another province or vice versa. However, 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. Consistent with this, Table IX reports no statistically significant difference in the proportion of parents born in Quebec among the group of exposed parents, compared to the other groups of parents. 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. I perform a robustness check by repeating the analyses using a subsample of parents who live in the province they were born in and find similar results.

Another concern is that of changes in sample composition, in particular due to changes in fertility that occurred in response to the reform. Since QPIP provided greater financial incentives to have children, are the longer-term effects I detect driven by the marginal couples that are induced to have a child by the new program, who may be different in some way? I address this concern in multiple ways. First, in terms of observable characteristics, there were no significant changes in GSS sample characteristics in Quebec compared to other provinces (see Table IX). Moreover, I control for personal and household characteristics in all my regressions, so the program effects I detect should not be biased by exposed households having different observable characteristics. Second, the marginal couples induced to have children may be different in some unobservable characteristic, e.g. beliefs about gender roles, which I cannot control for. However, the sample characteristics for new mothers in the EICS data (See Table IV) show that women who give birth in post-QPIP Quebec are more likely to be older, legally married, and less educated - all characteristics that are associated with less gender-egalitarianism. Third, even if we assume that QPIP’s effects were heterogeneous based on beliefs, it is unclear what direction this bias 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 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 reforms targeted more traditional families, which would bias me against finding reduced sex specialization. For example, QPIP increased 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 bias due to unobservable beliefs, if it exists, would take. An argument can be made that it may lead me to under-estimate the true program effect of QPIP on sex specialization.

6 Conclusion

This paper provides the first comprehensive study of the effects of the QPIP reform and offers an important contribution to the literature on paternity leave. It is the first to find that
daddy quotas can influence behavior through their labeling effects rather than the binding restrictions on who can use the leave - providing evidence that, when combined with generous compensation, ‘daddy-only’ labels can produce a flypaper effect that makes leave stick to fathers. This study also offers the first comprehensive causal analysis of the effect of 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 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 ‘daddy quotas’ may help fathers overcome such barriers to taking leave as social stigma and perceived professional penalties. Policymakers hoping to improve fathers’ leave participation should therefore consider quotas as an effective and complementary policy option to improving financial compensation for leave. Second, these results suggest that it is possible for policies that induce changes in short-term behavior to have persistent effects on people’s behavior, 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. Third, and perhaps most importantly, these 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.
References


Table I: Parental Leave Programs in Canada in 2006

<table>
<thead>
<tr>
<th></th>
<th>Employment Insurance</th>
<th>QPIP Basic Plan</th>
<th>QPIP Special Plan</th>
</tr>
</thead>
<tbody>
<tr>
<td>Eligibility Criteria</td>
<td>600 hours of insurable employment</td>
<td>2000CAD of insured income</td>
<td>2000CAD of insured income</td>
</tr>
<tr>
<td>Basic Replacement</td>
<td>55%</td>
<td>70%</td>
<td>75%</td>
</tr>
<tr>
<td>Rate for Earnings</td>
<td></td>
<td>75%</td>
<td></td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Insurable income</td>
<td>Up to 39,000CAD</td>
<td>Up to 57,000CAD</td>
<td>Up to 57,000CAD</td>
</tr>
<tr>
<td>Waiting Period</td>
<td>2 weeks</td>
<td>None</td>
<td>None</td>
</tr>
<tr>
<td>Duration</td>
<td>Total 50 weeks =</td>
<td>Total 55 weeks =</td>
<td>Total 40 weeks =</td>
</tr>
<tr>
<td></td>
<td>15 weeks maternity leave + 35 weeks parental leave + no paternity leave</td>
<td>18 weeks maternity leave + 32 weeks parental leave + 5 weeks paternity leave</td>
<td>15 weeks maternity leave + 25 weeks parental leave + 3 weeks paternity leave</td>
</tr>
</tbody>
</table>

Table II: Regression Discontinuities in Personal and Household Characteristics in EICS data

<table>
<thead>
<tr>
<th></th>
<th>Mother’s Age</th>
<th>Fathers’ Age</th>
<th>Children aged 0-1</th>
<th>Children aged 1-5</th>
<th>Family Size</th>
<th>Foreign born</th>
</tr>
</thead>
<tbody>
<tr>
<td>RD Estimate</td>
<td>0.138</td>
<td>0.624</td>
<td>0.012</td>
<td>-0.130</td>
<td>-0.324</td>
<td>0.110</td>
</tr>
<tr>
<td>[0.894]</td>
<td>[0.610]</td>
<td>[0.340]</td>
<td>[0.432]</td>
<td>[0.118]</td>
<td>[0.410]</td>
<td></td>
</tr>
<tr>
<td>Bandwidth (months)</td>
<td>16.69</td>
<td>17.14</td>
<td>4.06</td>
<td>9.65</td>
<td>9.80</td>
<td>8.15</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th>Mother’s Educ ≤</th>
<th>Mother’s Educ =</th>
<th>Father’s Educ ≤</th>
<th>Father’s Educ =</th>
<th>Father’s Educ =</th>
</tr>
</thead>
<tbody>
<tr>
<td>RD Estimate</td>
<td>0.014</td>
<td>-0.100</td>
<td>-0.101</td>
<td>0.222</td>
<td>-0.123</td>
</tr>
<tr>
<td>[0.913]</td>
<td>[0.540]</td>
<td>[0.484]</td>
<td>[0.264]</td>
<td>[0.191]</td>
<td>[0.464]</td>
</tr>
<tr>
<td>Bandwidth (months)</td>
<td>8.46</td>
<td>8.94</td>
<td>8.80</td>
<td>8.28</td>
<td>8.94</td>
</tr>
</tbody>
</table>

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 (2012). 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 III: Local Mean Impact of QPIP on Parents’ Leave Behavior

<table>
<thead>
<tr>
<th>OUTCOMES:</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Fathers’ Participation Rates</td>
<td>(2002-2005 Average)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Fathers’ Leave Duration (Weeks)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Mothers’ Participation Rates</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Mothers’ Leave Duration (Weeks)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Baseline for Quebec</td>
<td>0.213</td>
<td>2.011</td>
<td>0.725</td>
<td>42.494</td>
</tr>
</tbody>
</table>

I. RD analysis of Quebec

 Jump at Cutoff | 0.536*** | 3.088* | 0.155 | 0.427 |
 | [0.00] | [0.06] | [0.18] | [0.79] |
 Bandwidth (months) | 8.692 | 18.968 | 8.568 | 20.764 |

II. RD analysis of Quebec, including personal covariates

 Jump at Cutoff | 0.531*** | 3.126** | 0.188 | 1.050 |
 | [0.00] | [0.04] | [0.15] | [0.53] |
 Bandwidth (months) | 8.791 | 18.968 | 8.568 | 20.764 |

III. RD analysis of Quebec, including personal covariates and clustered errors

 Jump at Cutoff | 0.530*** | 3.125** | 0.187** | 1.050 |
 | [0.00] | [0.01] | [0.01] | [0.46] |
 Bandwidth (months) | 8.692 | 18.968 | 8.568 | 20.764 |

IV. RD analysis of control provinces

 Jump at cutoff | -0.026 | -1.430* | 0.036 | -0.991 |
 | [0.27] | [0.01] | [0.43] | [0.43] |
 Bandwidth (months) | 6.143 | 14.733 | 7.114 | 18.246 |

*** 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 (2012). When errors are clustered, they are done so at the level of the month of birth (assignment variable).
Table IV: Means and Difference-in-differences in Characteristics of Full EICS Sample

<table>
<thead>
<tr>
<th></th>
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<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Age of Mother</td>
<td>30.307</td>
<td>30.456</td>
<td>29.071</td>
<td>30.134</td>
<td>0.909***</td>
</tr>
<tr>
<td>Age of Spouse/Partner</td>
<td>32.732</td>
<td>32.851</td>
<td>31.900</td>
<td>32.586</td>
<td>0.552</td>
</tr>
<tr>
<td>Legally Married</td>
<td>0.862</td>
<td>0.846</td>
<td>0.366</td>
<td>0.389</td>
<td>0.040**</td>
</tr>
<tr>
<td>Immigrant</td>
<td>0.213</td>
<td>0.204</td>
<td>0.113</td>
<td>0.169</td>
<td>0.063</td>
</tr>
<tr>
<td>Family Size</td>
<td>3.784</td>
<td>3.830</td>
<td>3.711</td>
<td>3.760</td>
<td>0.006</td>
</tr>
<tr>
<td>Number of children aged 0-1</td>
<td>1.013</td>
<td>1.019</td>
<td>1.011</td>
<td>1.015</td>
<td>0.001</td>
</tr>
<tr>
<td>Number of children aged 1-5</td>
<td>0.526</td>
<td>0.582</td>
<td>0.497</td>
<td>0.528</td>
<td>-0.021</td>
</tr>
<tr>
<td>Number of children aged 6-17</td>
<td>0.243</td>
<td>0.251</td>
<td>0.205</td>
<td>0.249</td>
<td>0.036</td>
</tr>
<tr>
<td>Mother has high school degree or less</td>
<td>0.244</td>
<td>0.162</td>
<td>0.163</td>
<td>0.162</td>
<td>0.034***</td>
</tr>
<tr>
<td>Mother has some college</td>
<td>0.415</td>
<td>0.417</td>
<td>0.552</td>
<td>0.466</td>
<td>-0.087*</td>
</tr>
<tr>
<td>Mother has college degree</td>
<td>0.339</td>
<td>0.371</td>
<td>0.284</td>
<td>0.371</td>
<td>0.053</td>
</tr>
<tr>
<td>Father has high school degree or less</td>
<td>0.279</td>
<td>0.261</td>
<td>0.252</td>
<td>0.176</td>
<td>-0.060*</td>
</tr>
<tr>
<td>Father has some college</td>
<td>0.434</td>
<td>0.432</td>
<td>0.512</td>
<td>0.515</td>
<td>-0.005</td>
</tr>
<tr>
<td>Father has college degree</td>
<td>0.285</td>
<td>0.306</td>
<td>0.235</td>
<td>0.308</td>
<td>0.054</td>
</tr>
<tr>
<td>First Child</td>
<td>0.461</td>
<td>0.441</td>
<td>0.474</td>
<td>0.447</td>
<td>-0.007</td>
</tr>
</tbody>
</table>

*** 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 V: Difference-in-Differences in Parents’ Leave Outcomes

<table>
<thead>
<tr>
<th>OUTCOMES:</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Fathers’ Participation</td>
<td>0.213</td>
<td>2.011</td>
<td>0.725</td>
<td>42.494</td>
</tr>
<tr>
<td>Leave Duration Rates (Weeks)</td>
<td>0.725</td>
<td>42.494</td>
<td>0.725</td>
<td>42.494</td>
</tr>
<tr>
<td>Baseline for Quebec</td>
<td>0.213</td>
<td>2.011</td>
<td>0.725</td>
<td>42.494</td>
</tr>
<tr>
<td>(2002-2005 Average)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

I. D-in-D across Provinces & Time

<table>
<thead>
<tr>
<th>Quebec * Post-Reform</th>
<th>0.531***</th>
<th>3.225*</th>
<th>0.121*</th>
<th>1.960</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>[0.00]</td>
<td>[0.08]</td>
<td>[0.07]</td>
<td>[0.19]</td>
</tr>
<tr>
<td>N</td>
<td>8907</td>
<td>7157</td>
<td>8907</td>
<td>8907</td>
</tr>
</tbody>
</table>

II. D-in-D across Provinces & Time with Personal Controls and Province- & Year- and Birthmonth- Fixed Effects

<table>
<thead>
<tr>
<th>Quebec * Post-Reform</th>
<th>0.527***</th>
<th>3.241</th>
<th>0.125**</th>
<th>1.946</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>[0.00]</td>
<td>[0.15]</td>
<td>[0.05]</td>
<td>[0.14]</td>
</tr>
<tr>
<td>N</td>
<td>8905</td>
<td>7156</td>
<td>8905</td>
<td>8905</td>
</tr>
</tbody>
</table>

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

Notes: Table presents difference in differences in parents’ leave behavior between Quebec and other provinces before and after the introduction of QPIP in 2006. Difference-in-differences are taken across province and year. 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.
### Table VI: Heterogenous Difference-in-Differences by Parents’ Education and Birth Parity

<table>
<thead>
<tr>
<th>OUTCOMES:</th>
<th>Fathers’ Participation Rates</th>
<th>Fathers’ Leave Duration (Weeks)</th>
<th>Mothers’ Participation Rates</th>
<th>Mothers’ Leave Duration (Weeks)</th>
</tr>
</thead>
<tbody>
<tr>
<td>I. Couples where father is more educated than mother</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Quebec * Post-Reform</td>
<td>0.463</td>
<td>2.821</td>
<td>0.112</td>
<td>2.642</td>
</tr>
<tr>
<td></td>
<td>[0.11]</td>
<td>[0.96]</td>
<td>[0.99]</td>
<td>[0.55]</td>
</tr>
<tr>
<td>N</td>
<td>2359</td>
<td>1865</td>
<td>2359</td>
<td>2359</td>
</tr>
<tr>
<td>II. Couples where mother and father have same level of education</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Quebec * Post-Reform</td>
<td>0.545***</td>
<td>3.264***</td>
<td>0.164***</td>
<td>1.264</td>
</tr>
<tr>
<td></td>
<td>[0.00]</td>
<td>[0.00]</td>
<td>[0.00]</td>
<td>[0.89]</td>
</tr>
<tr>
<td>N</td>
<td>3030</td>
<td>2440</td>
<td>3030</td>
<td>3030</td>
</tr>
<tr>
<td>III. Couples where mother is more educated than father</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Quebec * Post-Reform</td>
<td>0.564***</td>
<td>3.487</td>
<td>0.105</td>
<td>1.831</td>
</tr>
<tr>
<td></td>
<td>[0.00]</td>
<td>[0.16]</td>
<td>[0.67]</td>
<td>[0.28]</td>
</tr>
<tr>
<td>N</td>
<td>3516</td>
<td>2851</td>
<td>3516</td>
<td>3516</td>
</tr>
</tbody>
</table>

**Panel B: Program Impacts by Birth Parity**

I. Couples having their first birth

| Quebec * Post-Reform | 0.600*** | 3.384 | 0.118*** | 1.062 |
| | [0.00] | [0.27] | [0.00] | [0.99] |
| N | 3805 | 3049 | 3805 | 3805 |

II. Couples having a birth of higher parity

| Quebec * Post-Reform | 0.465*** | 3.141 | 0.114 | 2.324*** |
| | [0.00] | [0.93] | [0.99] | [0.00] |
| N | 5100 | 4107 | 5100 | 5100 |

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

Notes: Table presents difference in differences in parents’ leave behavior between Quebec and other provinces before and after the introduction of QPIP in 2006. Panel A shows results for different samples based on differences in parents’ education and Panel B shows results for samples based on birth parity. Difference-in-differences are taken across province and year. Sample spans 2002-2010 of the EICS data and comprises mothers aged 18-40 in cohabitating or married relationships who gave birth in the last year. Heteroskedasticity-robust p-values, clustered at the province level and calculated using wild bootstrap procedures, are presented in brackets.
Table VII: Program Effect of QPIP on Joint Distribution of Parental Leave

<table>
<thead>
<tr>
<th></th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Mother takes 0 months of leave</td>
<td>Mother takes 1-5 months of leave</td>
<td>Mother takes 6-9 months of leave</td>
<td>Mother takes 10-11 months of leave</td>
<td>Mother takes 12+ months of leave</td>
</tr>
<tr>
<td>(A) Father takes 0 weeks of leave</td>
<td>0.001</td>
<td>-0.037***</td>
<td>-0.053***</td>
<td>-0.105***</td>
<td>-0.365**</td>
</tr>
<tr>
<td>Baseline for Quebec (2004-2005 Average)</td>
<td>[0.93]</td>
<td>[0.00]</td>
<td>[0.00]</td>
<td>[0.00]</td>
<td>[0.01]</td>
</tr>
<tr>
<td>(B) Father takes 1-5 weeks of leave</td>
<td>0.002***</td>
<td>-0.002***</td>
<td>0.070***</td>
<td>0.164**</td>
<td>0.272**</td>
</tr>
<tr>
<td>Baseline for Quebec (2004-2005 Average)</td>
<td>[0.00]</td>
<td>[0.00]</td>
<td>[0.00]</td>
<td>[0.03]</td>
<td>[0.03]</td>
</tr>
<tr>
<td>(C) Father takes 6+ weeks of leave</td>
<td>-0.001</td>
<td>-0.005**</td>
<td>0.013***</td>
<td>0.018</td>
<td>0.030**</td>
</tr>
<tr>
<td>Baseline for Quebec (2004-2005 Average)</td>
<td>[0.72]</td>
<td>[0.03]</td>
<td>[0.00]</td>
<td>[0.57]</td>
<td>[0.03]</td>
</tr>
</tbody>
</table>

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

Notes: Table presents difference in differences in parents’ leave behavior between Quebec and other provinces before and after the introduction of QPIP in 2006. Each cell represents a combination of mothers’ and fathers’ leave duration. 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. Difference-in-differences are taken across province and year. Sample spans 2004-2010 of the EICS data and comprises 7,156 mothers aged 18-40 in cohabitating or married relationships who gave birth in the last year. Heteroskedasticity-robust p-values, clustered at the province level and calculated using wild bootstrap procedures, are presented in brackets.
Table VIII: Mean Characteristics of GSS Data

<table>
<thead>
<tr>
<th></th>
<th>Parents with youngest child aged 1-3</th>
<th>Parents with youngest child aged 5-8</th>
<th>Difference-in-Difference</th>
</tr>
</thead>
<tbody>
<tr>
<td>I. Fathers’ Characteristics (N=1606)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age</td>
<td>34.896</td>
<td>34.815</td>
<td>34.407</td>
</tr>
<tr>
<td>Age of spouse</td>
<td>32.735</td>
<td>32.960</td>
<td>31.934</td>
</tr>
<tr>
<td>Canadian Born</td>
<td>0.767</td>
<td>0.736</td>
<td>0.855</td>
</tr>
<tr>
<td>Born in Quebec</td>
<td>0.035</td>
<td>0.032</td>
<td>0.964</td>
</tr>
<tr>
<td>Children under 14</td>
<td>1.963</td>
<td>1.895</td>
<td>1.842</td>
</tr>
<tr>
<td>Children in Household</td>
<td>1.939</td>
<td>1.920</td>
<td>1.857</td>
</tr>
<tr>
<td>Legally Married</td>
<td>0.889</td>
<td>0.851</td>
<td>0.499</td>
</tr>
<tr>
<td>Household size</td>
<td>4.023</td>
<td>4.034</td>
<td>3.875</td>
</tr>
<tr>
<td>II. Mothers’ Characteristics (N=1936)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age</td>
<td>32.465</td>
<td>32.600</td>
<td>31.748</td>
</tr>
<tr>
<td>Age of Spouse</td>
<td>34.709</td>
<td>35.332</td>
<td>33.996</td>
</tr>
<tr>
<td>Canadian Born</td>
<td>0.728</td>
<td>0.674</td>
<td>0.786</td>
</tr>
<tr>
<td>Quebec Born</td>
<td>0.031</td>
<td>0.028</td>
<td>0.962</td>
</tr>
<tr>
<td>Age of Youngest child</td>
<td>1.841</td>
<td>1.606</td>
<td>1.853</td>
</tr>
<tr>
<td>Children under 14</td>
<td>1.828</td>
<td>1.929</td>
<td>1.736</td>
</tr>
<tr>
<td>Children in Household</td>
<td>1.827</td>
<td>1.924</td>
<td>1.730</td>
</tr>
<tr>
<td>Legally Married</td>
<td>0.897</td>
<td>0.864</td>
<td>0.458</td>
</tr>
<tr>
<td>Household size</td>
<td>3.975</td>
<td>4.032</td>
<td>3.813</td>
</tr>
</tbody>
</table>

Notes: Data comprises a GSS sample of married/cohabitating parents aged 18-50 whose youngest child is 1-3 or 5-8 years old. Difference-in-difference-in-difference are across years, provinces and children’s age groups.
<table>
<thead>
<tr>
<th>Activity</th>
<th>Fathers</th>
<th>Mothers</th>
<th>Fathers/Mothers</th>
</tr>
</thead>
<tbody>
<tr>
<td>Daily minutes in non-market work</td>
<td>158.70</td>
<td>312.83</td>
<td>0.51</td>
</tr>
<tr>
<td>Daily minutes in childcare</td>
<td>88.69</td>
<td>165.42</td>
<td>0.54</td>
</tr>
<tr>
<td>- Interactive childcare</td>
<td>42.56</td>
<td>61.96</td>
<td>0.69</td>
</tr>
<tr>
<td>- Routine childcare</td>
<td>37.77</td>
<td>87.72</td>
<td>0.43</td>
</tr>
<tr>
<td>- Travel &amp; communication</td>
<td>8.36</td>
<td>15.73</td>
<td>0.53</td>
</tr>
<tr>
<td>Daily minutes in domestic work</td>
<td>70.01</td>
<td>147.41</td>
<td>0.47</td>
</tr>
<tr>
<td>- Cooking</td>
<td>31.78</td>
<td>70.88</td>
<td>0.45</td>
</tr>
<tr>
<td>- Housekeeping</td>
<td>16.04</td>
<td>64.38</td>
<td>0.25</td>
</tr>
<tr>
<td>- Shopping</td>
<td>48.91</td>
<td>50.86</td>
<td>0.96</td>
</tr>
<tr>
<td>- Other chores</td>
<td>12.36</td>
<td>10.45</td>
<td>1.18</td>
</tr>
<tr>
<td>- Maintenance &amp; repairs</td>
<td>9.80</td>
<td>1.69</td>
<td>5.79</td>
</tr>
<tr>
<td>Daily minutes at home</td>
<td>873.68</td>
<td>1,134.92</td>
<td>0.76</td>
</tr>
<tr>
<td>Daily minutes with family</td>
<td>409.07</td>
<td>509.32</td>
<td>0.80</td>
</tr>
<tr>
<td>Daily minutes in paid work</td>
<td>416.75</td>
<td>168.54</td>
<td>2.47</td>
</tr>
<tr>
<td>Employed</td>
<td>0.836</td>
<td>0.572</td>
<td>1.46</td>
</tr>
<tr>
<td>Daily minutes at workplace (if employed)</td>
<td>355.75</td>
<td>180.89</td>
<td>1.96</td>
</tr>
<tr>
<td>Full-time (if employed)</td>
<td>0.961</td>
<td>0.759</td>
<td>1.27</td>
</tr>
<tr>
<td>Usual weekly hours (if employed)</td>
<td>44.025</td>
<td>31.109</td>
<td>1.42</td>
</tr>
<tr>
<td>Weeks worked if (employed)</td>
<td>47.452</td>
<td>42.865</td>
<td>1.11</td>
</tr>
</tbody>
</table>

Notes: Table presents average minutes per day spent by parents in different activities, based on 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 X: Exposure to QPIP and Parents’ Market Outcomes

<table>
<thead>
<tr>
<th>OUTCOMES:</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
<th>(6)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Time in</td>
<td>Whether</td>
<td>Time at</td>
<td>Full-time</td>
<td>Usual</td>
<td>Weeks worked</td>
</tr>
<tr>
<td></td>
<td>Paid Work</td>
<td>Employed</td>
<td>Workplace</td>
<td>Worker</td>
<td>Weekly Hours</td>
<td>Last Year</td>
</tr>
<tr>
<td></td>
<td>if emp</td>
<td>if emp</td>
<td>(if emp)</td>
<td>(if emp)</td>
<td>(if emp)</td>
<td>(if emp)</td>
</tr>
</tbody>
</table>

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***]
N 988 988 901 888 885 888

Mothers:
Quebec * Post-reform: [60.147***, 0.046*, 79.908***, 0.054*, 1.353*, -3.257***]
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-reform: [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-reform: [0.23, 0.92] [0.00, 0.01] [0.03, 0.04]
N 1939 1939 1286 1278 1273 1271

Notes: Data from GSS sample of married/cohabitating parents aged 18-50 whose youngest child is aged 1-3 or 5-8 years old. Difference-in-differences are across years and provinces, and difference-in-difference-in-differences are across years, provinces and children’s age groups.
<table>
<thead>
<tr>
<th>OUTCOMES:</th>
<th>(1)</th>
<th>(2)</th>
<th>(3)</th>
<th>(4)</th>
<th>(5)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Total Time in Non-market Work</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Time in Domestic Work</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Time in Childcare</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Time at Home</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Time with Family</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

### I. Double-Differences using parents with youngest child aged 1-3

**Fathers:**
  - [0.00] [0.01] [0.00] [0.02] [0.22]
- N 988 988 988 988 988

**Mothers:**
  - [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**
  - [0.07] [0.02] [0.90] [0.11] [0.03]
- N 1596 1596 1596 1596 1596

**Mothers:**
  - [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 from GSS sample of married/cohabitating parents aged 18-50 whose youngest child is aged 1-3 or 5-8 years old. Difference-in-differences are across years and provinces, and difference-in-difference-in-differences are across years, provinces and children’s age groups.
Figure IA: Histogram of Maternity Leave Duration in Quebec, 2002-2005

Figure IB: Cumulative Distribution Function of Maternity Leave Duration in Quebec, 2002-2005

Source: Constructed by author using Employment Insurance Coverage Survey data from 2002-2005
Figures IIA-D: Discontinuities in Fathers’ Leave Outcomes

Source: Graphs created by STATA ‘rd’ program to plot local polynomials from non-parametric local linear regressions that identify jumps in outcomes at the cutoff using EICS data for Quebec and other provinces (See Panels I and IV of Table III). Statistics Canada’s disclosure requirements for restricted-access data prevent the plotting of raw data in these graphs.
Figures IIIA-D: Discontinuities in Mothers’ Leave Outcomes

Source: Graphs created by STATA ‘rd’ program to plot local polynomials from non-parametric local linear regressions that identify jumps in outcomes at the cutoff using EICS data for Quebec and other provinces. (See Panels I and IV of Table III. Statistics Canada’s disclosure requirements for restricted-access data prevent the plotting of raw data in these graphs.)
Figure III: Google Trends in Searches for the word ‘QPIP’

Source: Graph collected by author from using Google Trends Search. Similar results were obtained when using terms such as “Québec parental insurance program” or “Regime quebecois de l’assurance parentale”.
Figure IV: Parallel Trends in Fathers’ Leave Participation

Source: Graph constructed by author using Employment Insurance Coverage Survey data from 2002-2010. Graph plots average unadjusted participation rates for fathers in Quebec and other provinces.
Figure V: Event Study of Fathers’ Leave Participation Rates

Source: Graph constructed by author using Employment Insurance Coverage Survey data from 2002-2010. Event study was conducted by regressing the outcome (father’s leave participation) on indicators for birthyear as well as interaction terms between an indicator for Quebec and indicators for birthyear. Graph plots the coefficients from the interaction terms of indicator for Quebec and indicators for birthyear, 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.