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ABSTRACT

Although an increasing body of research promotes the development of universal early education and care programs, little is known about the extent to which these programs affect gender gaps in academic achievement and other developmental outcomes. Analyzing the introduction of universal highly-subsidized child care in Quebec, we first demonstrate that there are no statistically significant gender differences in the average effect of access to universal child care on child outcomes. However, we find substantial heterogeneity in policy impacts on the variance of developmental and behavioral scores across genders. Additionally, our analysis reveals significant evidence of differential parenting practices by gender in response to the introduction of the policy. The analysis is suggestive that the availability of subsidized child care changed home environments disproportionately, and may be responsible for the growing gender gaps in behavioral outcomes observed after child care is subsidized.

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An online appendix is available at http://www.nber.org/data-appendix/w23259

1 Introduction

Over the last three decades in North America, gender imbalances in many educational and labour market outcomes have rapidly evolved and boys are now falling badly behind girls at school. This result is in marked contrast to the media and policy discussions that surrounded The American Association of University Women's 1992 report titled "How Schools Shortchange Girls", which introduced gender equity as part of the debate in educational reform. Nowadays, a growing body of international evidence documents that even as early as age five, boys are lagging far behind girls in basic reading, writing and math.[1] Early childhood is also increasingly emerging as the point in the lifecycle where politicians, policy makers and researchers suggest that early education and care programs need to be introduced. Proponents argue that by not providing these programs to all children, especially those from disadvantaged backgrounds, boys and girls are being shortchanged and not being given an equal start to future learning. In this paper, we aim to provide evidence on whether the availability of, or access to, early universal publicly subsidized child care services would exacerbate or remediate gender differences in developmental outcomes early in life.

Baker (2011) and Cascio (2015) independently survey the small but growing literature on studies estimating the effects of introducing universal child care, concluding that there remains no consensus as to whether or not these programs significantly influence child development. Only a subset of this research examines gender differentiation in the impact of universal child care.[2] To the best of

[1] See Buchmann et al. (2008) for a survey of research on gender gaps in educational performance. More generally, early childhood has been identified by developmental scientists to be a period in the lifecycle in which there are gender gaps in the rate at which verbal skills develop (e.g. Feldman et al., 2000; Kramer et al., 1997; Bleses et al., 2008), spatial-mechanical play (e.g. Moore and Johnson, 2008; Levine et al., 1999) and sensitivity to environmental contexts (e.g. Crockenberg, 2003; Zaslow and Haynes, 1986); all of which are believed to translate into higher cognitive and non-cognitive skills for girls upon school entry.

[2] There is a larger literature exploring with data from targeted interventions to determine if boys and girls respond differently to early education and care. For example, Anderson (2008) conducts a re-analysis of the Perry Preschool and Abecedarian data (and a third intervention called the Early Training Project) and finds that the significant effects of the intervention were largely concentrated among girls. Heckman et al. (2010) conduct a careful reanalysis of the findings from the Perry program and are unable to find evidence of there being significant gender differences in the estimated treatment effects. Cost-benefit analysis of the Perry Preschool program conducted by Belfield et al. (2006) indicate that the net present value for participants was higher among girls than boys. Similarly, Oden et al. (2000) report that participation in targeted Head Start programs significantly increased high school graduation rates and lowered arrest rates, for girls only. While this literature has yet to reach a consensus on which gender benefits more from early childhood education, it provides substantial evidence that these programs have differential effects on boys and girls. Last, Campbell et al. (2014) and Conti, Heckman and Pinto (2015) present evidence that intensive early childhood educational programs implemented in the 1960s and 1970s differentially improved boys' health and health behaviours.

our knowledge only three studies presently exist: Datta-Gupta and Simonsen (2010) find that the introduction of universal child care in Denmark led to declines in non-cognitive skills for boys from families with lower education, while Havnes and Mogstad (2011) and Felfe et al. (2015) respectively find that girls received the majority of the long term benefits from reforms in Norway and Spain that increased the availability of regulated child care.[3]

Our study contributes to this literature by first exploring whether there are differential gender policy effects resulting from the only large scale universal subsidization of child care in North America. Specifically, we evaluate the 1997 Quebec Family Policy that provided child care for only $5-a-day for all children in the province under the age of four. This program was first formally evaluated by Baker, Gruber and Milligan (2008) (henceforth referred to as BGM) who provided evidence that the introduction of universal child care led to statistically significant reductions in a variety of child health, developmental, and behavioral measures. In addition, the authors' analysis indicated that parenting practices and family functioning in Quebec were negatively affected by the policy. This study, as well as other work evaluating the Quebec Family Policy, has not investigated whether the program had different effects on boys and girls, or their families.[4]

Second, since prior research in labour economics documents that reporting only the mean effects of a policy may mask policy relevant heterogeneity, we additionally contribute to the literature by exploring whether this policy changed any of the first four moments of the distribution of child outcomes for both girls and boys in Quebec. Motivating this analysis is not only current trends to develop policies aimed at reducing inequality in early life experiences but evidence from Autor et al. (2016) that indicates important differences in the sensitivity of the sexes to child-rearing environment. In particular, boys from disadvantaged background are found to be most sensitive. Since the Quebec policy was universal and did not solely target the most disadvantaged, we hypothesize that solely reporting estimates of the mean impact of the policy for both boys and girls may not capture the wide range of responses in child outcomes. Whether the availability of subsidized child care reduces inequality in outcomes at the start of children's lives within and between the genders appears relevant for many current child care policy debates.

[3]In related work evaluating the introduction of Kindergartens into U.S. public schools, Cascio (2009) finds that girls exhibit larger positive impacts than boys for educational outcomes later in life.

[4]That being said, several other papers have looked to see whether the impacts of child care varied across a variety of observed and unobserved dimensions. For example, Kottelenberg and Lehrer (2014) and Lefebvre et al. (2011) explored treatment effect heterogeneity across subgroups defined by child age.

Our analysis thus complements Kottelenberg and Lehrer (2016) who follow the idea of Bitler et al. (2006) of reporting distributional treatment effects. Yet, there are clear practical benefits from understanding if the heterogeneity in the impacts of child care are clearly larger among demographic subgroups. By estimating mean impacts between the genders, we can not only observe if the policy effects appear concentrated among one of the sexes, but can also speak to whether gender gaps in developmental outcomes emerge in early education and care.

Our empirical analysis reveals two key initial findings. First, we present evidence that the statistically significant reductions in four of the six child developmental and behavioral measures reported in BGM emerge only for one gender; since the policy effect is statistically insignificant for the other gender. However, formal tests reject that there are statistically significant differences in the estimated average policy effect between boys and girls for each child developmental, behavioral and health outcome. Second, we present evidence that the policy led to statistically significant gender differences on the higher order moments of several developmental and behavioral outcome variables. Specifically, our evidence indicates that the policy significantly increased both the variance and kurtosis of developmental scores for young females at a higher rate than their male counterparts. In contrast, we find the variance, skewness, and kurtosis of hyperactivity and inattention scores for males significantly increased following the introduction of the policy. These results suggest that there is significant heterogeneity in the policy impacts both between and within gender that would be masked by solely focusing on average causal effects.

To inform our understandings of the heterogeneous responses in child development to the Quebec Family Policy, the remainder of our analysis examines sex differences in child care usage and parenting behaviours. First, we find take-up differences in the mode of child care: boys are more likely to be placed in center-based care while are girls more likely to be placed in home-based care.[5] Consistent with prior research that has associated attendance at center-based care with increased cognitive performance, suggesting that gender gaps in performance would diminish as a result of this arrangement and thus, consistent with our main finding of increased variation in girl's developmental scores.

Further, we find that the introduction of universal child care led to substantial changes in the

[5] A related literature (e.g. Hiedemann et al. (2004)) demonstrates that among American families with white mothers, different child care decisions are made on the basis of the child's gender.

manner in which parents invested in sons and daughters, and in particular favours boys between the ages of 0-3.[6] This analysis provides suggestive evidence that differential changes in parenting practices by child gender may partially explain some of the gender differences in the negative effects of access to child care on developmental outcomes for girls only.[7] These finding reinforces recent evidence presented in Joo (2010) and Gelber and Isen (2013) who each document the importance of home learning environments in explaining future outcomes, and may have significant policy implications.[8] After all, recent research by Baker and Milligan (2013) has clearly shown that North American boys and girls receive different home inputs which may explain approximately half of the gender gaps in educational outcomes.[9]

This paper is organized as follows: Section 2 includes a description of data used for our analyses and further institutional details on the introduction of the Quebec Family Policy, section 3 includes a description of our empirical strategy, and section 4 includes a presentation and discussion of the results. Many of our findings are consistent with scientific findings on gender differences in the developmental process. We argue that findings from these literatures have the potential to enhance policy debates which arguably tend to concentrate on overly simplistic notions of causal mechanisms. Finally, in the concluding section we summarize our findings and discuss directions for further research.

[6] The finding of differential investment across gender in response to the policy is paralleled by more recent work from Baker and Milligan (2013). They examine and find gender specific effects of the Quebec Family Policy on parental time spent reading amongst children age 0-1.

[7] However, we do not suggest that equalizing home inputs may be desirable since research has also shown that children of different genders differ in their response to identical parental inputs. Evidence of heterogeneous response by child gender to a similar treatment / investment has been documented in numerous settings. For example, in the Moving to Opportunity housing lottery experiment, Kling et al. (2007) present evidence indicating that girls who moved to a better neighborhood saw improvement to both educational and health outcomes, whereas boys who moved to a better neighborhood had adverse outcomes in these domains. Related to this, Thomas (1994) reviews the child development literature on the impact of an absentee father and concludes that paternal absence has a greater influence on boys than girls.

[8] Specifically, evidence in Joo (2010) suggests that home environments early in life are more consistent and significant determinants of children's long-term outcomes than are early childhood care and education programs including Head Start. Gelber and Isen (2013) present convincing evidence that a significant portion of the positive effects of Head Start on child outcomes derive from changes in parental investment in their children. We concur with Gelber and Isen (2013) that investigating changes in home environments, including child rearing strategies, is highly relevant to a full welfare analysis of any child care policy.

[9] The economics literature documenting that boys and girls are raised in somewhat different family environments (and this may result in different investments) dates back to Ben-Porath and Welch (1976). See Lundberg (2005) for a recent survey of the literature on how child gender affects parental time allocation and investment decisions. More recently, Bertrand and Pan (2013) document that parental inputs in kindergarten vary significantly by child gender, particularly among single mother households, and that these investment differences partially explain the gender gap in disruptive behavior.

2 Data and Policy Setting

In 1997, the Quebec government implemented the Quebec Family Policy: a range of policies designed to strengthen governmental support of parents. A large part of this support consisted of an expansion to the child care system. Under the Quebec Family Policy, parents with children aged 0-4 were granted access to child care at a rate of \$5 per day (increasing to \$7 per day in 2004).[10] This program was implemented gradually, with access extended to children aged 4 in 1997, aged 3 in 1998, aged 2 in 1999 and aged 0-2 in 2000.

The policy also made important reforms to the structure of child care provision. The Ministère de la Famille et de l'enfance was established to develop a comprehensive early childhood program.[11] Formal qualifications for caregivers were raised and operational regulations were modified.[12] The delivery of child care services transited to larger facilities but the staff to-child ratios remained fixed at 1:8; with the exceptions of 4-5 year old children who saw an increase to 1:10.[13] Last, staff wages were scheduled to increase by 35-40 percents over a 4-year-long period. Given these large number of changes in child care delivery, we should explicitly state that in our analysis we cannot distinguish the impacts of these supply side interventions on the quality of care from the reduction in fees which occurred simultaneously.

The staggered pattern of birth cohort eligibility to subsidized child care spaces in the early years of the program was in part motivated by the need for additional child care spaces. Haeck et al. (2015) report that the number of regulated child care places in the province rose from 85,000 in 1997 to 217,000 in 2012, while provincial subsidies to child care rose from 288 million dollars in 1996/97 before the program to 2.2 billion dollars in 2011/12. Yet, the data also suggests that the

[10] The Quebec Family Policy also increased parental leave benefits and provided families with a standard child allowance based on income, family type (single parent, two parent), and number of children. Last, simultaneously full-day kindergarten was introduced for children age 5. Note, the price of childcare has recently increased but remained at \$7 per day from 2004 throughout the period in which the data we analyze was collected.

[11] Their expanded mandate was to oversee both center-based child care (group child care) for children ages 0 to 4; and family child care for children ages 0 to 12.

[12] As just one example of each, the educational requirements for the regulated daycare institutions' staff and whereas only 1/3 of the staff was previously required to be trained in early childhood education, this doubled to 2/3 of the staff by 2000.

[13] Child care under the program was provided in two venues. Child care centers (called centres de la petite enfance–CPE) were created out of existing nonprofit child care centers and were quite large. The second setting was home-based care staffed by regulated providers and organized into networks affiliated with a local CPE. Typically older children enrolled in the CPE-based care and younger children were enrolled in family home-based care. That said, to the best of our knowledge, there does not exist any data on how the number of spaces were allocated across children of different ages.

total number of spaces available in September 2000 could accommodate only about 20% of Québec children in September 2000.[14] Even with the continued increase in the number of subsidized spaces, reports of there being waiting lists for spaces at each daycare centre persist to this day. Further, since the position where individual children are placed on these lists was largely left at the discretion of the provider during the period of data we analyze, this indicates that who could attend subsidized child care was based on both parental and provider behavioral decisions.

To facilitate comparisons with the existing literature evaluating the developmental impacts of the Quebec Family Policy, we follow the sample restrictions and covariate definitions conducted in BGM. We use the first seven cycles of data from the National Longitudinal Study of Children and Youth (NLSCY),[15] a nationally representative longitudinal study tracking cohorts of Canadian children from early childhood. The first cycle of the NLSCY collected data on a random sample of Canadian children aged 0-11 in 1994-95.[16] These children were followed biannually, and a refreshment sample of approximately 2,000 children aged 0-1 is added in each new cycle of data collection.

In each cycle of the NLSCY data both child developmental scores and extensive questions relating to child care usage, parental labour supply, and other demographic characteristics are collected. Responses were collected from a child age standardized questionnaire administered in a face to face interview by a representative of Statistics Canada with the person most knowledgeable (PMK) about the child; which was the biological mother in 89.9% of cases in the NLSCY. These face-to-face interviews generally lasted between one to two hours.

We consider the exact same set of child and family outcomes that BGM used in their analysis. First, we consider a set of a binary indicators for the child being in any type of non-parental care or a specific type of care such as centre based care. Second, to measure child development we use the score on revised Peabody Picture Vocabulary Test (PPVT) score for children aged 4, and a

[14]It appears reasonable to conjecture that parents with low reservation prices for child care were now more likely under the policy to send their children to child care. Indeed, Lefebvre and Merrigan (2008) speculate that possibly liquidity-constrained low-income families, were induced to use these services once the policy was introduced.

[15]This paper makes use of data from 2004-2007, cycles six and seven from the NLSCY, which were not available to BGM. Some argued negative effects highlighted in BGM reflected short run changes in outcomes stemming from the large increase in supply to child care. This critique was addressed in Kottelenberg and Lehrer (2013) which showed that negative results persisted using this additional data. The analysis here builds on this set of findings and uses nearly identical data to Kottelenberg and Lehrer (2014) and Haeck et al. (2015) and yields efficiency gains.

[16]This sample was restricted to Canada's ten provinces and excluded both full time members of the Canadian Armed Forces and those living on Aboriginal reserves. These exclusions represent about 2% of the Canadian population.

age-standardized motor and social development (MSD) score for children aged 0-3.[17] To examine the child's health status we use an indicator variable based on the parent's subjective evaluation of whether their child is "excellent health", and reports indicating if the child never experienced either i) a nose/throat infection, or ii) an ear infection.[18]

The remaining child outcomes are collected for those who are at least 2 years of age and capture dimensions of child behavior ranging from hyperactivity to anxiety to physical aggression and opposition.[19] We should point out that solely the hyperactivity and inattention index we employ is not consistent to that used in BGM. This index, calculated as a sum of responses to questions related to frequency of various behaviors, was adjusted in cycle 4 of the NLSCY. Two questions making up part of the index were removed and one new question was added. We overcame this difference by the merging of the existing indices to produce one in which all questions are common.

In their analysis, BGM also consider indices related to family dysfunction, aversive parenting, and maternal depression.[20] These scales provide us with measures of: 1) a family dysfunction score; 2) a punitive aversive score; 3) a hostile/ineffective score; 4) an inconsistency score; 5) and a positive interactions score.[21] Last, we expand on measures related to the household environment by

[17]The total score obtained in the Motor and Social Development Section of the Child's Questionnaire is obtained from responses by the PMK to 15 questions about children in the 0 to 3 age group which is then standardized in the NLSCY by child's age in months. The underlying questions vary by child age and generally ask whether or not the child is able to perform a specific task. The scale is common in longitudinal surveys and has been used in collections of both the National Longitudinal Survey of Youth in the United States and recent versions of the National Child Development Survey in England.

[18]We also note that both the child and parental scales are shown to have reasonable levels of internal consistency (Statistics Canada, 1996), making them suitable for this analysis.

[19]Measures of these variables in the NLSCY were developed in accordance with established practices in developmental psychology and take the form of a raw score that is a simple aggregation up from responses to individual questions. For example, the Hyperactivity/Inattention Subscale ranges from 0 to 16 on the basis of answers to 8 questions (can't sit still, is easily distracted, can't concentrate or pay attention, can't settle for long, is inattentive, fidgets, or acts impulsive) that are each scored as 0 (not true), 1 (sometimes true) or 2 (often true). Details on the questions used to construct each index can be found in Statistics Canada (2003), the NLSCY documentation. We treat these indices as a continuous scale in the analysis.

[20]Similar to the child behavioural indices these are constructed from multiple questions that have responses never, sometimes, and often that are respectively assigned the values zero, one, and two. The respective indices are constructed by summing the values across all the questions. As an example, the family functioning index is constructed from the ranking on a four point scale from strongly agree to strongly disagree of the following statements: our family misunderstands each other; we can turn to each other for support; we cannot talk to each other about sadness; family members accepted as they are; we avoid discussing fears or concerns; we express feelings to each other; there are lots of bad feelings in the family; family members feel accepted for what they are; making decisions is a problem for family; we are able to make decisions/solve problems; we do not get along well together; we confide in each other; and, drinking is a source of tension in family. Details on the questions used to construct each index can be found in Statistics Canada (2003), the NLSCY documentation.

[21]Each of the four scales were derived by factor-analyzing parenting items included in the NLSCY (Special Surveys Division, 1996) and have been shown to have high levels of internal consistency (e.g. Jenkins et al.., 2003). There

including information in the NLSCY on the nature and quantity of parental time spent with their children. We create discrete indicators that measure whether the PMK reports that the amount they actively partake with their child in specific recreational and educational contexts such as reading to them or playing a sport crosses a specific threshold such as daily.

The NLSCY data quality has been shown to be of high quality since across cycles of the NLSCY, the cross-sectional response rate hovered between 85-93% in each province.[22] Each behavioral outcome utilized in this paper was shown to have a Cronbach's alpha coefficient indicative of high degrees of reliability. For example, using cycle 1 data the Cronbach's alpha were as follows: anxiety 0.59, hyperactivity 0.80, aggression 0.75 and prosocial behaviours 0.85.

Following BGM, our analysis is performed using only children aged 4 years or less living in two-parent families at the time of interview, thereby eliminating the contaminating effects of pre-policy subsidization that generally have higher utilization rates with single-headed households. This isolates an appropriate comparison group not affected by changes in other policies during this period. Further, since two-parent families remain a key focus of the universal child care debate that aims to extend subsidized access to child care to locations and individuals for which it was not previously made available, this sample is of interest to policy audiences.[23] Last, we use each child's final survey weight provided in the NLSCY that has been adjusted for nonresponse, and post-stratified by province, age and sex to match known population totals at the time of sample selection for the full sets of estimates and summary statistics.

Both Kottelenberg and Lehrer (2013) and BGM present evidence that there are few substantial differences in the unconditional rates of sample characteristics between Quebec and the rest of Canada both before and after the introduction of the Quebec Family Policy. Thus, we begin by making comparisons in household characteristics between child gender using the pooled sample. Table 1 presents summary statistics on a subset of parent and family variables that are used as control variables in our analysis, for samples defined by child gender. The third column contains tests of difference in means between the genders, finding several minor differences. Girls have higher

are five ordinal responses to the questions on these scales that range from "never" to "many times each day." As one example, the positive interaction scale includes 5 questions such as "How often do you and s/he laugh together?". In our analyses we treat these scales as a continuous variable.

[22] Statistics Canada additionally reports that in the second cycle of the NLSCY out of the 24,692 PMKs a valid answer was obtained for more than 90% of questions submitted.

[23] See Kottelenberg and Lehrer (2016) for an analysis of single-parent households that focuses on distributional causal effects.

odds of being in a family that includes: i) a younger sibling, ii) a father with lower level of education, iii) older parents, and iv) a mother who did not drop out of high school.

Summary statistics on the full set of outcome variables explored in the paper are presented in Table 2. The first panel of this table focuses on child care and maternal labour supply decisions. Over half of Canadian mothers are employed. Roughly two-thirds of working moms use child care and slightly more than 10% of non-working moms use these services. Notice there is substantial heterogeneity in the number of hours a child is in care. Based on child gender, there are no significant differences in either the type, amount, and use of child care services. The same is also true of labour supply decisions. The lack of a gender difference contrasts sharply with existing evidence using US data. For example, Hiedemann et al. (2004), show that child care decisions are related to child gender and argue that these differences in usage may reflect appropriate responses to children's developmental needs.

The second panel of Table 2 presents large gender differences in nearly every child developmental outcome. These differences are consistent with prior research examining, among other data sources, the children of the NLSY 1979 sample. In early childhood, girls perform better on motor and social development skills tests and cognitive measures including the PPVT. In our sample, girls also tend do better than boys along every health dimension. Not surprisingly, boys display higher rates of behavioral problems including hyperactivity and inattention, emotional anxiety and physical aggression relative to girls. The only measure where the gender gap in early childhood appears small is the separation anxiety index. This small gender gap may reflect the lack of significant gender differences reported for parental work and care decisions, as shown in the top panel.

The bottom panel of Table 2 exhibits large statistically significant gender differences in specific parental child rearing practices.[24] Boys are more likely to experience higher levels of ineffective parenting during early childhood. Parents of boys are also significantly more likely to do a special activity that a child enjoys, as well as plays games or sports. In contrast, girls have a slightly higher chance of being read to daily, a gap that increases in magnitude as the child ages. Girls are also more likely to be taken to a library. These descriptive statistics indicate that parents are more likely to engage in educational activities with daughters and recreational activities with sons.

[24] This is not surprising since prior research including Hiedemann et al. (2004) notes that the birth of a son is often correlated with more father and paternal grandmother involvement.

3 Empirical Strategy

Our main specification follows BGM by using a linear difference in differences (DID) model that compares changes over time in outcomes in Quebec with changes in outcomes in the rest of Canada. To compare effect estimates between genders the models are estimated separately for boys and girls. Formally, the estimating equation for an outcome of interest Y is expressed as:

$$Y_{ipt} = \beta_0 + \delta Policy_{ipt} + \beta_2 PROV_p + \beta_3 YEAR_t + \beta_4 X_{ipt} + \varepsilon_{ipt} \qquad (1)$$

where i, p, and t index individual, province, and year. The vector of covariates X, includes controls for child, parent, family, and geographic characteristics[25] while $PROV$ and $YEAR$ are respectively a series of province and time dummies. The $Policy$ variable is an interaction between the indicator for living in Quebec after 1998, the year the Quebec Family Policy was introduced.[26] The main coefficient of interest δ should be interpreted as an intent to treat (ITT) parameter. To conduct inference with 60 clusters, we first follow the guidance from the burgeoning literature that followed Bertrand et al. (2004) and report corrected standard errors by province-time cells. Second, since we are estimating the effect of the policy on multiple related outcomes, we make the Simes p-value adjustment for multiplicity.

This causal interpretation of δ relies on the maintenance of three assumptions in the underlying data. These assumptions are commonly referred to in the literature on linear difference in difference estimators as common trend, common support, and no anticipation effects. In the top 4 panels of Figure 1 we separately provide graphical evidence that trends in child care usage and maternal labour supply variables were identical between regions in Canada prior to the introduction of the policy (1995-1997), suggesting the assumption of common trend is met.[27] Prior to the introduction

[25] We use the exact same set of controls as BGM (with the sole exception of child gender since that would introduce perfect collinearity) a subset of which is presented in Table 1. To reduce issues related to misspecifying the functional form of the estimating equation, all variables included are discretized. For example, we create a host of dummy variables for the various categories measured by parental education, number of siblings and community size. Parental age is also categorized, grouping parents in five year categories as in BGM.

[26] We also considered specifications similar to Lefebvre and Merrigan (2008) that allowed the effect of the policy to vary by post-policy cycle. Since there was no systematic major differences in the estimated effects over time, we report the effect that averages across all four post-period cycles.

[27] We also formally tested this assumption by running the following regression $Y_{ipt} = \alpha_o + \alpha_1 PROV_p + \alpha_2 QUEBEC_p * YEAR_t + \alpha_3 YEAR_t v_{ipt}$, using data from the first two cycles where the outcomes investigated are maternal labour supply and child care use. The full set of results are available upon request and we did not find evidence of differential trends in Quebec prior to the implementation of the policy.

of the policy, parents in Quebec were less likely to use child care but there were no ex ante differences in maternal labor supply. The remaining assumptions of both common support and the absence of anticipatory behavior in Quebec appear plausible since ex-ante we would expect neither the observed or unobserved characteristics of individuals living in Quebec to differ substantially from those living in other provinces, nor would we expect that parents in Quebec would have altered their child care decisions prior to the implementation of the Quebec Family Policy. The concept of common support requires that individuals residing in Quebec with a particular set of observed covariates have counterparts, individuals with similar covariates, in the rest of Canada.

By maintaining these three assumptions, it is additionally possible to estimate the effects of the availability of child care policy on higher order moments of the outcome variable. Specifically, Firpo et al (2009) propose a regression based estimation method which estimates the impact of changing the distribution of explanatory variables on functionals of the marginal distribution of the outcome variable.[28] This estimator involves using OLS to reestimate Equation (1), where the dependent variable is replaced by the corresponding recentered influence function (RIF) for the distributional statistics of interest.[29] Not only is this estimator easy to implement but this strategy offers the additional benefits of both imposing identical assumptions to,[30] and controlling for the same set of

[28]This estimator is popularly referred to as unconditional quantile regression and is frequently used to estimate the impact of changing the distribution of explanatory variables on the marginal distribution of the outcome variable. We do not follow this strategy to estimate quantile treatment effects, since similar to Meyer et al. (1995) and Poterba et al. (1995), we would need to impose an additional assumption that the quantile treatment effects across time are identical across all quantiles to point identify the quantile policy effect. This additional (unattractive) assumption is not required to identify policy effects on higher order moments of the outcome variable.

[29]The influence function $IF(y; v(F_y))$ captures the impact of a single observation on the distribution statistic $v(F_y)$. For example, the impact of an observation y_i on the variance of the distribution F_y can be calculated using the variance influence function, $IF(y;\sigma^2) = (y_i - \mu_y)^2$, which equates to calculating the portion of total variance contributed by this single observation. With the Firpo et al. (2009) estimator, the independent variables are subsequently transformed by recentering the influence function at the distributional statistic of interest; ensuring that $E(RIF) = v(F_y)$.The RIF transformations used for estimating policy impacts on the variance, skewness, and kurtosis of the outcome variable are respectively defined as:

$$\begin{aligned} RIF_{variance} &= \sigma^2 + (y_i - \mu)^2 \\ RIF_{skewness} &= \frac{\mu^3}{\sigma^3} + \frac{(y_i - \mu)^3}{\sigma^3} \\ RIF_{kurtosis} &= \frac{\mu^4}{\sigma^4} + \frac{(y_i - \mu)^4}{\sigma^4} \end{aligned}$$

where μ^k is the the k^{th} moment about the mean and σ is the standard deviation.

[30]Naturally, common trend would now require testing the equality of pre-treatment trends in the respective higher order moments of the outcome variables; similar to what is reported in Online Appendix Table 1 for the mean. These results as well as tests on the equality of standard deviations that correspond to the last column of Table 2 are available from the authors upon request.

control variables; as the linear DID estimates of the intent to treat effect.

The linear DID model in Equation (1) only admit location shifts, that is the effect of the policy on mean outcomes. Yet, with many policies, substantial attention is now being paid to how they may influence inequality in outcomes, which would require an alternative model and estimator that can estimate distributional effects. As many of these alternative estimators are computationally challenging, utilizing the Firpo et al. (2009) estimator to recover policy impacts on higher order moments of the outcome variable provides new insights on how the policy operates and can help establish whether there is any evidence of policy effect heterogeneity within gender to help determine whether estimating distributional effects is likely to be fruitful.[31]

4 Results

4.1 Do the effects of access to universal child care differ by gender?

To begin our analysis we examine whether there are similar responses in child care use and maternal work from families with a child of each gender. Because our estimated parameter for the child outcome variables are intent-to-treat parameters which directly reflect the changes in levels of treatment, child care use and maternal work, it is important for the comparison across sexes that each experiences similar changes to these take-up variables. Table 3 presents by sex the percentage point in labor supply and child care attendance outcomes, estimates of δ from Equation (1), showing statistically different estimates between the genders by reporting results in bold font. This table shows that while there were significant changes in patterns of take-up following the policy, there are relatively few statistical differences by sex. We document roughly 10-20% larger increases in the magnitude of child care attendance, hours in care, and maternal work for boys, but these differences are not statistically significant. An interesting pattern in the take-up that does emerge is that the type of child care used by parents differs significantly by gender. In general, there is a shift toward institutional or center based care but this shift is disproportionately represented by changes for male

[31] In the online appendix available from the CJE online archive at economics.ca/cje/fr/archive.php, we discuss and provide evidence of similar patterns of policy effect heterogeneity from an alternative estimation strategy that is motivated by Solon et al. (2015) and involves reweighting. To the best of our knowledge, only Kottelenberg and Lehrer (2016) provide evidence of distributional policy effects of Quebec's child care policy using the change in changes estimator that makes a different set of assumptions than the linear difference in differences estimator.

children. Boys' child care attendance shifts away from care in another person's home while this is not the case for girls.[32] In summary, Table 3 establishes a similar change in treatment between boys and girls but highlights important differences in the type of treatment received.

The main question of our analysis is to determine whether the negative impacts to child developmental and behavioural outcomes reported in Kottelenberg and Lehrer (2013) and BGM are primarily driven by boys or girls. To this end, the first column of Table 4 reports the intent-to-treat estimates for each child outcome by sex and reports statistical differences between the sexes reported in bold font.[33] This analysis finds several statistically different from zero effect estimates for one sex but not the other. For boys, access to subsidized child care leads to statistically significant declines in the MSD score and increases in the hyperactivity and inattention score. Similarly, we observe that the statistically significant policy effect for both emotional anxiety and separation anxiety scores emerges only for girls. While the magnitude and statistical significance in the estimated policy effects often appears to differ between genders, formal tests indicate that there are no statistically significant gender differences in the average policy effect for any outcome.

The remaining columns of Table 4 explore the impact of the Quebec Family Policy on the other summary statistics that measure the dispersion, peakedness and skewness of the outcome's distribution. There are several interesting findings pertaining to the policy effect on the distribution of both the MSD score and PPVT score. Outcomes for girls are found to become relatively more variant relative to boys, both the effect of the policy on the variance and kurtosis of these scores are significantly different between sexes. One explanation for this result may be found in the relatively higher proportion of boys switching into center or institutional based care which is traditionally associated with increased cognitive performance.

Turning to the effect of access to subsidized child care on the higher order moments of behavioral outcomes, we observe that it may exacerbate existing externalizing behavior for boys. We find statically significant increases both in the skewness of hyperactivity and inattention score and the skewness and kurtosis of the physical aggression score. Despite the policy not leading to a statistically significant difference between the genders at the mean for these outcomes, we find

[32] Kottelenberg and Lehrer (2014) present evidence that after the policy was implemented the gains in attendance are somewhat similar for children age 1-4. Their results indicate less heterogeneity in take up of center based care on the age than the gender dimension.

[33] We do not report estimates for the health outcomes reported in Kottelenberg and Lehrer (2013) and BGM but find no significant difference between boys and girls.

evidence that there are gender differences in the higher order moments of these distributions. Since these variables generally predict extreme forms of delinquency later in life, this observed polarization points to the possibility that this policy may contribute to the widening gender gap reported in Baker et al (2015).

Together, the plethora of statistically significant policy impacts both within and between genders on the dispersion, peakedness and skewness of the childs' outcome distribution suggests that this policy operates in a heterogeneous manner. This heterogeneity would be masked by only using the traditional linear DID estimator that only allows for location shifts. However, the complete absence of gender differences in average policy effects but statistically significant gender differences in the effect of the policy on the variation and kurtosis of developmental scores appears puzzling. After all, collectively earlier research on universal child care programs points to girls faring better than their male counterparts (Datta-Gupta and Simonsen, 2010; Havnes and Mogstad, 2011; Felfe et al., 2015). While these results may not extend to Quebec's experience, we next attempt to determine whether other mechanisms proposed in the literature can reconcile our results at the mean level with the findings in prior research.

Unfortunately, without imposing strong structural assumptions it is not possible to isolate the influence of subsidy receipt from the influence of maternal employment as the mechanism linked to declines in outcomes observed in Table 4. Indeed, there are multiple causal pathways through which child care subsidies could affect family well-being. For example, Herbst and Tekin (2014) hypothesize three main pathways which could all cause changes in maternal and child well-being: maternal transitions from leisure to labor (through entering the labor market), familial changes in income and consumption, and changes in the nature and quantity of time mothers spend with children due to subsidy receipt.

The latter pathway suggests child rearing may be influenced by the policy. A central message from the education production function literature, dating back to the Coleman Report, is that home inputs such as parental time spent challenging children in math and reading has a positive and significant impact on subsequent achievement measures.[34] It has also been hypothesized that some

[34] Recent evidence on gender gaps in educational attainment including Bertrand and Pan (2013) and Autor et al. (2016) indicate that boys and girls are differently affected by the quantity and quality of inputs received in childhood. Further, these studies show that gender gaps in behavioral outcomes are larger for children from more disadvantaged families. Family disadvantage is highly correlated with lower neighborhood and school quality as well as less parental education and the study by Autor et al. (2016) exploits within family differences in circumstances between siblings

parents have a preference for boys over girls and therefore they invest more in boys. For instance, a large body of research surveyed in Bharadwaj et al. (2012) and Baker and Milligan (2013) present evidence that even in countries where sex preferences are thought to be small or non-existent there are sex differences in the amount of time parents report spending on various activities with their children. These differences may arise from parental beliefs regarding child ability based on gender.[35] As Table 2 documents, such differentiation exists within Canadian households. We next consider to what extent does the availability of subsidized child care affects specific child rearing practices.[36] Assuming that both parental investments and child care are inputs into a human capital production function for the child and that these inputs have different returns, parents face trade-offs between work, leisure, and child well-being. The Quebec Family Policy reduces the price of child care relative to other investments parents can make to the production process. In equilibrium, when making child care decisions the marginal rate of substitution between leisure and parental time investment into their children is equal to the price of leisure divided by the price of child care. Some parents will change the manner in which their children receive care when the costs of child care are lowered.

To examine how investments in home environments responded to the introduction of the Quebec Family Policy on average, we follow a strategy in Kottelenberg and Lehrer (2016) by reestimating Equation (1). We extend the earlier work by exploring gender differences in these policy impacts and the results of this exercise for each of the outcomes summarized in the last panel of Table 2 are presented in Table 5.[37] The top panel examines the impacts on measures of parental health and household environment. There are few significant changes to parental health after the policy was introduced but in general families with girls increasingly experience worse home environments. The intention-to-treat estimates indicate that following the introduction of subsidized child care, on average girls face significantly lower levels of parent consistency and lower levels of positive

to generate their findings.

[35] For example, see Furnham et al. (2002) and Frome and Eccles (1998) for evidence that parents believe that their sons' mathematical ability is higher than their daughters.

[36] Only a handful of economic studies in the child care literature jointly examine labor supply decisions and the demand for specific modes of child care. For example, Blau and Hagy (1998) and Powell (2002) show that price elasticities for center based care are quite elastic. These studies are sensitive to model specification and to the identifying assumptions maintained. To the best of our knowledge, studies have yet to consider either changes in parental child rearing and investment practices or differences by gender directly.

[37] Note, an important potential drawback of our parental investment measures investigated in this table is that they only measure investments as a flow at a certain point in time, rather than a stock that has accumulated since birth. Thus, we cannot rule out that differences in the stock of parental investments are much more important for explaining the gender gap in responses to the policy.

interactions with their parents relative to boys. In addition, only girls experience statistically significant declines in parent consistency. Conversely, the family dysfunction index is statistically significant for boys.

The final two panels of Table 4 present evidence regarding changes to specific parental inputs for children aged 0-3 and 4 respectively. The middle panel highlights that following the policy introduction parents of children aged 0-3 significantly decreased the amount of time spent doing activities with their child, focusing on their child, reading to their child, and laughing with their child. These estimated declines are approximately twice as large for girls relative to boys.[38] This removes the existing gender gap in this outcome that existed prior to reform where girls generally had 3-4% slightly higher rates of receiving these activities.

Examining the intention to treat estimates suggests that once subsidized child care was introduced girls experienced increases in focused time with their parents but experienced decreases in the time spent doing either special activities or being taken to the library with their parents. These latter activities are more time-intensive and are more likely to take place outside of the home. The evidence from Table 5 clearly demonstrates that there are significant gender differential responses to the policy in home investments influencing the home environments. These results extend the findings in Kottelenberg and Lehrer (2016) who report estimates by strata of the developmental score measures separately for children aged 0-3 and children aged 4. The results in Table 5 focus on whether there are significant differences in parenting measures across gender rather than across the unconditional developmental score distribution, which we argue is more informative since parents may respond differentially to the policy by child gender but since they are unlikely to know the exact rank of their child in the distribution making the distributional analysis more challenging to interpret.

To provide suggestive evidence that parental response are a primary mechanism through which the reform may have affected child outcomes at the mean level, we re-estimate Equation (1) for the same set of outcomes considered in Table 4 where we stepwise include outcomes in Table 5 as

[38]More recent work by Baker and Milligan (2013) also examine gender specific changes in parenting that occur in response to the Quebec Family Policy. Also using the NLSCY, they examine the effects of parenting children aged 0-1, and do not restrict their sample to two parent families. In their study, time spent reading is calculated on a 5 point scale from rarely/never to everyday. In contrast to the decline in everyday reading we find for children aged 0-3, they illustrate an increase in overall reading to children as a response to the policy amongst children aged 0-1. Similarly,they find that male children are better off than their female counterparts.

controls for home response. In these specifications, we additionally interact each of these additional regressors with an indicator for residing in Quebec to capture any geographic differences in these behaviors that existed in the absence of the policy. Intuitively, if the estimated policy effect from Equation (1) is found to be robust to the inclusion of any of these variables, the evidence is not consistent with the hypothesis that the reform effect is due to a particular change in parental behavior or household environment.[39] Prior to presenting the results, we should explicitly state this evidence is suggestive since we are treating the parenting and household variables that are being included in the model as exogenous.[40]

The results from this exercise are presented in Table 6. Each row corresponds to a different child outcome and we restrict the sample to those with full records on all the parenting measures. Thus, the first column corresponds to what was presented in Table 4 but uses a sub-sample of the data and fortunately there are no major changes in the ITT estimates. The remaining three columns differ based on which parental practices or parental activities are being included in the specification. The final column provides the results when all of these measures are jointly included. Notice, for children of both genders that after controlling for the complete set of parental variables the effects of the policy in nearly all cases attenuates towards zero. Excluding the PPVT score, the observed changes are larger for girls and generally corresponds to roughly 75% of the policy effect, whereas for boys it is closer to 20% of the policy effect. The different degrees of attenuation correspond to the observed larger negative policy responses to the parenting scales and activities by gender reported in Table 5. Alluding to the findings in Autor et al. (2016), the analysis of the PPVT score, the variable most closely linked to school-readiness, highlights a greater degree of responsiveness in the policy estimates for boys to the inclusion of parenting variables. Despite presenting differing channels in this analysis, via the inclusion of different subsets of the parenting behaviors and inputs, we can not clearly indicate the importance of one parental activity over another given the interconnected nature of parental responses. Overall, this analysis suggests that home responses are one of the main mechanisms through which the reform affected child outcomes particularly among girls, and

[39] We are grateful to an anonymous reviewer who suggested we undertake this analysis.

[40] In Section 3 of the online appendix we follow a robustness check undertaken by Baker and Milligan (2016) to evaluate how different assumptions about selection bias due to unobserved factors might explain the policy impacts, while controlling for parenting inputs. Our analysis presented and discussed in the online appendix demonstrates our main conclusions are indeed robust to reasonably sized failures of there being correlations in the unobserved factors that explain both parenting measures and child behavioral outcomes.

that there is likely a myriad of parental responses that are acting together rather than a single specific pathway.[41]

4.2 Discussion

In their analysis of Quebec's subsidized child care policy, BGM posited that the negative impacts on developmental outcomes might arise from changes in maternal well-being due to labour force entry. However, ex ante there is no reason to expect the differences between the genders in the causal estimates reported in Table 4. We find that there are numerous differences in the estimated effects of parental and family outcomes reported in the top panel of Table 5. Together, these results in combination with the evidence in Table 6 are suggestive that the policy may have had indirect effect of changing the household environment and home inputs.[42] Finally, we would like to reinforce that in Table 3, we do not see any significant difference in either attending care or the amount of time spent within in care, suggesting that differences in compliance with the policy cannot explain the differences in the estimated impacts.

The evidence of significant treatment effect heterogeneity by gender uncovered in Tables 3 to 5 parallels evidence from the developmental sciences literature and suggests an additional mechanisms may be at work. While neuroscientists do not yet fully understand the biological basis of brain development, several stylized facts are emerging from this literature. First, early childhood has been shown to be a point in time where gender differences in brain anatomy and functionally emerge.[43] Second, progressive myelination and regressive pruning are processes known to differentially affect

[41] Table 6 only reports changes to the mean estimates. We did conduct a similar analysis on the higher moments explored in Table 4. In general, we found that the addition of parenting variables to this analysis resulted higher amounts of variation being attributed to the policy variable. From this result we would infer that the move to child care is met with a wide variety of responses that are dampened by parenting behaviours.

[42] We also note that consistent evidence in Booth et al. (2002), we believe additional analyses presented in an online appendix cast doubt on the hypothesis that stress from entering the workplace is driving the negative effects of subsidized child care. These authors compared the child rearing practices of mothers with careers to mothers who stayed at home to raise their children, and found very small differences in the quality of time spent by mothers with infants, although the quantity of time spent by mothers with infants was much greater for stay at home mothers. Working mothers appear to compensate for their time away from their child by spending more social time, particularly on the weekends. In addition, Booth et al. (2002) report that fathers of children in child care also appeared to be more involved with parenting and interaction, perhaps as another means of compensating for the non-maternal care.

[43] Recent survey of the brain anatomy by Cosgrove et al. (2007) indicate that women have a larger caudate and hippocampus; regions known to be involved with spatial memory and correlate with ability to learn language. This appears consistent with the gender gap in PPVT scores documented in Table 2 as well as Kimura (2000) and Berninger et al. (2008) who respectively show that the average 20-month old girl has twice the vocabulary of the average 20-month old boy and that girls have more advanced spelling and grammar skills at early ages.

the functionality of the brain by gender. These processes are determined in part as a response to how environmental factors interact a child's own characteristics and the policy substantially altered child care arrangements as well as our evidence of gender differences in changes in the home environment.

Finally, a potential biological mechanism to explain increased poor health and stress-related behavioral problems are increases in the level of cortisol; a hormone associated with stress. Bradley and Vandell (2007) summarize research showing that i) children in child care are at elevated risk of increased cortisol secretions if they have either difficulties interacting with their peers or insensitive parents,[44] and ii) irrespective of gender, children who begin care earlier in life and were in care 30 or more hours a week experience higher odds of poor outcomes.[45] Watamura et al. (2010) present evidence that within a child care center, cortisol secretions are higher for boys than girls. While child care quality levels unsurprisingly vary significantly across settings, a striking finding from those evaluating Quebec's policy is that the overall child care quality was reported to be minimal and this may explain some of the negative consequences. Specifically, both Japel et al., (2005) and Drouin et al. (2004) conclude that as a whole, the general level of quality in Quebec's child care settings had not attained the levels needed to foster the social, emotional and cognitive development of the children. Our results showing that the policy, on average, increased time spent in child care disproportionately for boys who, and led to more varied outcomes in hyperactivity, inattention and physical aggression for boys relative to girls, appears consistent both with this research and that in the developmental sciences. We suggest that incorporating advances from the scientific literature may help inform whether there are gender-specific strategies or policies that can be undertaken in either home or child care environments to improve the developmental process.

[44] This is consistent with a mechanism postulated in BGM that the negative behavioral effects documented for children may be the result of difficulties children may have experienced when they were moved into child care centers, a more social environment where they needed to interact with more children.

[45] This body of research is also consistent with Blau (1999, 2001) highlighting the importance of the heterogeneity in the quality of child care across provider locations. See Kottelenberg and Lehrer (2014) for evidence on the significant age differences in policy effects from this reform in Quebec.

5 Conclusion

This paper extends earlier research evaluating the developmental consequences of the Quebec Family Policy by first documenting that the reductions in four of the six child developmental and behavioral measures reported in earlier research are driven by children of one gender. Once subsidized child care is made available, only boys face statistically significant reductions in motor social development and increased hyperactivity and inattention scores. However, formal tests reject that the estimated average policy effect between boys and girls is statistically significant differences for each child developmental, behavioral and health outcome. In addition, we find significant evidence of differential parenting practices by child gender in response to the policy. When we account for these changes in household environment and investment decisions assuming they are exogenous, we find substantial reductions in average policy effects. These results are suggestive that behavioral responses in the home related to child investments are likely one of the main mechanisms through which this child care reform negatively affected many child outcomes.

In this paper, we additionally propose an easy to implement strategy to analyze data from natural experiments to determine if the policy reduced inequality in child outcomes. We find that the availability of subsidized child care generally increased the variation in most child outcomes excluding separation anxiety for both boys and girls, suggesting within gender that children are getting off to different starts. Further, our analysis of the policy impacts on higher order moments uncovers several significant and meaningful differences between the genders. For developmental outcomes, we find the policy leads to greater peakedness in the distribution for girls relative to that of boys. When examining hyperactivity and inattention scores, we observe that the policy significantly increases variation in these measures for boys but led to a significant decrease for girls, which may explain why Baker et al (2015) report exacerbated behavioural consequences later in life for boys from the availability of child care.

Future research is needed to more fully explore whether these universal early education and care programs have impacts that persist differentially by gender later in life.[46] In addition, our findings of large changes in child rearing practices in response to the policy make us question

[46] As noted earlier, there is mixed evidence from European studies that differ if long term benefits are more likely to accrue to girls than boys. At present, only Baker et al. (2015) and Kottelenberg (2015) present some negative long run consequences to the introduction of universal child care in Canada, but neither study formally explored gender differences.

the extent to which differences in the use of child care reflects either appropriate responses to children's developmental needs or parental knowledge of the actual quality of child care centers. The technology underlying most every outcome in early childhood remains poorly understood in the research community, and it appears increasingly unlikely that parents could make optimal investment decisions and that these decisions would be invariant to child gender. Thus, we believe that the negative effects of providing access to subsidized child care are simply indicative of parents' limited knowledge of their child's human capital production function and related propensity to make optimization errors when choosing inputs. That being said, our analysis also implies that policymakers may need to supplement early childhood education policies with training in parenting skills to reinforce these policies. In conclusion, we believe our findings suggest that in an effort to improve human capital and developmental outcomes, educators and policymakers should not ignore the scientific evidence of gender differences in developmental processes when designing policies that change learning environments.

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Table 1: Summary Statistics by Gender

	Girls	Boys	P-Value
Observations	18832	19816	
Mother's Characteristics			
Younger than 21	0.009	0.010	(0.341)
Age 21-30	0.380	0.382	(0.602)
Age 31-40	0.555	0.557	(0.686)
Older than 40	0.056	0.050	(0.015)**
Did not complete High School	0.094	0.099	(0.072)*
Completed a University Degree	0.278	0.279	(0.861)
Immigrant	0.207	0.208	(0.757)
Father's Characteristics			
Younger than 21	0.003	0.002	(0.053)*
Age 21-30	0.234	0.244	(0.029)**
Age 31-40	0.615	0.615	(0.927)
Older than 40	0.148	0.139	(0.012)**
Did not complete High School	0.128	0.120	(0.016)**
Completed a University Degree	0.255	0.265	(0.017)**
Immigrant	0.203	0.214	(0.004)***
Family Characterisitcs			
Resides in Rural Region	0.133	0.134	(0.872)
Resides in a Large City (>500K+)	0.479	0.462	(0.001)***
Child has younger siblings	0.235	0.227	(0.058)*
Child has older siblings	0.570	0.569	(0.828)

— Note: We present the proportion of the sample with the corresponding characteristic. Columns one and two split the data in sub-samples by child's gender, as is indicated in the column headings. Standard deviations are not provided but are easily calculable using the proportion and the number of observations. The NLSCY sample weights, designed to accurately reflect the make up of the Canadian population, are applied in these and all calculations throughout the paper. Finally, we examine whether the sample of girls and boys have different proportions of a given characteristic and thus present the p-values from this statistical test in column 3. ***, ** and * indicate significance at the 1%, 5% and 10% level respectively.

Table 2: Dependent Variable Summary Statistics

	Ages	Girls		Boys		P-Value
		Obs.	Mean (Std.Dev)	Obs.	Mean (Std.Dev)	
Child Care and Work Decisions						
In Care	0-4	18591	0.418 (0.498)	19575	0.413 (0.498)	(0.771)
Other's Home	0-4	18591	0.2242 (0.417)	19575	0.2364 (0.416)	(0.404)
Own Home	0-4	18591	0.0737 (0.295)	19575	0.0699 (0.296)	(0.671)
Center Based Care	0-4	18591	0.1137 (0.337)	19575	0.1059 (0.341)	(0.473)
Hours In Care	0-4	18568	13.587 (18.707)	19529	13.828 (19.083)	(0.724)
Full Time Care	0-4	18568	0.3231 (0.471)	19529	0.3304 (0.473)	(0.652)
Mother Works	0-4	18740	0.5383 (0.488)	19731	0.5222 (0.489)	(0.353)
Mother Works / Uses Child Care	0-4	18530	0.362 (0.49)	19514	0.3496 (0.49)	(0.454)
Mother Works / Does Not Use Child Care	0-4	18530	0.1747 (0.405)	19514	0.1719 (0.4)	(0.83)
Mother Does Not Work / Uses Child Care	0-4	18530	0.0557 (0.222)	19514	0.063 (0.227)	(0.371)
Mother Does Not Work / Does Not Use Child Care	0-4	18530	0.4076 (0.474)	19514	0.4155 (0.475)	(0.644)
Child Development, Behavior, and Health Outcomes						
MSD Score	0-3	14824	101.42 (14.702)	15648	97.07 (15.209)	(0.000)***
PPVT Standardized Score	4	3265	101.40 (15.22)	3320	99.94 (15.085)	(0.000)***
Hyperactivity and Inattention Score	2-3	7118	3.372 (2.305)	7555	3.833 (2.434)	(0.000)***
Emotional Anxiety Score	2-3	7175	1.158 (1.424)	7606	1.247 (1.504)	(0.000)***
Physical Aggression Score	2-3	7098	4.683 (2.825)	7527	5.044 (3.033)	(0.000)***
Separation Anxiety Score	2-3	7180	2.590 (1.946)	7632	2.661 (2.001)	(0.028)**
Child in Excellent Health	0-4	18781	0.681 (0.466)	19758	0.635 (0.481)	(0.000)***
Never had a Nose/Throat Infection	0-4	15197	0.469 (0.499)	16055	0.435 (0.496)	(0.000)***
Never had an Ear Infection	0-4	15187	0.549 (0.498)	16043	0.501 (0.5)	(0.000)***
Has been Injured	0-4	18779	0.070 (0.255)	19752	0.089 (0.284)	(0.000)***

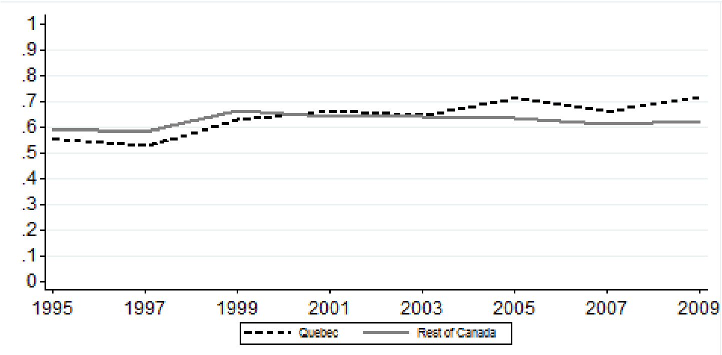
Table 2: Dependent Variable Summary Statistics

	Ages	Girls		Boys		P-Value
		Obs.	Mean (Std.Dev)	Obs.	Mean (Std.Dev)	
Parent Well-Being Outcomes						
Mother in Excellent Health	0-4	18687	0.3982 (0.49)	19678	0.393 (0.488)	(0.291)
Father in Excellent Health	0-4	18645	0.4195 (0.493)	19654	0.4135 (0.492)	(0.233)
Mother's Depression Score	0-4	16597	4.0512 (4.669)	17408	4.1333 (4.702)	(0.106)
Parenting Outcomes						
Family Dysfunction Index	0-4	18451	7.888 (5.068)	19391	7.862 (5.109)	(0.616)
Ineffective Parenting	2-4	10557	8.540 (11120)	11120	9.030 (3.6)	(0.000)***
Parent Consistency	2-4	10441	14.950 (10990)	10990	14.970 (3.2)	(0.653)
Positive Interaction	2-4	10321	16.080 (11258)	11258	16.110 (2.52)	(0.381)
Spends 5 minutes of focused time - many times a day	0-3	15054	0.720 (0.449)	15887	0.715 (0.451)	(0.181)
Laughs with child - many times a day	0-3	15054	0.833 (0.373)	15880	0.830 (0.375)	(0.267)
Does a special activity that the child enjoys - Once or twice a day or more	0-3	15007	0.636 (0.481)	15841	0.647 (0.478)	(0.02)**
Plays a sport, game, or hobby with child - Once or twice a day or more	0-3	15042	0.744 (0.437)	15877	0.759 (0.428)	(0.001)***
Reads to child - daily	0-3	9056	0.689 (0.463)	9582	0.664 (0.472)	(0.000)***
Spends 5 minutes of focused time - many times a day	4	3554	0.508 (0.5)	3672	0.490 (0.5)	(0.062)*
Laughs with child - many times a day	4	3553	0.730 (0.444)	3672	0.693 (0.461)	(0.000)***
Does a special activity that the child enjoys - Once or twice a day or more	4	3554	0.332 (0.471)	3670	0.358 (0.479)	(0.011)**
Reads to child - daily	4	3522	0.687 (0.464)	3627	0.634 (0.482)	(0.000)***

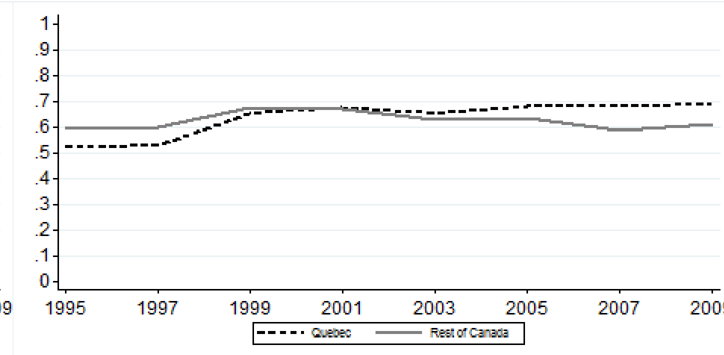
— Note: Each row corresponds to a variable of interest, with column 1 indicating for which age group the variable is measured. Columns 3 and 5 contain the mean and standard deviation (in parentheses) specific to the gender specific sub-sample, as identified in the column header. The corresponding sample size for these statistics are presented in columns 2 and 4. Finally, we test for difference in the mean (or proportion) between girls and boys for each variable of interest and present the p-values from this statistical test in column 6. ***, ** and * indicate significance at the 1%, 5% and 10% level respectively.

Figure 1: Trends in Uptake and Development by Child Gender

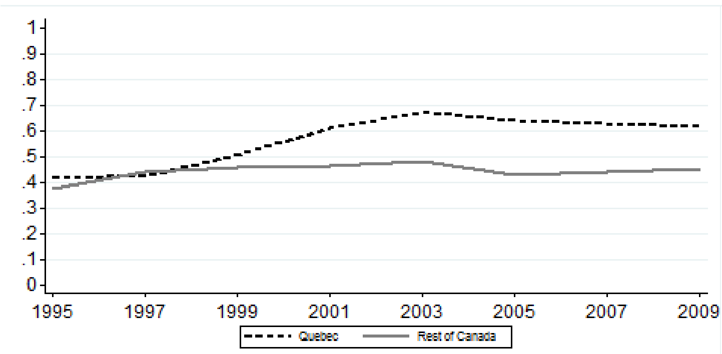
(a) Girls: Proportion of Mother's Working



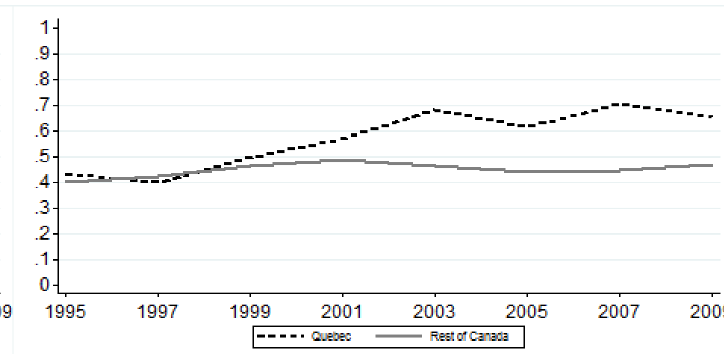
(b) Boys: Proportion of Children in Care



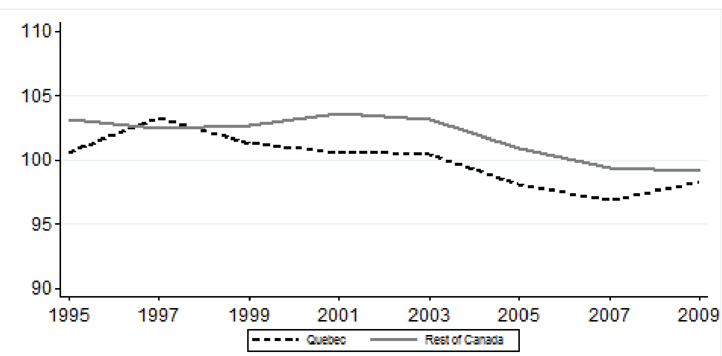
(c) Girls: Proportion of Mother's Working



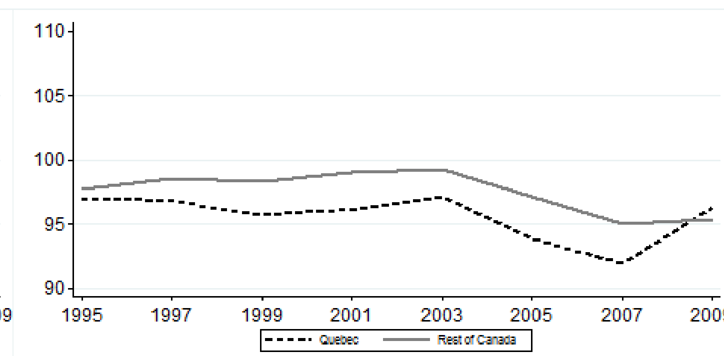
(d) Boys: Proportion of Mother's Working



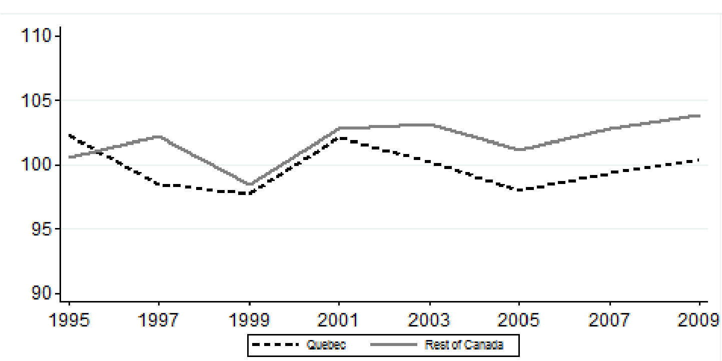
(e) Girls: Motor and Social Development Score



(f) Boys: Motor and Social Development Score



(g) Girls: Peabody Picture and Vocabulary Test



(h) Boys: Peabody Picture and Vocabulary Test

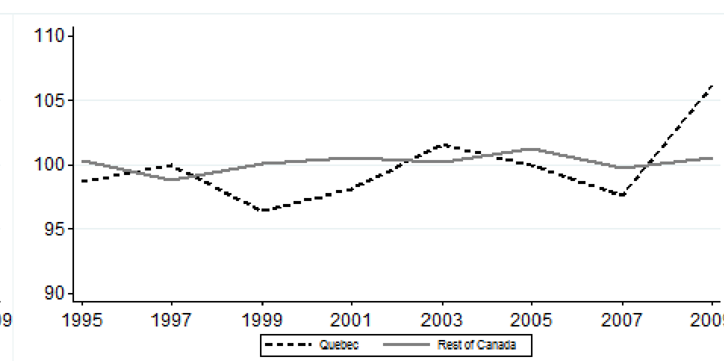


Table 3: Estimates of the Causal Effect of Access to Universal Child Care by Gender

	Intention to Treat	
	Girls (P-Value)	Boys (P-Value)
Child Care and Work Decisions		
In some type of care	0.186	0.205
	(0.000)***	(0.000)***
Care in Another's Home	**-0.007**	**-0.055**
	(0.788)	**(0.001)*** **
Care in own home	-0.028	-0.007
	(0.031)**	(0.617)
Care in institutional setting	**0.224**	**0.263**
	(0.000)* **	**(0.000)*** **
Hours in all Child Care Arrangements	8.136	8.874
	(0.000)***	(0.000)***
In full time care - More than 20 hours	0.221	0.222
	(0.000)***	(0.000)***
Mother Works	0.095	0.122
	(0.000)***	(0.000)***
Mother Works / Uses Childcare	0.143	0.165
	(0.000)***	(0.000)***
Mother Works / Does not use Childcare	-0.046	-0.049
	(0.000)***	(0.000)***
Mother does not Work / Uses Childcare	0.046	0.041
	(0.000)***	(0.000)***
Mother does not Work / Does not use Childcare	-0.143	-0.157
	(0.000)***	(0.000)***

— Note: For the outcome variable in each row we present the estimates of the policy coefficient δ as specified in Equation (1). We split the sample by gender as denoted in the column header. These regressions also include a set of dummies derived from the covariates listed in Table 1 as well as province and cycle indicators. We test the reported coefficients for statistical difference from zero using a two-tailed test and report adjusted p-values (presented in parentheses) corresponding to the estimate in the row above. These p-values make use of a Simes p-value adjustment procedure to account for testing effects on multiple related outcomes. The standard errors underlying the hypothesis tests are also corrected at the province-year level. ***, ** and * indicate significance at the 1%, 5% and 10% level respectively. Statistically different estimates between girls and boys are presented in bold face.

Table 4: Estimates of the Causal Effect of Access to Universal Child Care on Distributional Moments of Child Outcomes

	Girls			
	Mean	Variance	Skewness	Kurtosis
MSD Score	-1.56	**67.531**	-1.217	**6.489**
	(0.198)	**(0.000)*** **	(0.002)***	**(0.004)*** **
PPVT Standardized Score	-0.912	**34.220**	-0.697	**0.319**
	(0.449)	**(0.266)**	(0.087)*	**(0.800)**
Hyperactivity and Inattention Score	0.136	**-0.600**	**0.048**	**-0.549**
	(0.449)	**(0.228)**	**(0.859)**	**(0.475)**
Emotional Anxiety Score	0.333	0.393	0.221	-1.330
	(0.002)***	(0.187)	(0.732)	(0.684)
Physical Aggression Score	0.718	1.203	**0.416**	**0.826**
	(0.000)***	(0.133)	**(0.156)**	**(0.343)**
Separation Anxiety Score	0.21	**-0.044**	0.204	**0.134**
	(0.079)*	**(0.899)**	(0.435)	**(0.852)**
	Boys			
MSD Score	-1.74	**7.001**	-0.377	**0.379**
	(0.004)***	**(0.704)**	(0.235)	**(0.796)**
PPVT Standardized Score	-0.033	**-49.452**	0.241	**-3.623**
	(0.977)	**(0.106)**	(0.576)	**(0.015)** **
Hyperactivity and Inattention Score	0.511	**0.591**	**0.847**	**1.136**
	(0.000)***	**(0.287)**	**(0.002)*** **	**(0.138)**
Emotional Anxiety Score	0.1	0.327	0.559	2.162
	(0.506)	(0.290)	(0.341)	(0.418)
Physical Aggression Score	0.527	1.360	**1.026**	**2.149**
	(0.000)***	(0.128)	**(0.000)*** **	**(0.008)*** **
Separation Anxiety Score	0.13	**-1.331**	-0.431	**-2.471**
	(0.323)	**(0.001)*** **	(0.130)	**(0.002)*** **

— Note: For the outcome variable in each row we present the estimates of the policy coefficient δ as specified in Equation (1) for the distributional statistic denoted in the column header. We split the sample by gender as denoted in the column header. These regressions also include a set of dummies derived from the covariates listed in Table 1 as well as province and cycle indicators. We test the reported coefficients for statistical difference from zero using a two-tailed test and report adjusted p-values (presented in parentheses) corresponding to the estimate in the row above. The standard errors underlying the hypothesis tests are also corrected at the province-year level. ***, ** and * indicate significance at the 1%, 5% and 10% level respectively. Statistically different estimates between girls and boys are presented in bold face.

Table 5: Estimates of the Causal Effect of Access to Universal Child Care on Parenting and Family Outcomes

	Intention to Treat	
	Girls (P-Value)	Boys (P-Value)
Mother in Excellent Health	-0.015	-0.013
	(0.578)	(0.536)
Father in Excellent Health	0.007	-0.01
	(0.761)	(0.694)
Mother's Depression Score	0.846	0.483
	(0.002)***	(0.049)**
Family Dysfunction Index	**-0.207**	**0.465**
	(0.578)	**(0.012)****
Ineffective Parenting	0.943	0.602
	(0.000)***	(0.000)***
Parent Consistency	**-0.509**	**0.041**
	(0.012)**	**(0.799)**
Positive Interaction	**-0.821**	**-0.474**
	(0.000)***	**(0.014)****
Ages 0-3		
Spends 5 minutes of focused time - many times a day	**-0.119**	**-0.069**
	(0.017)***	**(0.029)****
Laughs with child - many times a day	**-0.079**	**-0.039**
	(0.014)***	**(0.012)*****
Does a special activity that the child enjoys	-0.061	-0.056
- Once or twice a day or more	(0.017)***	(0.011)***
Plays a sport, game, or hobby with child	0.006	0.018
- Once or twice a day or more	-0.018	(0.010)*
Reads to child - daily	**-0.085**	**-0.032**
	(0.022)***	**-0.024**
Estimated Days Read to a Month	-1.56	-1.014
	(0.448)***	(0.587)*
Age 4		
Spends 5 minutes of focused time - many times a day	**0.084**	**-0.078**
	(0.025)***	**(0.042)***
Laughs with child - many times a day	**-0.029**	**-0.165**
	-0.022	**(0.036)*****
Does a special activity that the child enjoys	-0.097	-0.066
- Once or twice a day or more	(0.041)**	-0.045
Reads to child - daily	0.016	-0.05
	-0.031	-0.053
Estimated Days Read to a Month	**-1.005**	**-2.479**
	-0.77	**(1.097)****

— Note: For the outcome variable in each row we present the estimates of the policy coefficient δ as specified in Equation (1) first without weighting (Intention to Treat) and second with the inverse propensity weights specified in Equation (2) (Composition Adjusted). For each specification we split the sample by gender as denoted in the column header. These regressions also include a set of dummies derived from the covariates listed in Table 1 as well as province and cycle indicators. We test the reported coefficients for statistical difference from zero using a two-tailed test and report adjusted p-values (presented in parentheses) corresponding to the estimate in the row above. These p-values make use of a Simes p-value adjustment procedure to account for testing effects on multiple related outcomes. The standard errors underlying the hypothesis tests are also corrected at the province-year level. ***, ** and * indicate significance at the 1%, 5% and 10% level respectively. Statistically different estimates between girls and boys are presented in bold face.

Table 6: Exploration of Mediating Effects of Parenting Activities and Behaviours

	Girls				
	No Additions	Parenting Activities	Parenting Scales	Reading	All
MSD Score	-1.543	-0.938	-0.541	-1.230	-0.810
	(0.128)	(0.352)	(0.570)	(0.231)	(0.413)
PPVT Standardized Score	-0.931	-0.948	-1.192	-0.809	-1.144
	(0.363)	(0.326)	(0.222)	(0.405)	(0.216)
Hyperactivity	0.326*	0.186	-0.012	0.264	-0.033
	(0.081)	(0.287)	(0.945)	(0.172)	(0.847)
Emotional Anxiety	0.274***	0.222**	0.165*	0.263***	0.164*
	(0.002)	(0.011)	(0.064)	(0.004)	(0.074)
Physical Aggression	0.501***	0.343*	0.062	0.445**	0.050
	(0.006)	(0.063)	(0.745)	(0.021)	(0.796)
Separation Anxiety	0.206	0.134	0.045	0.180	0.036
	(0.102)	(0.292)	(0.712)	(0.179)	(0.769)

	Boys				
	No Additions	Parenting Activities	Parenting Scales	Reading	All
Males					
MSD Score	-2.124***	-1.849***	-1.572*	-1.962***	-1.745**
	(0.007)	(0.010)	(0.058)	(0.007)	(0.021)
PPVT Standardized Score	-0.075	0.225	0.106	0.377	0.306
	(0.957)	(0.862)	(0.936)	(0.769)	(0.804)
Hyperactivity	0.692***	0.613***	0.607***	0.630***	0.577***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Emotional Anxiety	0.068	0.070	0.029	0.074	0.050
	(0.399)	(0.403)	(0.665)	(0.380)	(0.493)
Physical Aggression	0.626***	0.519**	0.491**	0.556***	0.450**
	(0.002)	(0.014)	(0.012)	(0.005)	(0.013)
Separation Anxiety	0.251***	0.194**	0.258***	0.201**	0.214***
	(0.001)	(0.030)	(0.000)	(0.018)	(0.004)

— Note: We present the intent-to-treat estimates of the policy coefficient δ as specified in Equation (1) in the first column. In subsequent columns we add the additional parenting variables to check for any mediation of the policy effect. We check parenting activities, parenting scales, and reading separately in columns 2,3, and 4 and then combine all additional variables together in column 5. The parenting activities included in column 2 are spends 5 minutes of focused time many times a day, laughs with child many times a day, does a special activity that the child enjoys once or twice a day or more, plays a sport, game, or hobby with child once or twice a day or more, and reads to child daily. In column 3, we include the family dysfunction, ineffective parenting, parental consistency, and positive interaction scales. Column 4 adds the indicator variable for reading to the child daily only. Note we use a sample that is consistent across the addition of all of the variables, accounting for the differences in the estimates in column 1 and those in Table 3. We test the reported coefficients for statistical difference from zero using a two-tailed test and report adjusted p-values (presented in parentheses) corresponding to the estimate in the row above. The standard errors underlying the hypothesis tests are also corrected at the province-year level. ***, ** and * indicate significance at the 1%, 5% and 10% level respectively. In this table we do not test for significant differences between girls and boys.

Table A1: Testing the Common Trend Assumption

	In some type of care	Mother Works	PPVT Standardized Score	MSD Score
Girls	0.003	-0.031	-1.164	-0.157
	(0.807)	(0.000)***	(0.124)	(0.775)
Boys	-0.003	-0.043	0.077	-0.839
	(0.862)	(0.000)***	(0.888)	(0.006)***

— Note: For the outcome variable in each row we present the test of the common trend assumption as laid out in the following equation: $Y_{ipt} = \beta_0 + \beta_1 YEAR_t + \beta_2 YEAR_t * QUE + \varepsilon_{ipt}$. Using data in the period before the policy implementation we report the coefficents β_2 for each corresponding outcome and test these coefficients for statistical difference from zero using a two-tailed test. P-values are presented in parentheses corresponding to the estimate in the row above. ***, ** and * indicate significance at the 1%, 5% and 10% level respectively.