The Effects of Maternity leave on Children's Abiliites. Evidence from a Maternity Leave Reform in Chile

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This paper estimates the effects of maternity leave on several children's outcomes exploiting the exogenous variation induced by a maternity leave reform in Chile. The reform increased the paid leave from 12 to 24 weeks for mothers of children born on July 25th, 2011 (and after). Using the eligibility status as an instrumental variable for the duration of the leave we estimate the effects of maternity leave on children's cognitive and noncognitive abilities, as well as some mother outcomes such as parental stress, employment, and wages. We find significant and positive effects of maternity leave on children's cognitive abilities and motor skills, specially for children with less educated mothers. We find also a significant reduction on parental stress index for mothers and an increase in the probability of being employed after the leave but no effect on wages. We identify an increase of breastfeeding for exposed children as one channel driving the results.The evidence supports recent findings on the effects of maternity leave on children's outcomes on developed countries.

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I. Introduction

Early childhood has taken a central role in the human capital literature over the last decade. Investments during this stage are fundamental to the human capital accumulation process, especially for cognitive skills (Cunha and Heckman, 2007, 2008; Cunha, Heckman and Schennach, 2010), and they explain a wide range of adult outcomes (Heckman, Stixrud and Urzua, 2006; Almond and Currie, 2011). Parental care (or time spent with the child) is usually thought as an important input to the human capital production function, and one policy tool aiming at increasing it is paid maternity leave (Dahl et al., Forthcoming).

Maternity leave legislation varies widely across countries in terms of duration of the leave, job protection, and income replacement. Over the last 20 years there has been a global trend towards longer paid leave periods (meeting or exceeding the International Labour Organization (ILO) standard of 14 weeks), higher levels of income replacement and less reliance on employer liability (ILO, 2014). Although it is commonly believed that longer leaves should have a positive effect on children's skill development, the empirical evidence is still scarce, based on developed countries only, and has mixed results.[1]

This paper estimates the effects of maternity leave on cognitive and noncognitive children abilities and some mother outcomes exploiting the arguably exogenous variation induced by a maternity leave reform in Chile. In 2011 the 12 weeks paid leave was extended to 24. Children older than 24 weeks at the date of the reform were not eligible for the extended leave. This allows us to use children's eligibility status as an instrumental variable for the effective duration of the leave. This is relevant since, as argued by Klerman and Leibowitz (1997), the timing of return to work after childbirth (i.e. the duration of maternity leave) is a mother's decision. As such, it will be likely influenced by some unobserved factors correlated with child's development. For instance, there could be a reverse causal relation if mothers of disadvantaged children (in terms of initial endowment) invest more time to compensate for this gap. This would generate a downward bias in the Ordinary Least Squares (OLS) estimator and Instrumental Variable estimation helps us to reduce this bias.

The outcomes analyzed in this paper include children's cognitive and noncognitive skills as measured by the Child Learning and Development Test (*Test de Aprendizaje y Desarrollo*

[1]While Carneiro, Løken and Salvanes (2015) find positive effects, other studies find no overall effect (Baker and Milligan, 2010; Dustmann and Schönberg, 2012; Liu and Skans, 2010; Würtz, 2010; Danzer and Lavy, 2013; Dahl et al., Forthcoming), and some even find negative effects (Baker and Milligan, 2011). When addressing the possibility of heterogeneous effects, some articles find a positive impact for children of highly educated mothers (Liu and Skans, 2010; Danzer and Lavy, 2013) and others for children of less educated mothers (Carneiro, Løken and Salvanes, 2015).

Infantil, TADI) that was developed in Chile and evaluates four dimensions: cognitive, language, motor, and socioemotional. To the best of our knowledge, this is the first paper that evaluates these type of measures and, thus, is novel to the literature.

Our identification strategy is similar to that of Baker and Milligan (2011) who use children's exposition to a Canadian maternity leave expansion as an instrument for the time spent with the child during her first year of life. However, our findings are different from theirs. While they find no effects on non-cognitive skills and a small but negative effect on cognitive ones, we find positive and statistically significant effects on cognitive and motor skills and no effects on language and non-cognitive (socioemotional) skills. Certainly the counterfactual of maternity leave in Chile and Canada is different (such as the quality of public and private childcare). There are also differences in the outcome variables analyzed and the children's age of exposure in the Chilean and Canadian reform. Their measure of non-cognitive skill is solely based on parental report, and thus it is exposed to a systematic measurement error.[2] Instead we use formal psychological tests that were constructed and standardized with a Chilean sample and, thus are more precise measures of children abilities. Another difference in our study is the age at which the expansion of maternity leave affected the children. While the reform analyzed by Baker and Milligan affected them during the second half of their first year of life, ours did during the first six months. As they note, their results could be driven by the development of stranger and separation anxiety, which are usually observed during the second six months of life. In contrast, our results are not exposed to this issue.

Our results show that mothers exposed to the reform significantly extend their paid maternity leave (it does not crowd out unpaid leave) and they support the idea of a reverse causal relation between abilities and leave duration. While OLS estimates suggest null effects, 2SLS estimates reveal positive gains of 7% and 6% of a standard deviation on the cognitive and motor skills test scores, respectively. These results are mainly driven by children of mothers with low education and are likely to remain in the long run. As pointed out by Elango et al. (2014), "*Early Childhood Interventions strongly boost IQ in the short-run for disadvantaged children. However, treatment effects often fade out. For very early interventions (before age 3), impacts persist into adulthood.*"[3] As mentioned above, the exposure to the reform occurred during the first 6 months of life, so these children were quite younger than three

[2] Parents who spend more time with their children should be more aware of their skill development.

[3] Early Childhood Interventions often target especially disadvantaged subpopulations. We analyze a broader treatment applied to children of labor force participating mothers, regardless of their socioeconomic status.

years.

We analyze the mechanisms through which the effects manifest. In addition to the arguably increased time spent with the mother (as proxied by maternity leave) we identify breastfeeding as another mechanism. We estimate duration models for breastfeeding length finding positive and significant effects for those exposed to the reform compared to those not exposed. This has a biological channel (see Victora et al. (2015) and Belfield and Kelly (2012)) but it also supports the argument relating maternity leave and time spent with the child. An increase in breastfeeding duration necessarily increases time spent with the child and mother-infant bonding.

Lastly, we also look at some mothers' outcomes such as the probability of being employed and wage at the time of the survey, and the Parent Stress Index (PSI). While we find no effects of the reform on wages, there is a small but positive and significant effect on the probability of being employed, and we find a reduction in the incidence of sick child leave use between the third and sixth month of life of the child. This suggests that at least part of the cost of the subsidy for maternity leave is paid by a reduction in sick child leave subsidy payments. Furthermore, we find a reduction in the PSI as measured at the time of the survey. This result shows that maternity leave legislation can also improve mothers' health.

The paper is organized as follows. Section II describes the reform that extended maternity leave in Chile. In section III, we explain in detail our instrument and our identification strategy. Data are described in section IV. Section V presents and discusses the main results. Finally, section VI concludes.

II. Institutional Background

Maternity leave in Chile was first enacted in 1925. Before October 17, 2011, legal maternity leave consisted of 12 mandatory weeks with full income replacement.[4] Job protection continued until one year after completion of the leave. Law No. 20.545 extended this permit for another additional 12 full time or 18 part-time weeks (in case the beneficiary decides to come back to work after the 12 mandatory weeks) through the establishment of the Parental Leave. This new leave operates under the same conditions as the former.[5] However, women

[4]The benefit limit is about USD \$2,640 at an exchange rate of USD/CLP 0.0016 and is slightly above the 95^{th} percentile of the women wage distribution aged 25 to 55 in 2011 (calculations based on the National Socio Economic Characterization Survey, CASEN, 2011).

[5]Just as the former leave, this new Parental Leave covers formal workers who have a signed employment contract and are contributors to the social security system. It may also cover independent workers, but under some requirements which are difficult to meet in the short run. To become eligible, independent workers must have at least one year

are allowed to return to work on a part-time arrangement after the 12 mandatory weeks and they are allowed to transfer a fraction of their extended leave to the father. Thus, the extension of the leave varies.[6]

The reform implementation considered a transition period during which it also covered mothers who were using or had finished their mandatory maternity leave on October 17, 2011, up until the child turned 24 weeks old. This means that children born after May 2, 2011 were eligible for the extended leave to some extent, varying from 0 to 12 extra full-time weeks. Those born on May 2 turned 24 weeks on October 17, so their mothers had access to 0 extra weeks. Those born between May 2 and July 25 were older than 12 weeks but younger than 24, so their mothers had access to an extra time $0 < t_i < 12$. Finally, mothers of children born after July 25 had access to the full 12 additional weeks. Hence, the interpretation of the instrumental variables estimation can be viewed as a Local Average Treatment Effect (LATE) for women induced to extend their maternity leave by the reform (compliers).

We argue about the exogeneity of the enactment date of the reform. Extending the leave was a longstanding idea. Some projects were submitted over the last decades but none of them was approved.[7] The idea was so rooted that there were even some prospective evaluations, for example Aedo (2007). The legislative process took almost eight months. It began on February 28, 2011 with a presidential message. The law was enacted on October 6, 2011 and published in the Official Journal on October 17, 2011. With all of this, the date of the reform seems not to be correlated to children's development through any mechanism other than the duration of the leave. Thus, this reform generates suitable exogenous variation to instrument the duration of maternity leave to estimate its impact on child development, as detailed in the following section.

There are also other characteristics of this reform that make it suitable for our analysis. Baker and Milligan (2011) argue that stranger anxiety and separation anxiety are both observed during the second half of the first year of life and they could compensate the effect of

of affiliation to the social security system and at least six (continuous or discontinuous) contributions during the six months prior to prenatal leave. Before the reform independent workers had no strong incentives to enter the social security system. The reform might induce some women to start contributing, anticipating the benefits of the subsidy. However, the time span we study is too short for this behavior to be relevant.

[6]Mothers are allowed to return to work on a half workday basis after the 12 mandatory weeks of leave, which is an inalienable right. If the part-time leave is chosen, the length of the extension increases to 18 weeks (instead of 12) and the subsidy is reduced by half. This means that, after the reform, female workers can take a maximum of 24 weeks in case of full time Parental Leave or 30 weeks in case of part-time Parental Leave (12 full time leave plus 18 part-time leave). Mothers can also transfer some weeks to the father: up to 6 weeks in case of full time, or up to 12 weeks in case of part-time leave. In any case, the weeks used by the father must be the last weeks of the leave. The subsidy for the father is calculated based on his income, up to the same limit of USD $2,640. By this means, even though the first 12 weeks of leave remain inalienable, the additional 12 (or 18) weeks are not completely mandatory.

[7]For more information see the "*Trámite de Proyectos*" (processing of projects) section on the Senate's website, `www.senado.cl`.

additional maternal care. Our reform is not exposed to this issue. As far as we know, this is one of the earliest treatments studied in the maternity leave literature, covering the period between the third and sixth month of life. Another advantage of this reform is its homogeneity regarding income substitution. The extended leave is paid, just as the pre-reform leave, producing no income effects.[8] This is also the first paper to estimate the effects of maternity leave in the context of developing countries.[9]

There is also a sick child leave for mothers of children under age one. It has been reported that some mothers might use this leave as a mechanism to extend their maternity leave beyond the legal permission (SUSESO, 2012; Congress, 2012). To assess the sensibility of our results to changes in sick child leave patterns we construct an alternative measure of maternity leave that includes the sick child leave weeks and estimate the first stage. As we discuss later, the incidence of sick child leave is small.

III. Empirical Strategy

We consider the estimation of the following structural equation

$$Y_i = X_i'\gamma + \beta L_i + u_i \tag{1}$$

where Y_i denotes a developmental outcome of child i and L_i the duration of maternity leave, is exposed to a potential endogeneity problem. As discussed (for example) by Bernal and Keane (2010); Baker and Milligan (2011); Carneiro, Løken and Salvanes (2015); Danzer and Lavy (2013), mothers' choice of L_i might depend on some unobserved factors that affect child development. If mothers of children with lower initial skill endowments take longer leaves as a compensation mechanism, there will be a reverse causal relation. In this case the OLS estimator of β will be inconsistent and downward biased.

Thus, we follow an instrumental variable approach in which we exploit exposure to the

[8]As we commented on footnote 6, the limit of $2,640 to the subsidy represents a higher income than the 95^{th} percentile of the income distribution of Chilean women aged 25 to 55 in 2011. It is also close to the 99^{th} percentile of mother's income distribution in our sample.

[9]Carneiro, Løken and Salvanes (2015) study a reform that extended the leave in Norway from 12 weeks of unpaid leave to 4 months of paid leave and 12 months of unpaid leave, thus covering children between their third and twelfth month. Dahl et al. (Forthcoming) study a series of posterior Norwegian reforms that expanded paid leave from 18 to 35 weeks. Baker and Milligan (2010, 2011) study a Canadian reform that extended the paid leave from 25 to 50 weeks. Dustmann and Schönberg (2012) study three reforms in Germany: the first extended the job-protected leave from 8 weeks to 6 months, the second from 6 to 10 months, and the last from 18 to 36 months (there were other reforms in between). Liu and Skans (2010) analyze a Swedish reform that extended the paid leave from 12 to 15 months. Würtz (2010) studies a reform that extended the paid leave from 14 to 20 weeks in Denmark. Finally, Danzer and Lavy (2013) study a reform that extended the paid leave from 12 to 24 months in Austria.

reform as an instrument for maternity leave duration (L_i). As described in section II, child's exposure to the extended leave is determined by the difference between her date of birth and the date of the reform. This variable is plausibly exogenous with respect to the human capital accumulation process. Besides, the reform should induce some mothers to choose a longer leave.

Formally, let e_i be the degree of exposure to the reform of mother i measure by the number of eligible weeks (on top of the regular leave). Hence, $e_i \in [0, 12]$ with 0 meaning non exposed and 12 fully exposed. Since there was a transition period we observe about 40% of mothers partially exposed to reform with $e_i \in (1, 11)$. For the sake of interpretation, we use define our instrumental variable $Z_i = e_i/12$, thus it varies from 0 to 1. Certainly this transformation does not alter the result. Another alternative is to use a binary instrument. We present results with a binary instrument in the appendix. They are robust to the instrument definition.

Then, we estimate the following triangular system by 2SLS.

$$Y_i = X_i'\gamma + \beta L_i + u_i \tag{2}$$

$$L_i = X_i'\delta + \pi Z_i + v_i \tag{3}$$

where X_i is a vector of covariates including child and mother's age, number of people in the household, and indicators of mother's education level, region, child's (male) gender, urban area, father's presence, student mother, and mother's civil status.

Under the standard IV assumptions (clean and fully correlated instruments), β is identified. In a heterogeneous treatment effects framework, under an additional monotonicity assumption, it can be interpreted as a local average treatment effect for the compliant subpopulation, i.e., the children of mothers who were induced to choose a longer leave by the reform (Imbens and Angrist, 1994). As pointed out by Angrist and Pischke (2009), in this framework the linear response function assumption implied by model (2)-(3) need not be taken literally, for the identified parameter is a weighted average of the unit causal responses. Now we turn to the plausibility of the IV assumptions.

The exclusion restriction would be violated if the date of the reform was chosen for some reason related to child development. However, the length of the legislative process and the longstanding intention to extend the leave point in the opposite direction. The political process delayed the entry into force of the law. During the Congress sessions the prospect of a maternity leave extension was not controversial. The discussion focused on other details, especially the limit of USD $1,200 to the new subsidy that was considered in the initial bill. It was voted to delete the sentence that established this limit, giving rise to a three months episode that delayed the enactment of the reform (Congress, 2012).[10] This dilation was arguably not related to children development.[11]

Another threat to the exclusion restriction might occur if there was manipulation of the birth dates. If certain mothers postponed their childbirths anticipating the reform, and if they were systematically different form the rest in some unobserved aspect related to children's development, this assumption would not hold. However, the argument explained above applies to this case too. The date of enactment of the law was not readily predictable so it seems sensible to assume its exogeneity. In the appendix we present evidence on the evolution of weekly births suggesting no bunching around the enactment date of the law.

The first stage assumption is reasonably plausible. Microeconomic theory suggests that an extension of the legal permission reduces the number of women quitting their jobs to extend their leave and increases the number of women on leave (Klerman and Leibowitz, 1997). Thus, the coefficient of the instrument in the first stage, π, should be positive and statistically significant. Our results (shown in section V) are consistent with this hypothesis. Similarly, the monotonicity assumption should hold by a revealed preference argument. The extension of the legal permission represents an expansion of the feasible choice set. Some mothers might choose not to change the duration of their leave and some might choose a longer leave that was not feasible before. But no rational mother would choose a shorter duration that was feasible and not chosen in absence of the reform.

We also estimate model (2)-(3) for a set of mother outcomes (separately), namely, a measure of parental stress, an indicator of being employed at the time of the survey, and her last reported wage. The specification of vector X_i varies depending on the response variable.[12] The arguments explained above regarding the validity of the IV assumptions apply for these response variables too.

[10] The elimination of that sentence set the limit at the default of USD $2,640 established for the former leave. A heated argument followed, as this implied an additional expenditure and the Constitution establishes that it is an exclusive presidential attribution. The President argued for the unconstitutionality of the project. The presidential request was finally accepted after a long process, but then the President set the limit to USD $2,640 again.

[11] This kind of exogeneity argument has been used, for instance, for the case of California's Paid Family Leave reform. See Huang and Yang (2015) for details.

[12] Let X_{1i} be a vector including child and mother's age, gender of the child, and indicators of urban area, father's presence, region, and mother's education. In the PSI estimation we use X_{1i}, number of people in the household, and an indicator for the mother being a student. For the employment status specification the covariates are X_{1i}, number of people in the household, and dummies for student mother and mother's civil status. Finally, in the wage equations we add month and year dummies to the last set of covariates.

IV. Data

We use the Early Childhood Longitudinal Survey (*Encuesta Longitudinal de La Primera Infancia*, ELPI) data. The ELPI is a nationally representative survey of children designed to study early childhood and support public policy evaluation (Behrman, Bravo and Urzúa, 2010). The first wave, conducted in 2010, covers approximately 15,000 children born between January 1, 2006 and August 31, 2009. The second wave, conducted in 2012, covers these children plus a new sample of approximately 3,000 children born between September 1, 2009 and December 31, 2011. This survey contains a rich set of information comprising family background, child care, and mother's labor market outcomes. It also includes test scores on a formal psychological evaluation of children's development.

Regarding children outcomes, the second wave contains children cognitive and noncognitive skills as measured by the Child Learning and Development Test (*Test de Aprendizaje y Desarrollo Infantil*, TADI). The latter was developed in Chile and evaluates four dimensions of children development: cognitive, language, motor, and socioemotional abilities.

Using the 2012 wave of ELPI we are able to identify children exposed and not exposed to the reform since we have the exact date of birth. Furthermore, we have a measure of total maternity leave of eligible mothers (excluding mothers out of the labor force at childbirth). There is a specific subsection of the survey in which there are questions about the extension of maternity leave both prior and after childbirth. In particular the question: *how many weeks of maternity leave did the mother of the chosen child take?* The question asks about the "chosen child" since the ELPI survey focused on the children. As noted, the question does not refer to paid or unpaid leave so we interpret it as a total leave measure. We will get back to this point in the results section.

A small fraction of the sample (near 3%) choose the part-time parental leave and we considered as our measure of total leave the sum of full time and part-time weeks.[13] We focus on children up to six years. Taking the non-eligible mothers out of our analysis, we have a sample size of 3,458 observations.

We present some descriptive statistics for the sample of eligible mothers in Table 1. Roughly one third of them are highly educated (more than a high school diploma), 19% are

[13] We chose this specification in order to allow for decreasing marginal returns to maternity leave duration. In this scenario mothers spending an additional half of a workday with their children can have similar results than mothers spending a whole extra workday. We could have alternatively specified our measure of duration as the sum of full time weeks plus half the amount of part-time weeks, or we could have just used only full time weeks. Our results are not sensitive to any of theses changes in the specification, partly because less than 3% of mothers in our sample made use of the part-time leave. Similar results using alternative specifications are available upon request.

single mothers, 5% are students, and 79% were employed at the time of the survey. 91% were workers at childbirth and 96% of them (workers) had a formal employment contract. The incidence of father's presence reaches 72% and 96% of them were employed. The mean duration of maternity leave is 17 weeks for exposed children versus 13.5 weeks for non-exposed children.

Children were evaluated under the TADI (*Test de Aprendizaje y Desarrollo Infantil*) scale. It is a standardized test that allows the evaluation of development and learning of children between 3 months and 6 years.[14] The TADI evaluates four dimensions of development, each with its own score: Cognition, Language, Motor Skills and Socioemotional Development. We estimate model (2)-(3) for each dimension separately, using the corresponding T-scores as shown below, and for the total score (which is the mean of the four dimensional scores).[15] T-scores are distributed with mean $\mu = 50$ and standard deviation $\sigma = 10$ but we standardized them for the sake of interpretation.

TADI scores deliver some advantages as skill measures. First, the socioemotional score is a more precise measure of non-cognitive skills than the parent-reported indexes typically used in the literature. It is not exposed to the systematic measurement error that threatens the parent-reported measures. This allows a more precise estimation of the non-cognitive effect of maternity leave. Second, this scale was constructed and standardized with a Chilean representative sample, so it picks up the cultural background.

The ELPI survey also contains the Parenting Stress Index Short Form (PSI) scores of the mothers. We use the percentile scores of the PSI as our measure of maternal stress. Finally, we use a binary indicator for employment status of the mother (at the time of the survey) being equal to one for employed ones, and the last wage reported by the mother as labor market outcomes.

[14] This test was constructed by the Centre for Advanced Research in Education (CIAE) at the University of Chile in conjunction with the Centre for Studies in Development and Psychosocial Stimulation (CEDEP), under guidance from a multidisciplinary group of national and international experts. Its construction was based on the Early Learning Standards, an exhaustive literature review, and a set of development and learning evaluation instruments (Edwards and Pardo, 2013).

[15] The TADI scale is divided into 13 age groups. The T-scores indicate the relative position of the child within her reference population. For further information see Edwards and Pardo (2013).

V. Results

A. *First stage*

As mentioned before, we use exposure to the reform as an instrumental variable of maternity leave. The exposure is measure as the fraction of eligible days over the total number of days in regime. Thus, it varies continuously from 0 to 1 and the interpretation of the coefficient accompanying the instrument is the increase in weeks induce by the reform for a fully exposed mother.

To assess the effects of the reform on total leave (paid and unpaid) we present results for three measures of the dependent variable. Column (1) shows the result for the reported leave. This measure does not distinguish by paid or unpaid leave since the question in the survey does not indicate that type of distinction. Column (2) shows the results for the reported leave for those plus the number of weeks out of labor force after the reported leave and Column (3) shows the results when using reported leave plus the number of weeks due to sick child leave.[16]

Table 2 shows the first stage results. As can be seen, the coefficient on the excluded instrument has the expected sign: the reform induced some mothers to extend their leave. The estimated impact of the reform on the duration of the leave is an increase of roughly one month when using the self reported leave (4.17 weeks).[17] When using alternative measures the results does not change that much, showing increases of 3.73 and 3.21 weeks in columns (2) and (3). The magnitude of the effects found comparable to those in (Dahl et al., Forthcoming) where the increase of the leave was 3 weeks.

As we mentioned in Section 2, an impact of 12 additional weeks was not expected because of the non-mandatory nature of the reform and the fact that an important proportion of exposed children were in the transition period, where the leave could be extended only for the weeks remaining until the child turned 24 weeks old. Indeed, 60% of exposed mothers in our sample had access to the reform in regime. The remaining 40% was exposed to the transition, meaning they had access to a shorter extension. The mean potential extension for this group is 5.44 weeks. Additionally, 4% of mothers in our sample transferred some

[16] To construct the measure in Column (2) we add the number of weeks out of the labor force after the number of weeks of reported leave. We censored them in 52 weeks which is the largest leave in the sample. For sick child leave we have the number of leaves but not their duration. From administrative data we obtain the average duration of sick child leave by year and use them to construct the alternative measure in Column (3).

[17] Table B.1 in the appendix shows similar results for a binary specification of our instrument ($Z_i = 1$ for exposed children and $Z_i = 0$ for non-exposed children).

weeks to the father. Lastly, some substitution of paid for unpaid leave may occur. Since the question about the length does not specify the nature of the leave, there might be a sizable share of always takers that would have extended their legal (paid) maternity leave without the reform.

Regarding the strength of the instrument, we can see that the F-stat of the excluded instrument is 33.72 (reported leave), greater than the usual rule of thumb of 10. The Cragg and Donald (1993) statistic allows a formal test for weak instruments, using the critical values tabulated by Stock and Yogo (2005). Since we have only one endogenous variable, the Cragg-Donald statistic is just the F-stat of the excluded instrument. The critical value for the null hypothesis that the maximum actual size of the Wald test with nominal size 5% is greater than 10% is 16.38. Rejecting this null hypothesis, we conclude that our instrument is strong.

For the rest of the paper we use the first stage results from Column (1) given the strength of the instrument and the similarity in the results using the three alternative measures.

B. *Child Development*

As a benchmark for our 2SLS results, we show the results of the OLS estimation for TADI scores in Table 3. If we were to rely on OLS estimation, we would conclude that maternity leave has no effect on child development. However, exploiting the exogenous variation of maternity leave durations generated by our instrument, we show that this conclusion is not correct. Table 4 shows our 2SLS results for the TADI T-scores. We estimate a positive effect of 0.04 standard deviations on total TADI scores (significant at the 5% level), driven by effects of 0.07 and 0.06 standard deviations on the cognitive and motor dimensions respectively (both significant at the 1% level). We find no evidence of any impact on the language and socioemotional dimensions. The 2SLS coefficients are an order of magnitude greater than the OLS coefficients. Table B.2 shows similar results for the binary specification of the instrument, although the point estimates are somewhat bigger and the estimate for the language T-Scores becomes significant at the 10% level.

This result is consistent with the previously discussed hypothesis of a reverse causal relation and with other results in the (broader) maternal employment literature.[18] Using time-diary data and exploiting exogenous variation induced by shifts in the opportunity cost

[18]There is a large literature on the developmental effects of maternal employment during early childhood. See, for example, Blau and Grossberg (1992), Hill and O'Neill (1994), Waldfogel, Han and Brooks-Gunn (2002), Brooks-Gunn, Han and Waldfogel (2002), Ruhm (2004) and James-Burdumy (2005). Normally failing to account for endogeneity, this literature has produced mixed evidence (Bernal, 2008).

of time, Villena and Ríos (2012) obtain positive LIML estimates an order of magnitude greater than their OLS estimates.[19] Bernal (2008) and Bernal and Keane (2010) estimate structural models of maternal work and child care decisions. They find a statistically significant negative impact of maternal work on child development, which is consistent with our finding of a positive impact of maternity leave. Bernal and Keane (2011) use exogenous variation in welfare policy rules, also finding negative effects of child care use (maternal employment).

Next, we study the possibility of heterogeneous effects in two dimensions: mother's education and child's gender. In columns 1 and 2 of Panel A in table 5 we estimate separate 2SLS models for highly (more than highschool) and less educated (highschool or less) mothers. Column 3 shows the difference between estimated coefficients and its standard error. These standard errors were calculated under the assumption of independence and were adjusted to account for differences in sample size. The results suggest that, while children of less educated mothers benefit from a longer maternity leave, children of more educated mothers do not. While the pooled model suggested no impact in the language dimension, we estimate a positive effect of 0.06 standard deviations (significant at 5%) for children of less educated mothers. Panel B explores if there are heterogeneous impacts by child's gender, repeating the previous exercise for this variable. Differences in coefficients are not statiscally significant with the exception of the socioemotional dimension (marginally significant at the 5% level, suggesting a higher gain in socioemotional T-Scores for boys). This result could be consistent with other evidence suggesting higher noncongnitive returns to parental inputs for male children (Bertrand and Pan, 2013). Because of the independence assumption and the reduction in sample sizes, these results should be interpreted with caution.

Finally, we do two robustness checks. In table 6 we show that our 2SLS estimates are robust to the specification of the vector of covariates. In column 1 we only control for child's age (which is essential for the independence assumption to hold). Next we add controls for other child's characteristics (namely child's gender and birthweight). In column 3 we also add mother's characteristics (mother's age and indicators for the mother being a student and her schooling level and civil status). Finally, in column 4, we add other family background control variables (number of people in the household and indicators of father's presence, region and

[19] Note that they study the relation between maternal care and child development in a different framework, yet find similar results to ours. They use the Child Development Supplement data, a complementary survey of the Panel Study of Income Dynamics. It is a longitudinal study of children aged 0 to 12 in the first round. Successive rounds are conducted in five-year intervals. Their sources of exogenous variation in maternal time are shifts to the opportunity cost of time induced by changes in the average cost of childcare, housekeeping costs, average offered wages for the mothers, and welfare programs (subsidies).

urban area), mimicking the specification in column 1 of table 4. The point estimate remains virtually unchanged across columns. In table 7 we replicate our main estimations, setting the cutoff of our binary instrument one year before the cutoff for the actual instrument and excluding truly exposed children. It might be seen as a false experiment in which we expect to find no effects. This is exactly what we find, suggesting that our main results are not driven by a spurious correlation.

Our main results are consistent with the findings of Carneiro, Løken and Salvanes (2015) regarding the long run effects of maternity leave. They study a reform that extended the 12 weeks of unpaid leave to four months of paid and 12 months of unpaid leave in Norway. Their analysis suggests a decline in high school dropout and an increase in wages at age 30, both stronger for children of mothers with low education. The Chilean reform is similar in two important aspects. Children were exposed to an extension of the paid leave during their first six months of life and the main alternative to staying home with their mother was informal care. Figure 1 illustrates this last point. Among mothers who did not stay home between the child's first and third month, 83% used an informal alternative of child care. Among those who did not stay home between the fourth and sixth month, 67% turned to informal childcare arrangements. Similarly, Carneiro, Løken and Salvanes (2015) show that at the time of the Norwegian reform the coverage of formal child care was very low.

Substitution of informal care by prolonged maternal care might have a greater impact than the substitution of formal care. This would be true if the quality of informal care was lower. It is a plausible assumption, as in formal child care centers care is usually provided by trained professionals and children are exposed to greater peer interaction. This argument could also explain the heterogeneity of our results regarding mother's education. On the one hand, less educated mothers could be expected to have lower incomes and thus turn to informal care with a higher probability. On the other hand, we expect education levels to be positively correlated among relatives, so mothers with low education who turn to informal care would be leaving their child under care of less educated people. On the same grounds, we expect less educated mothers who turn to formal care to be more likely to use public childcare centers, which could have a lower quality.[20]

The evidence shown in table 8 below is consistent with the first argument. We show the distribution of child care arrangements for children with $Z_i = 0$ who where not with their

[20]For instance, Noboa and Urzúa (2012) argue that the negative impact of public childcare center attendance on adult interactions that they find "*may be related to the low quality of individual care provided by a limited number of teachers and caregivers at public childcare centers*".

mother during their second three months of life. As can be seen, mothers with low education were 10.8 percentage points less likely to use a formal arrangement. If they tend to rely more on informal arrangements as the table suggests, then their children should benefit more when they are exposed to prolonged maternal care.

Our main result regarding cognitive abilities points in the opposite direction than those of Baker and Milligan (2011). As commented above, they find a negative cognitive effect of the Canadian reform. There is at least one difference that could explain this contrast: the age at which children were exposed to the prolongation of maternal care. The Chilean reform affected children at very young ages, thus, they were not exposed to stranger anxiety and separation anxiety, which could compensate a positive effect of prolonged maternal care.[21] On the other hand, our findings regarding noncognitive skills are in line with Baker and Milligan (2011). Our measure of noncognitive skills is more precise, but we obtain the same result.

C. Mechanisms

So far we have shown that a longer maternity leave causes an increase in total TADI T-scores, driven by increases in the cognitive and motor T-scores. The ELPI data allows us to go further and look for evidence on the mechanisms generating this impact. We focus on the duration of breastfeeding (exclusive or mixed). There is a widely held view of breastfeeding as a promoter of cognitive and motor development. For instance, in a population based birth cohort study Victora et al. (2015) find a positive association between the duration of breastfeeding and IQ, educational attainment and income at age 30. On the same ground, Belfield and Kelly (2012) find a positive association between the incidence of breastfeeding (versus formula feeding) at birth and health status at age 9 months, and nutritional status (less obesity) and cognitive outcomes at age 24 months. They also find that breastfeeding for six months or more is associated with higher motor scores at age 9 months. Breastfeeding can also improve the quality of mother-child interactions. Papp (2013) finds a positive association between breastfeeding and maternal sensitivity. Furthermore, it seems sensible to expect an increase in average breastfeeding duration due to an increase in average maternity leave duration. There is also some empirical evidence on this hypothesis. Baker and Milligan (2008*b*) find an increase in breastfeeding durations induced by the Canadian reform although

[21] This is explained by Baker and Milligan (2011).

they find no impact of breastfeeding on health outcomes.

In this section we explore the effects of the Chilean reform on breastfeeding. The ELPI contains a question on durations which is only answered by mothers who have finished breastfeeding their children. This means that we do not observe durations for mothers of children who are still being breastfed. But, under the assumption of a continuous duration of breastfeeding, if we know that child i is still being breastfed at age a_i, we also know that breastfeeding duration will be at least a_i.[22] Thus, imputing child's age as the duration of breastfeeding in these cases we turn a truncation problem into a censoring problem. This strategy allows the use of nonparametric and semiparametric tools of survival analysis which take account of (independent) censoring.

We start our analysis with a comparison of the survivor functions of breastfeeding for children with $Z_i \in (0,1]$, i.e. eligible children, and $Z_i = 0$, i.e. non-eligible children.[23] In figure 2 we plot the Kaplan-Meier estimate of the survivor function for both groups. This nonparametric estimator, also known as the product-limit estimator, is consistent under independent censoring (Kaplan and Meier, 1958). We plot the estimates for durations $t \leq 20$ because exposed children are younger than non-exposed children and there is no data for this group beyond a duration of 20 months. The survivor functions seem to be equal until a duration of 3 months. From this point on the probability of survival is greater at any duration for the $Z_i \in (0,1]$ group. The difference is more pronounced from $t = 6$ onwards. This evidence suggests that the reform induced longer breastfeeding durations.

Next, we introduce covariates. This requires adding some structure. We estimate a Cox proportional hazards model. This semiparametric estimator consistently estimates a proportional hazards function with independent censoring (Cox, 1972). The hazard rate for individual i is modeled as

$$h(t|Z_i, \mathbf{X}_i) = h_0(t) \exp\left(\phi Z_i + \mathbf{X}_i'\omega\right) \tag{4}$$

where Z_i is our instrument, $\mathbf{X}_i$ is a vector of covariates and no assumption is made about the functional form of the baseline hazard $h_0(t)$.[24] Table 9 shows the results of estimating

[22] This is a reasonable assumption. Once breastfeeding is stopped, milk production is ceased so it is at least difficult to resume the process. In other words, breastfeeding should be a continuous process.

[23] The survivor function, defined as $S(t) \equiv Pr\,(T > t)$, is the probability of survival past time t. By survival we mean remaining in the "breastfeeding" state. It can be equivalently seen as the probability of failing after time t, where by failing we mean getting out of the "breastfeeding" state.

[24] The hazard rate, defined as $h(t) \equiv \lim\limits_{\Delta t \to 0} \frac{Pr(t+\Delta t\, >\, T\, >\, t\ |\ T>t)}{\Delta t}$, is the instantaneous probability of leaving the

this model under four different specifications of the vector $\mathbf{X}_i$. The estimated hazard ratio is 0.82, meaning that the hazard rate for exposed children is 0.82 times that of non-exposed children. We reject the null hypothesis $H_0 : \phi = 0$ at the 10% significance level. This test is asymptotically equivalent to $H_0 : \exp(\phi) = 1$, but is preferred for its better small-sample properties.[25] We show similar results using a binary instrument in table B.3 in the appendix. This evidence points in the same direction as our previous analysis with the Kaplan-Meier estimator, implying that the Chilean reform caused an increase in breastfeeding durations. In addition to the biological channel, an increase in breastfeeding duration necessarily implies an increase in time spent with the child. This result agrees with those of Baker and Milligan (2008*b*) for Canada and Berger, Hill and Waldfogel (2005) for the US. This mechanism could partially explain the positive cognitive and motor effects of maternity leave found in the previous section.

D. Mother Outcomes

Now we turn to mother's outcomes. Table 10 shows the 2SLS results for the parental stress index and last wage reported, and the IV-Probit result for employment. The dependent variable in Column (1) is the PSI percentile score. We find that a longer leave reduces maternal stress according to this measure. This implies that a longer maternity leave might have a positive mental health effect for the mother. This effect has the expected sign: entrusting child care to other people during the early stages might cause some stress to the mother. Column (2) shows the estimates for the logarithm of the last wage reported by the mother. We find no evidence of a wage effect of maternity leave. Finally, the dependent variable in Column (3) is an indicator for the mother being employed at the time of the survey. The point estimate is positive and significant at 10%, indicating an increase in the probability of being employed of 1.4 percentage points.

Our mother employment result is consistent with the findings of Ruhm (1998) regarding women's employment, although he finds a negative impact in their relative wages. In contrast, Lalive and Zweimüller (2009) find that the Austrian reform induced a short run decrease in women's employment and earnings, but produced no long run effects. An increase in

"breastfeeding" state conditional on survival to time t. The specification of a proportional hazards model implies that, for two individuals $i \neq j$, the ratio of the hazard rates is a constant: $\frac{h(t|Z_i,\mathbf{X}_i)}{h(t|Z_j,\mathbf{X}_j)} = \frac{\exp\left(\phi Z_i + \mathbf{X}_i'\omega\right)}{\exp\left(\phi Z_j + \mathbf{X}_j'\omega\right)}$. In other words, one individual's hazard rate is a multiple of another's.

[25] The exponentiated coefficient is the hazard ratio for a one unit change in the regressor. Particularly, in the case of our instrument, we have $\frac{h(t|Z_i=1,\mathbf{X}_i=\mathbf{X})}{h(t|Z_i=0,\mathbf{X}_i=\mathbf{X})} = \frac{h_0(t)\exp\left(\phi+\mathbf{X}'\omega\right)}{h_0(t)\exp(\mathbf{X}'\omega)} = \exp(\phi)$.

mother's employment could be explained by job continuity. The possibility to return to her pre-childbirth job (assured by job protection) allows the mother to continue accumulating specific human capital, thus reducing her incentives to quit the labor market. Baker and Milligan (2008*a*) find evidence that maternity leave entitlements increase job continuity with the prebirth employer in Canada.[26]

VI. Conclusions

This study estimated the causal effect of the length of maternity leave on children's abilities and mother outcomes. Our identification strategy exploits the exogenous variation induced by a reform that extended the legal maternity leave and affected differently children born before and after its entry into force. Using a rich data set that includes cognitive and noncognitive ability tests we assess the effects of maternity leave on total, cognitive, language, motor, and socioemotional ability as measured by TADI T-scores.

Our main findings indicate that the short term effects of maternity leave are positive and significant for total abilities for the children of mothers who extended their leave induced by the reform. The effects are driven by cognitive and motor skills and are stronger for children whose mothers are less educated. We find no effects of exposure to reform on socioemotional skills. We hypothesize as a potential channel the substitution of informal care by time spent with mother. Last, we identify breastfeeding as another channel through which the effects manifest. We estimate duration models for breastfeeding length finding positive and significant effects for those exposed to the reform compared to those not exposed.

When looking at mother outcomes, we find a reduction of mother's stress (as measured by the Parental Stress Index) and a slightly positive effect on the probability of being employed but no effects on wages.

Our findings complement the mixed evidence reported in the literature in two ways. First, it is one of the first investigations using maternity leave reforms that affected children at very young ages (between 3 and 6 months old). Second, it is one of the first evaluations for a developing country which contributes to the understanding of human capital formation in a different context to that of developed countries.

[26]In Chile, Perticara and Sanhueza (2010) find that women face about 40% higher risk of leaving employment when the child is three month old but it disappears after the child is one year old. Our result is different from theirs. We find that the longer the maternity leave the more probability of being employed at the time of the survey, which is at least 7 months after birth.

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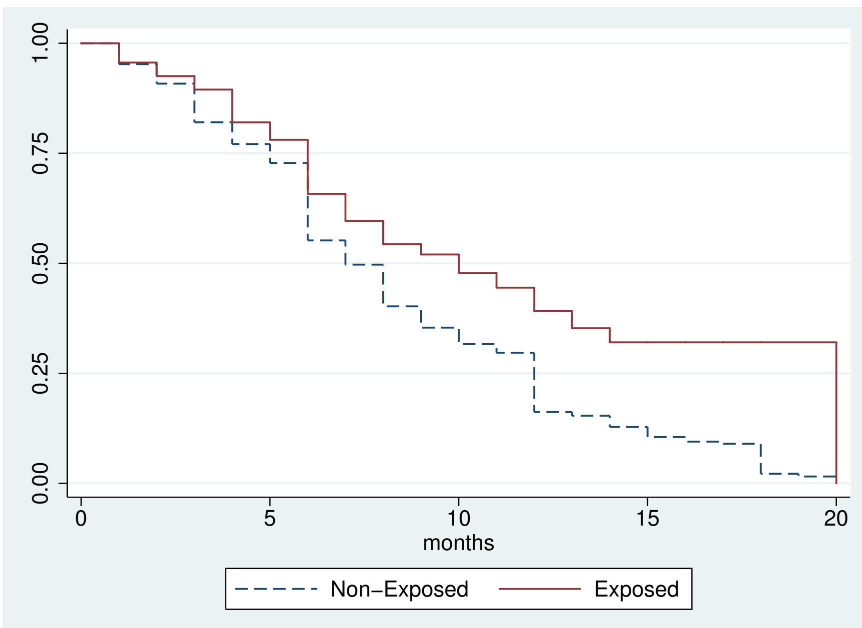
Figure 1. Child care arrangements

Figure 2. Kaplan-Meier estimates of the survivor function



Note: Kaplan-Meier estimates of the survivor function for children who where exposed ($Z > 0$) and not exposed ($Z = 0$) to the reform.

Table 1—Descriptive Statistics

	Obs	Mean	Std. Dev.	Min	Max
Maternity Leave Duration					
Exposed	237	16.966	8.065	0	48
Non-exposed	3,221	13.477	8.056	0	54
Response Variables					
Maternity leave duration (weeks)	3,458	13.716	8.104	0	54
TADI, total	3,387	52.647	8.672	23	81
TADI, cognitive	3,402	52.827	11.071	23	81
TADI, language	3,405	52.061	10.997	23	81
TADI, motor	3,401	53.229	11.919	23	81
TADI, socioemotional	3,406	52.419	11.646	23	81
Used SCL, 0-3 months	982	.202	.401	0	1
Used SCL, 3-6 months	1,739	.254	.435	0	1
Used SCL, 0-3 months	2,469	.258	.437	0	1
Parental Stress Index (PSI)	2,894	42.005	33.546	1	99
Mother employed	3,450	.79	.407	0	1
Mother's wage (CLP, thousands)	3,026	329.89	287.04	0	4,000
Child's Characteristics					
Age (months)	3,458	45.749	16.645	7	72
Gender (male)	3,458	.509	.5	0	1
Birthweight (kg.)	3,233	3.36	.479	2	4.98
Family Background					
Mother's age (years)	3,450	32.999	6.155	16	71
Mother was a worker at childbirth	3,453	.912	.284	0	1
Mother had an employment contract at childbirth	3,040	.959	.198	0	1
Mother is married	3,450	.424	.494	0	1
Mother is in a domestic partnership	3,450	.318	.466	0	1
Mother is a widow	3,450	.005	.068	0	1
Mother is single	3,450	.191	.393	0	1
Mother is separated	3,450	.062	.241	0	1
Mother is a student	3,450	.053	.224	0	1
Mother with low education	3,442	.669	.471	0	1
Mother with high education	3,442	.331	.471	0	1
Father is present	3,458	.723	.447	0	1
Father (whenever present) employed	2,501	.961	.194	0	1
Home is located in urban area	3,458	.938	.241	0	1
Number of people in the household	3,458	4.545	1.517	2	23

Notes: TADI T-scores shown in this table are not standardized. SCL stands for sick child leave

Table 2—Child Development First Stage

	Maternity Leave Duration (weeks)		
	Reported leave (1)	Reported + inactivity (2)	Reported + Sick child leave (3)
Child's Exposition	4.1710***	3.7361***	3.2153***
	(0.718)	(0.974)	(0.816)
Child's age (months)	-0.0006	0.0090	-0.0151
	(0.010)	(0.014)	(0.011)
Mother's age (years)	0.0399	-0.0769**	0.0210
	(0.025)	(0.034)	(0.029)
Child's gender (male)	-0.5509**	-0.5485	-0.5233
	(0.280)	(0.378)	(0.318)
Birthweight (kg.)	-0.2823	-0.5958	-0.1733
	(0.294)	(0.397)	(0.335)
Urban area	-0.0545	-0.5797	0.1386
	(0.619)	(0.828)	(0.704)
Father's presence	0.2366	0.2667	0.8202
	(1.055)	(1.409)	(1.199)
# of people in the household	0.1382	0.2117*	0.1203
	(0.093)	(0.125)	(0.106)
Student mother	-1.4494**	-0.9542	-1.2872*
	(0.686)	(0.922)	(0.780)
Constant	10.1148***	20.1742***	10.5026**
	(3.669)	(4.746)	(4.172)
Observations	3151	3242	3151
F-stat (excluded instrument)	33.72	14.70	15.52

Notes: Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Excluded instrument is child's exposition. It is equal to the number of additional weeks of leave allowed due to the reform divided by the number of weeks in regime (so it varies form 0 to 1). Other control variables not shown are region and mother's education level and civil status dummies.

Table 3—Child Development OLS

	TADI				
	Total	Cognitive	Language	Motor	Socioemotional
Maternity leave duration (weeks)	0.0017	-0.0011	0.0024	0.0048*	0.0020
	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)
Child's age (months)	0.0051***	0.0062***	0.0017	0.0099***	0.0028**
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Mother's age (years)	-0.0025	-0.0007	-0.0024	-0.0072*	-0.0008
	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)
Child's gender (male)	-0.1749***	-0.1035***	-0.1864***	-0.1745***	-0.2401***
	(0.030)	(0.039)	(0.038)	(0.042)	(0.041)
Birthweight (kg.)	0.0311	0.0411	-0.0195	0.0986**	0.0149
	(0.031)	(0.040)	(0.040)	(0.044)	(0.043)
Urban area	0.0724	0.1879**	0.1113	-0.0586	0.0715
	(0.066)	(0.080)	(0.084)	(0.094)	(0.097)
Father's presence	-0.2320**	-0.1934	-0.3110**	-0.2267	-0.1928
	(0.108)	(0.146)	(0.146)	(0.160)	(0.147)
Number of people in the household	-0.0318***	-0.0276**	-0.0345***	-0.0293**	-0.0345**
	(0.010)	(0.013)	(0.012)	(0.014)	(0.014)
Student mother	0.0895	0.0400	0.0412	0.0413	0.2287**
	(0.072)	(0.097)	(0.091)	(0.100)	(0.103)
Constant	0.1136	0.0734	0.1047	0.2959	-0.0632
	(0.403)	(0.587)	(0.494)	(0.491)	(0.487)
Observations	3,151	3,166	3,169	3,165	3,170

Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Other control variables not shown are region and mother's education level and civil status dummies.

Table 4—Child Development 2SLS

	TADI				
	Total	Cognitive	Language	Motor	Socioemotional
Maternity leave duration (weeks)	0.0392**	0.0691***	0.0341	0.0611***	-0.0036
	(0.017)	(0.025)	(0.024)	(0.023)	(0.022)
Child's age (months)	0.0063***	0.0083***	0.0027*	0.0116***	0.0026*
	(0.001)	(0.002)	(0.001)	(0.002)	(0.002)
Mother's age (years)	-0.0040	-0.0035	-0.0036	-0.0093**	-0.0006
	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)
Child's gender (male)	-0.1525***	-0.0640	-0.1674***	-0.1401***	-0.2436***
	(0.033)	(0.046)	(0.042)	(0.047)	(0.043)
Birthweight (kg.)	0.0418	0.0610	-0.0115	0.1131**	0.0134
	(0.033)	(0.046)	(0.042)	(0.047)	(0.043)
Urban area	0.0733	0.1896**	0.1094	-0.0621	0.0718
	(0.070)	(0.091)	(0.086)	(0.100)	(0.096)
Father's presence	-0.2406**	-0.2098	-0.3187**	-0.2402	-0.1914
	(0.113)	(0.166)	(0.144)	(0.170)	(0.146)
Number of people in the household	-0.0372***	-0.0371***	-0.0390***	-0.0374**	-0.0336**
	(0.010)	(0.014)	(0.013)	(0.015)	(0.014)
Student mother	0.1496*	0.1540	0.0921	0.1322	0.2195**
	(0.079)	(0.114)	(0.100)	(0.115)	(0.109)
Constant	-0.3348	-0.7645	-0.2699	-0.3720	0.0035
	(0.446)	(0.694)	(0.531)	(0.620)	(0.560)
Observations	3,151	3,166	3,169	3,165	3,170

Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Other control variables not shown are region and mother's education level and civil status dummies.

Table 5—Child Development Heterogeneity

Panel A: Mother's Education

	(1) High school or less	(2) More than high school	(1)−(2) Difference
TADI total	0.0621***	0.0112	0.0509**
	(0.023)	(0.030)	(0.026)
	[2,105]	[1,046]	{1.99}
TADI cognitive	0.1001***	0.0297	0.0704*
	(0.035)	(0.043)	(0.038)
	[2,115]	[1,051]	{1.86}
TADI language	0.0643**	-0.0070	0.0713*
	(0.030)	(0.048)	(0.037)
	[2,118]	[1,051]	{1.93}
TADI motor	0.0831***	0.0325	0.0414
	(0.031)	(0.043)	(0.035)
	[2,116]	[1,049]	{1.17}
TADI socioemotional	0.0034	-0.0053	-0.0087
	(0.026)	(0.044)	(0.033)
	[2,119]	[1,051]	{-0.26}

Panel B: Child's Gender

	(1) Female child	(2) Male child	(1)−(2) Difference
TADI total	0.0284	0.0597*	-0.0313
	(0.019)	(0.034)	(0.028)
	[1,552]	[1,599]	{-1.13}
TADI cognitive	0.0589**	0.0821*	-0.0232
	(0.028)	(0.048)	(0.039)
	[1,559]	[1,607]	{-0.59}
TADI language	0.0428	0.0224	0.0204
	(0.027)	(0.047)	(0.038)
	[1,561]	[1,608]	{0.53}
TADI motor	0.0443*	0.0999**	-0.0556
	(0.027)	(0.049)	(0.040)
	[1,557]	[1,608]	{-1.4}
TADI socioemotional	-0.0300	0.0425	-0.0725**
	(0.028)	(0.044)	(0.037)
	[1,560]	[1,610]	{-1.96}

Notes: Robust standard errors in parentheses, sample sizes in square brackets and t-statistics in braces. *** p<0.01, ** p<0.05, * p<0.1. Standard errors for the differences of coefficients were calculated assuming independence and an adjustment for differences in sample size was applied. Control variables not shown are child's and mother's age, birthweight, number of people in the household, and region, urban area, father's presence, and mother's civil status dummies. In panel A we also control for child's gender. In panel B we also control for mother's education level

Table 6—Different specifications of covariate vector

	TADI (total)	TADI (total)	TADI (total)	TADI (total)
Maternity leave duration (weeks)	0.0340**	0.0350**	0.0358**	0.0392**
	(0.016)	(0.016)	(0.017)	(0.017)
Child's characteristics	No	Yes	Yes	Yes
Mother's characteristics	No	No	Yes	Yes
Other family background variables	No	No	No	Yes
Observations	3,387	3,166	3,151	3,151

Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. All specifications control for child's age.

Table 7—False instrument

	TADI				
	Total	Cognitive	Language	Motor	Socioemotional
Maternity leave duration (weeks)	-0.2427	-0.0933	-0.1479	-0.4545	-0.2642
	(0.202)	(0.115)	(0.146)	(0.364)	(0.223)
Child's characteristics	Yes	Yes	Yes	Yes	Yes
Mother's characteristics	Yes	Yes	Yes	Yes	Yes
Other family background variables	Yes	Yes	Yes	Yes	Yes
Observations	2,924	2,939	2,940	2,936	2,941

Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 8—Child care arrangement for non-exposed children between 3 and 6 months

Child care arrangement	More than high school	High school or less	Diff.
Formal care	0.406	0.298	0.108***
			(0.0276)
Informal care	0.594	0.702	

Notes: Third column: Standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 9—Breastfeeding Duration

	(1)	(2)	(3)	(4)
Child's exposition to the reform	0.8179*	0.8173*	0.8177*	0.8278*
	(0.091)	(0.091)	(0.091)	(0.093)
Per capita household income (last month)	1.0000***	1.0000***	1.0000***	1.0000***
	(0.000)	(0.000)	(0.000)	(0.000)
Mother's age (years)	0.9969	0.9991	0.9994	0.9998
	(0.003)	(0.003)	(0.003)	(0.003)
Birthweight (kg.)	0.8634***	0.8652***	0.8635***	0.8674***
	(0.032)	(0.032)	(0.032)	(0.033)
Father's presence		1.0261	1.0283	0.9189
		(0.043)	(0.044)	(0.125)
Inactive mother		0.9484	0.9348	0.9458
		(0.044)	(0.043)	(0.044)
Number of children of the mother		0.9598**	0.9614*	0.9814
		(0.020)	(0.020)	(0.021)
Student mother			1.1899**	1.1924**
			(0.095)	(0.104)
Dummies for risky behaviors during pregnancy	Yes	Yes	Yes	Yes
Dummies for child care arrangement during 0-3 months	Yes	Yes	Yes	Yes
Region fixed effects	No	No	Yes	Yes
Dummies for mother's civil status and schooling level	No	No	No	Yes
Observations	2,937	2,936	2,936	2,931

Notes: Hazard ratios (exponentiated coefficients) reported. Delta-method standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 10—PSI and Labor Market Outcomes 2SLS

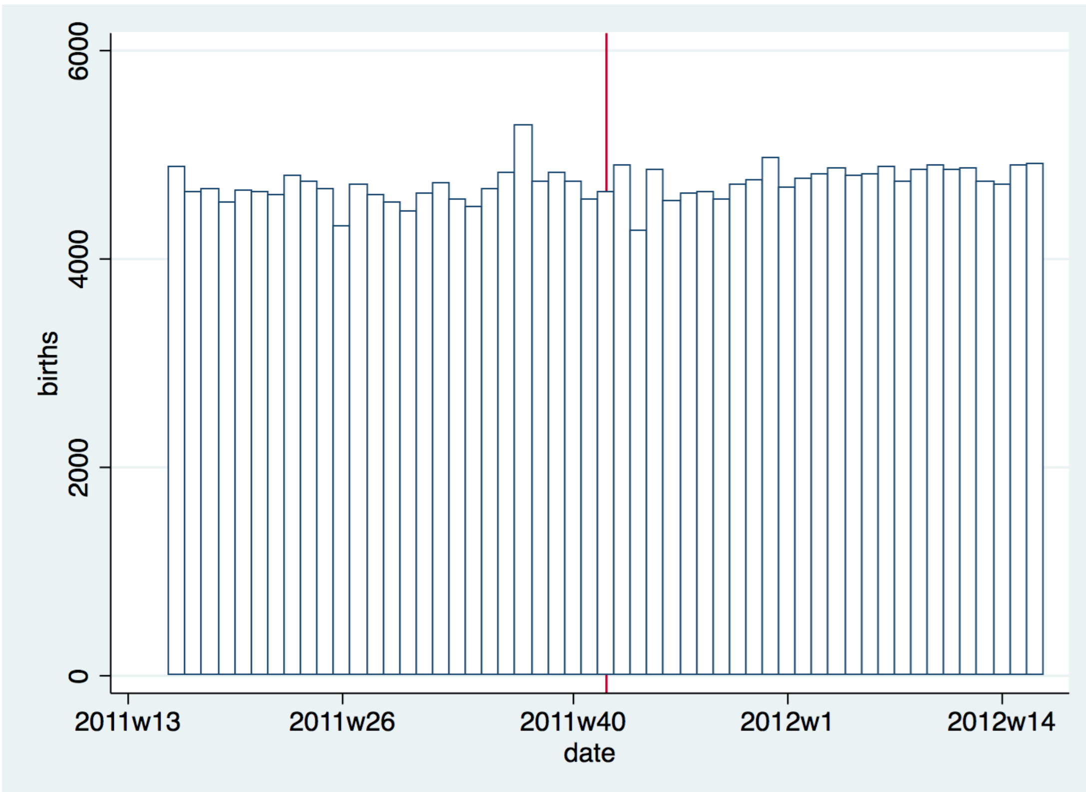
	PSI	Log Wage	Employment
Maternity leave duration (weeks)	-2.8913***	-0.0027	0.0492*
	(1.038)	(0.013)	(0.029)
Child's age (months)	0.0767	-0.0008	0.0019
	(0.049)	(0.001)	(0.002)
Mother's age (years)	-0.0539	0.0125***	0.0202***
	(0.146)	(0.002)	(0.005)
Child's gender (male)	1.8812	-0.0271	-0.0441
	(1.537)	(0.021)	(0.052)
Urban area	-0.8848	-0.0349	0.1620*
	(3.228)	(0.039)	(0.096)
Father's presence	-6.2858***	0.0123	0.0387
	(1.697)	(0.075)	(0.165)
Number of people in the household	0.3720	-0.0188***	-0.0046
	(0.484)	(0.007)	(0.016)
Student mother	-10.6429***	-0.0044	-0.0393
	(3.970)	(0.059)	(0.129)
Constant	95.7993***	12.0375***	-2.3351***
	(21.164)	(0.526)	(0.618)
Observations	2,882	3,011	3,437

Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Estimation method is 2SLS in columns 1 and 2, and IV-Probit in column 3. Common control variables not shown are region and mother's education dummies. Mother's civil status dummies not shown in columns 2 and 3. Month and year dummies not sown in column 2. The implied marginal effect of an additional week of maternity leave in column 3 is 0.0138

Appendix

A. Weekly Births

Figure A1. Weekly births before and after enactment



Note: Weekly births before and after the enactment of the law. Vertical red line is the enactment date October 17, 2011 (week 42).

B. Binary instrument

Table B.1—Child Development First Stage, Binary Instrument

	Maternity Leave Duration (weeks)	
	FT and PT leave (1)	FT leave only (2)
Child's Exposition	3.5445***	2.9184***
	(0.644)	(0.639)
Child's age (months)	-0.0005	-0.0022
	(0.010)	(0.010)
Mother's age (years)	0.0392	0.0440*
	(0.025)	(0.025)
Child's gender (male)	-0.5574**	-0.4528
	(0.280)	(0.278)
Birthweight (kg.)	-0.2779	-0.3805
	(0.295)	(0.293)
Urban area	-0.0297	0.0026
	(0.619)	(0.615)
Father's presence	0.2428	0.2834
	(1.056)	(1.048)
# of people in the household	0.1407	0.1138
	(0.093)	(0.093)
Student mother	-1.4198**	-1.2766*
	(0.687)	(0.682)
Constant	10.1236***	10.7189***
	(3.674)	(3.648)
Observations	3,151	3,151
F-stat (excluded instrument)	30.34	20.87

Notes: Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Excluded instrument is child's exposition. In its binary version it is equal to one if the child is exposed to any extension whatsoever and zero otherwise. Other control variables not shown are region and mother's education level and civil status dummies.

Table B.2—Child Development 2SLS, Binary Instrument

	TADI				
	Total	Cognitive	Language	Motor	Socioemotional
Maternity leave duration (weeks)	0.0573***	0.0871***	0.0425*	0.0895***	0.0153
	(0.019)	(0.029)	(0.025)	(0.027)	(0.024)
Child's age (months)	0.0068***	0.0089***	0.0029*	0.0125***	0.0032**
	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)
Mother's age (years)	-0.0047	-0.0042	-0.0039	-0.0104**	-0.0014
	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)
Child's gender (male)	-0.1417***	-0.0539	-0.1624***	-0.1227**	-0.2319***
	(0.035)	(0.049)	(0.042)	(0.051)	(0.043)
Birthweight (kg.)	0.0470	0.0661	-0.0094	0.1204**	0.0185
	(0.036)	(0.049)	(0.042)	(0.051)	(0.043)
Urban area	0.0737	0.1900**	0.1090	-0.0638	0.0707
	(0.075)	(0.097)	(0.088)	(0.107)	(0.096)
Father's presence	-0.2447**	-0.2139	-0.3208**	-0.2470	-0.1961
	(0.119)	(0.174)	(0.144)	(0.180)	(0.148)
Number of people in the household	-0.0398***	-0.0395***	-0.0402***	-0.0415***	-0.0364**
	(0.011)	(0.015)	(0.013)	(0.016)	(0.014)
Student mother	0.1784**	0.1833	0.1055	0.1782	0.2504**
	(0.085)	(0.121)	(0.102)	(0.126)	(0.109)
Constant	-0.5502	-0.9795	-0.3686	-0.7099	-0.2212
	(0.469)	(0.732)	(0.534)	(0.684)	(0.555)
Observations	3,151	3,166	3,169	3,165	3,170

Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Other control variables not shown are region and mother's education level and civil status dummies.

Table B.3—Breastfeeding Duration, Binary Instrument

	(1)	(2)	(3)	(4)
Child's exposition to the reform	0.8150**	0.8137**	0.8148**	0.8208**
	(0.078)	(0.078)	(0.078)	(0.080)
Per capita household income (last month)	1.0000***	1.0000***	1.0000***	1.0000***
	(0.000)	(0.000)	(0.000)	(0.000)
Mother's age (years)	0.9971	0.9996	0.9998	1.0002
	(0.003)	(0.003)	(0.003)	(0.003)
Birthweight (kg.)	0.8619***	0.8639***	0.8621***	0.8662***
	(0.032)	(0.032)	(0.032)	(0.032)
Father's presence		1.0210	1.0226	0.9169
		(0.043)	(0.043)	(0.124)
Inactive mother		0.9504	0.9376	0.9490
		(0.044)	(0.043)	(0.045)
Number of children of the mother		0.9566**	0.9582**	0.9778
		(0.020)	(0.020)	(0.021)
Student mother			1.1818**	1.1796*
			(0.093)	(0.101)
Dummies for risky behaviors during pregnancy	Yes	Yes	Yes	Yes
Dummies for child care arrangement during 0-3 months	Yes	Yes	Yes	Yes
Region fixed effects	No	No	Yes	Yes
Dummies for mother's civil status and schooling level	No	No	No	Yes
Observations	2,951	2,950	2,950	2,945

Notes: Hazard ratios (exponentiated coefficients) reported. Delta-method standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.