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ABSTRACT

The Cost of Following Traditional Gender Norms: Evidence from a Paid Leave for Seriously Ill Children*

In this paper we exploit the introduction in February 2018 of a new paid parental leave program to care for a seriously ill child in Chile (SANNA) to identify the role of both economic incentives and gender norms on families' decisions regarding market versus home production specialization. To measure the impact of economic incentives, we utilize the design of the SANNA program, which covers the beneficiary's wages up to a specific threshold, beyond which the benefit remains fixed. The efficient allocation of this benefit depends on the income levels of family members and whether their income exceeds the threshold. To investigate the role of gender norms, we compare the effect of economic incentives among older, more traditional families and younger families. Our results indicate that both gender norms and economic incentives affect parental leave allocation. We estimate that older families pay a cost of USD 1,200 for adhering to traditional gender norms compared to younger families.

JEL Classification: D13, J22, J16

Keywords: gender identity norms, parental childcare, gender wage gap

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1 Introduction

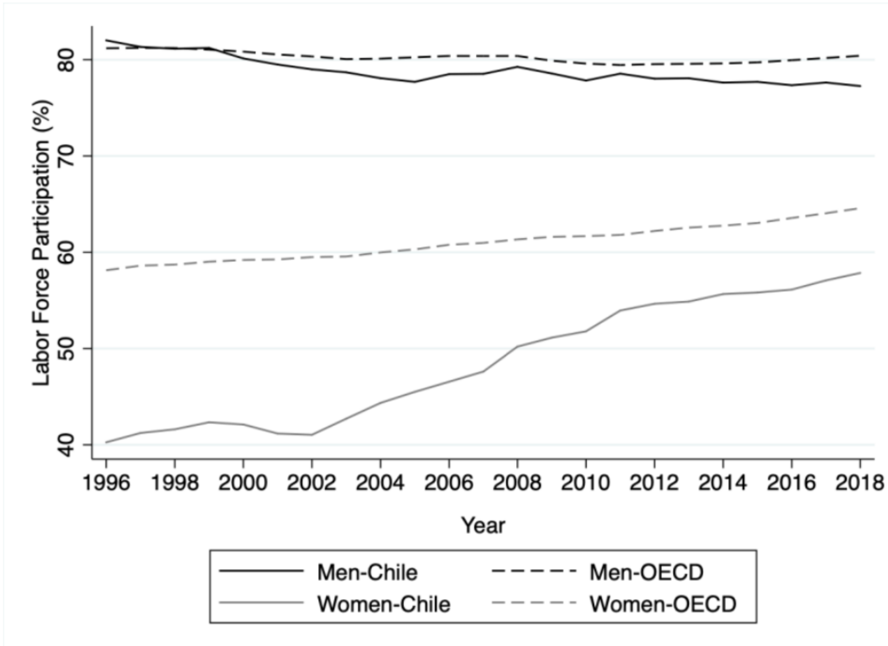
Large gender gaps persist in terms of time spent in domestic and market work in Latin American countries, in particular, Chile. Even when female labor participation has been rising (between 1996 and 2018 it increased to about 58%), it remains low in comparison to other countries from the Organisation for Economic Cooperation and Development (OECD). In turn, male labor participation in Chile has remained relatively stable since 1996 at about 80%, which is close to the OECD average (Figure 1). A similar pattern emerges when we compare Chile to other Latin American countries.[1] Standard factors such as age, education, number of children, and marital status, are not enough to explain the relatively small group of Chilean women who decide to participate in the workforce (Contreras and Plaza, 2010). Most of these standard models fail to consider other variables such as cultural factors and gender identity (Akerlof and Kranton, 2000). One such factor is the prevalence of traditional gender norms, deeply rooted and transmitted over time across generations (Alesina et al., 2013).

The role of gender identity norms on women's labor market participation in Latin American countries has received little attention in the literature due to the lack of adequate data for examining this hypothesis. A notable exception is Contreras and Plaza (2010), which uses survey data to directly control for beliefs about gender roles. The aim of this paper is to fill this gap and identify the role of both economic incentives and gender norms on families' decisions to specialize in market versus home production. We exploit the introduction in February 2018 of a new paid parental leave to care for a seriously ill child in Chile (from here referred to by its Spanish acronym, "SANNA").[2] SANNA gives each parent of seriously ill children between 45 and 90 days of paid leave per year depending on the type of illness, and each parent can transfer to the other a fraction or all the days of the paid leave that their are entitled to. To be eligible for these benefits, parents must provide a medical certificate for

[1]Chile ranks 99 out of 153 countries in terms of women labor force participation according to the Global Gender Gap Index 2020 published by the World Economic Forum.

[2]Law No. 21,063 called "Seguro para el Acompañamiento de Niñas y Niños" (SANNA).

Figure 1: Labor force participation by sex (ages 15-64, percentage)



Source: OECD.

the seriously ill child, and meet certain requirements in terms of accumulating a minimum number of contributions to the social security system at the time of the diagnosis. They can be either employed, self-employed or even unemployed.

Our study examines the allocation of parental leave between fathers and mothers in families, with a particular focus on the role of social norms. Specifically, we anticipate that prevailing societal norms may lead to a higher likelihood of mothers taking advantage of parental leave compared to fathers. Additionally, we explore the economic costs associated with conforming to traditional gender roles in parental leave allocation. To achieve this, we leverage the structure of the SANNA program, which provides coverage for the wage of the beneficiary up to a pre-determined threshold, with any excess being fixed. Therefore, when one parent's income exceeds the threshold and the other's falls below it, it may be economically advantageous to transfer the benefit to the lower-earning parent. We hypothesize that mothers may be less inclined to transfer this benefit to fathers compared to parents.

Using a difference-in-differences approach, we examine whether families with one parent who earns above the parental leave threshold are more likely to allocate the benefit to the lower-earning parent and whether this likelihood varies based on the gender of the lower-earning parent. Our results show that economic incentives play a significant role in the allocation of parental leave within families. For families where the mother has a higher (past) wage compared to the father, and her wage is above the threshold and thus the efficient allocation of days indicates that the father should take the whole leave, the share of days taken by the mother is 46-50 p.p. lower. This implies that when families are deciding who takes care of a seriously ill child, economic incentives are still important.

We also provide estimates of the cost of following traditional gender norms. To achieve this, we compare younger to older generations of parents. We argue that the former are less likely to adhere to traditional gender norms in parental leave allocation. Indeed, data from a survey conducted in Chile in 2017 shows that younger cohorts adhere less to traditional gender roles towards mothers' paid work and the division of labor (Encuesta Bicentenario UC).[3] Our results suggest that the impact of gender norms differs between younger and older families. For younger families, the coefficient for the interaction between the mother having the highest past income and the mother's past income is 57 to 68 percent of that for older families. We interpret this as traditional gender norms moderating the effect of economic incentives. Although we lack sufficient power to establish statistical significance for this difference at conventional levels, our findings suggest that older families may incur costs for adhering to traditional gender norms.

The rest of the document is organized as follow. We provide a background section describing the SANNA program in detail, to then analyze the current literature and establish our contribution. We then describe the data, present our empirical strategy and the results. In the final section we conclude.

[3] Figure A.1 in Appendix A shows the percentage of the population that agrees with each statement by age category.

2 The SANNA program

In February 2018, Chile introduced a new paid parental leave program called SANNA to support parents who need to care for a seriously ill child between the ages of one and eighteen. Under this program, each parent is entitled to a specific number of days of paid leave on a "per episode" and per year basis, which means that they can take a certain number of days each time their child requires care, with a limit on the total number of days they can take in a year. The program covers serious health conditions such as cancer, solid organ and hematopoietic stem cell transplants, end-stage or terminal illness, and serious accidents with a risk of death or serious and permanent functional sequela. Table 1 shows the number of days of entitlement, depending on the specific health condition, with cancer and end-stage or terminal illness providing the largest number of days for parents to take time off work.[4]

Table 1: Summary of days of paid leave provided by SANNA program

	Days of paid leave (per episode per year)
Cancer	90 days (up to 2 years)
Solid organ and hematopoietic stem cell transplants	90 days
End-stage or terminal illness	No limit, determined on a case-by-case basis
Serious accident with risk of death or serious and permanent functional sequela	45 days

Note: The table shows the number of days each parent is entitled to 'per episode' and per year according to Law No. 21,063 called "Seguro para el Acompañamiento de Niñas y Niños" (SANNA). In the case of cancer, parents are entitled to 90 days per year for up to 2 years.

The SANNA program offers the flexibility for each parent to transfer some or all of their entitled days of paid leave to the other parent. For example, if a family has a child with cancer, the father can transfer all of his days to the mother, allowing her to take up to 360

[4]The law imposes a maximum number of days per leave permit for the proper use of the program, which means that parents may need to provide multiple medical certificates to use all their available days in a given year. For example, a mother that decides to make use of all her available days in a given year (90 days) in the case of a sick child with cancer, she has to provide 6 medical certificates for 15 days each.

days off work over a two-year period. In the event of a serious accident, parents can transfer a maximum of 2/3 of their entitled days, meaning that 1/3 of the paid leave is reserved for the mother or father as a non-transferable right that must be used within a certain period or it will be forfeited ("use it or lose it").

During the first two years of the Covid-19 pandemic the government exceptionally increased the number of days provided by SANNA program. In 2020 and 2021 by decree the government set out four extensions: i) for 90 days on June 17th, 2020; ii) for 30 days on November 27th, 2020; iii) for 30 days on April 5th, 2021, and iv) for 90 days on July 12th, 2021.[5] To access these extra days of leave provided by the SANNA program, one of the parents must commit to using all of the days of any of the extensions, and these days cannot be transferred between parents. To qualify for the first extension, both parents must have already used all of the days originally assigned to them by the program. Likewise, to qualify for the second, third, and fourth extensions, both parents must have used all of the days assigned to them in the preceding extension.

To qualify for the SANNA benefits, parents must provide a medical certificate indicating the serious illness of their child. Both employed and self-employed individuals are eligible, provided they have made at least 8 and 12 contributions, respectively, to the social security system within the past 24 months. Furthermore, for employed individuals, the last 3 contributions must be consecutive and from the same employer. Unemployed parents are also eligible if they have been affiliated with the social security system for at least 12 months, and have made at least 8 contributions as an employee within the last 24 months.

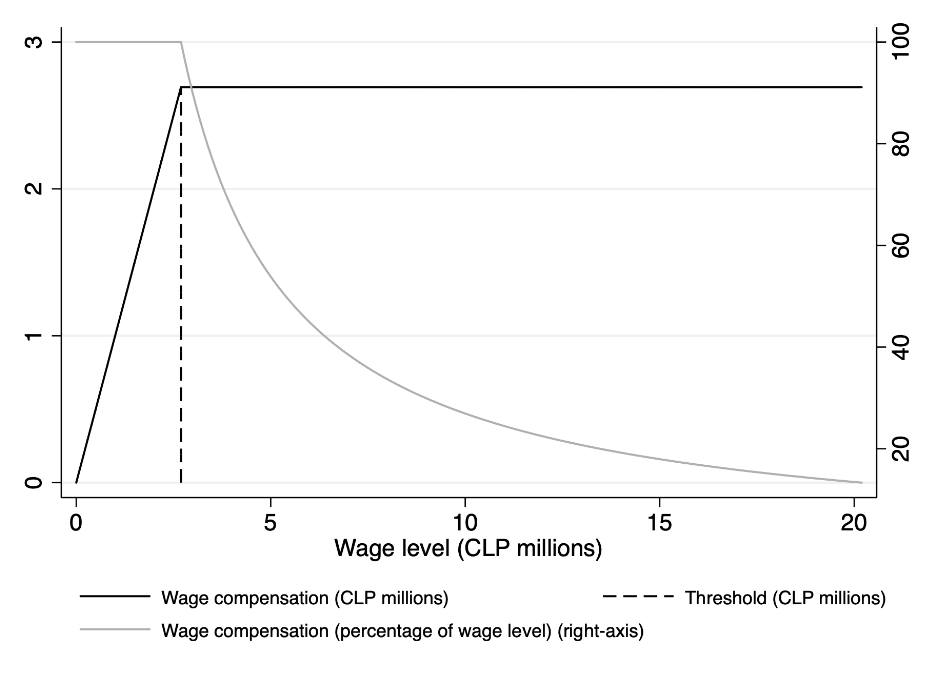
The SANNA program provides paid parental leave up to a certain amount that is determined annually by law.[6] In other words, if a person's wages fall below the threshold, they will receive full compensation for their lost wages, but if their wages exceed the threshold, they will only receive compensation up to this fixed level (Figure 2). This alters the economic

[5]Parents will be able to access to these extra days as long as the pandemic has not been declared "officially" over by the Chilean government.

[6]As reference for year 2022 this level was set in 81.6 UF equivalent to USD 3,300, approximately.

cost of taking the leave and thus can change the optional allocation of the leave within a family.

Figure 2: Wage compensation in SANNA program



There are several reasons why Chile is an interesting setting to investigate the impact of gender norms. Firstly, the implementation of SANNA in February 2018 provides new administrative data on how parents divide the time required to care for their sick children during regular working hours. This data can be linked to the Unemployment Insurance administrative data to determine the parents' monthly labor earnings and to the Registro Civil register to obtain information on family composition. Secondly, the SANNA wage compensation scheme creates exogenous variation in the cost of adhering to gender norms. Additionally, the eligibility criteria for accessing SANNA benefits based on social security contributions allow us to control for eligibility to the program. Thirdly, although Chile has experienced rapid economic growth and increased female educational attainment, its female labor force participation rate is low compared to other OECD and Latin American countries. This suggests that traditional gender norms may still be a constraint for women's labor market participation, and our findings could inform policies aimed at promoting female labor force participation. Specifically, increasing the number of women entering the labor force by 100,000 could result in a 0.65 percentage point increase in Chile's GDP (Berlien et al., 2016).

3 Related Literature

The early literature that studied gender gaps in the labor market failed to include factors such as cultural differences and gender identity norms. This began to change in the last two decades, when George A. Akerlof and Rachel E. Kranton introduced the concept of identity into economics. Akerlof and Kranton (2000) develop a model where different social categories exist, for example, men and women. Each one of these social categories is associated with specific behavioral prescriptions. Individuals belong to one of these categories, which will conform their identity. If the individual behaves according to what is expected from their category, their sense of belonging to their category will increase. On the other hand, if the individual doesn't behave according to their category, their utility will decrease. Therefore, gender identity directly enters into the utility function, and can influence economic outcomes such as labor force participation, among others.

Several studies have estimated the importance of gender identity on the labor market. Fortin (2005, 2009) show that women's attitudes towards gender roles explain an important part of women's employment and incomes. Charles et al. (2009) use men's responses to gender role questions and find a strong correlation between men's responses and the gender gap in employment and earnings. Bertrand et al. (2015) show that the distribution of the share of income earned by the wife displays a sharp drop at 0.5, which they attribute to an aversion to the condition where the wife earns more than the husband due to gender identity norms. Bursztyn et al. (2017) show that unmarried MBA female students reported lower desired salaries and willingness to work long hours. Folke and Rickne (2020) show that job promotions can increase the probability of divorce for women, but not for men. They find that divorces are concentrated in more gender-traditional couples.

The paper closest to our study is Ichino et al. (2021). Exploiting variation from a tax reform in Sweden that generates changes in the marginal tax rates of spouses, the authors analyze whether this had an impact on the spousal division of home production. They find an elasticity of substitution in home production substantially lower than the one for market production. Our paper differs from theirs in two dimensions. First, the reform we exploit directly affects the allocation of home production within the family. Second, we do not directly observe time allocated to home production as they do. However, we have very detailed information on market production and paid/unpaid leave, which we will use as proxies for time allocated to home production.

If gender identity is an important factor that explains the labor force participation gender gap, then it is relevant to study the determinants of gender identity norms and how these norms have changed over time. Several studies have shown the persistence of these norms (Alesina et al., 2013; Grosjean and Khattar, 2019; Teso, 2019). One mechanism that explains the persistence of the gender identity norms is intergenerational transmission (Fernandez et al., 2004; Acemoglu et al., 2004; Farré and Vella, 2013). Even though these norms are persistent, some research has shown different factors that can change them. Some of these factors are the introduction of the Pill (Goldin and Katz, 2002; Goldin, 2006) and schooling environment (Dasgupta and Asgari, 2004; Maccoby, 1990, 1998; Lee and Marks, 1990; Paredes, 2014). Policy measures aimed at reducing gender differences in the labor market can also change views regarding gender roles. For example, Unterhofer and Wrohlich (2017) show that the father's quota in parental leave introduced in Germany in 2007 changed the attitudes in gender roles in the grandparents' generation. Therefore, it is also possible that the SANNA program, which default option gives the same number of days for mothers and fathers, could have an impact on gender norms.

4 Data

We use data from several administrative registers from 2010-2022 in Chile. All citizens in Chile have a mandatory identification number (RUT), which is used to link individuals to different administrative registers. We pool different registers to build a novel data set.

Our primary source of data is the SANNA register compiled by the Superintendencia de Seguridad Social (Suseso), which contains information on start- and end-dates of SANNA leave spells for individuals (SUSESO, 2023). The SANNA register also includes information on rejected permits due to parents not meeting the eligibility requirements related to the number of months contributing to social security. We link the SANNA records to the Registro Civil dataset using parents' RUT to obtain information on demographics and family composition, including the other parent's RUT, siblings of the cared-for child, age and gender of parents, marital status, age and gender of the seriously ill child, and gender and age of siblings.[7]

We obtain labor market variables from the Unemployment Insurance administrative data. This dataset includes information on workers' monthly contributions to the unemployment insurance and their employer (employer-employee dataset). We can determine each worker's gross monthly wage from their contributions to the system since contributions represent a fixed percentage of their gross wages.[8] We can also infer employment attachment and status.[9] Every worker with open-ended and fixed-term contracts employed in the private sector after 2002 is required to contribute to the system.

Summary statistics from SANNA register matched with Registro Civil register for the complete data set are presented in Table 2. From February 2018 until February 2022, 939 mothers and 726 fathers of 1,453 sick children have access to the SANNA program. Mothers take, on average, 167 days of leave. In turn, fathers take, on average, 115 days. Since SANNA

[7] We define a family as the mother, father and siblings of the cared-for child.

[8] Workers with open-ended and fixed-term contracts contribute 3% of their gross wages up to a predetermined level established yearly by law.

[9] One caveat is that we only have information of the formal sector.

Table 2: Descriptive Statistics

(a) Mothers

Variables	Obs.	Mean	Median	Std. Dev.	Min	Max
Days on leave	939	166.66	123	132.41	1	899
By health condition:						
Cancer:						
Days on leave	842	151.93	106	117.18	1	627
Transplants:						
Days on leave	142	142.39	99.5	132.41	14	441
End-stage or terminal illness:						
Days on leave	57	132.63	75	139.72	2	630
Serious accident:						
Days on leave	9	67.56	45	76.39	8	253

(b) Fathers

Variables	Obs.	Mean	Median	Std. Dev.	Min	Max
Days on leave	726	114.61	90	103.14	4	652
By health condition:						
Cancer:						
Days on leave	630	103.49	89	84.10	3	599
Transplants:						
Days on leave	120	104.13	90	90.42	15	338
End-stage or terminal illness:						
Days on leave	56	94.88	47.5	111.00	1	582
Serious accident:						
Days on leave	3	25.0	22	11.79	15	38

Source: SANNA register and Registro Civil.

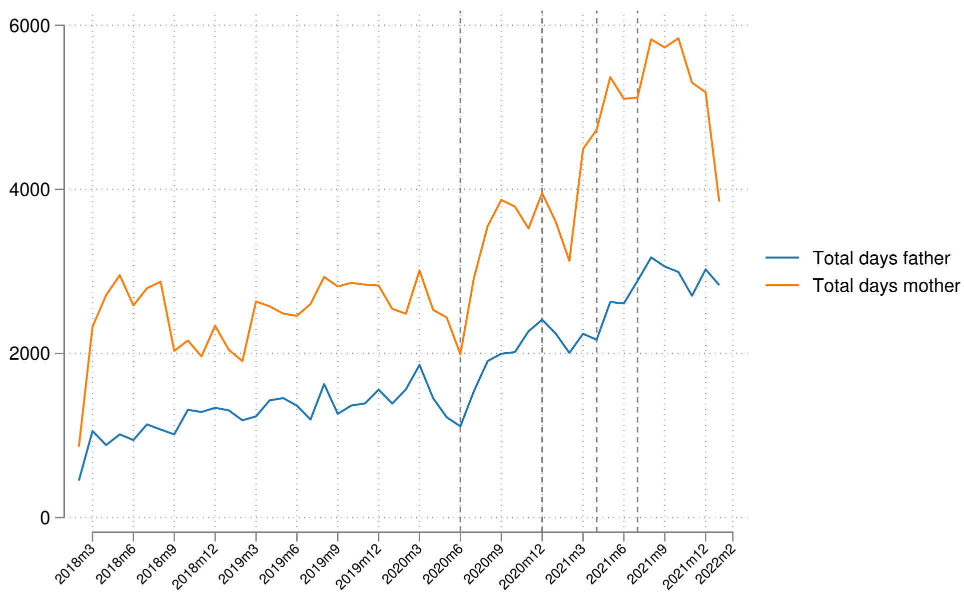
program's inception, the number of monthly days taken by parents have been increasing as more parents become aware of the program, and the coverage of new illnesses begin as established in the Law (Figure 3).[10] For the aforementioned reasons we project the number of SANNA beneficiaries to continue growing.[11] Due to the pandemic, the government

[10]The paid leave coverage schedule is the following: in February 2018 coverage begins for children sick with cancer and end-stage cancer; July 2018 coverage begins for children that need a solid organ and hematopoietic stem cell transplants; January 2020 for end-stage disease, and finally December 2020 for a child involved in a serious accident with risk of death or serious and permanent functional sequela.

[11]The estimated yearly number of children that will develop any of the serious illnesses covered by SANNA program is the following: cancers 958, transplants 64, end-stage or terminal illness 398, and accidents 2,622 (History of Law No. 21,063).

exceptionally increased the number of days available for parents. By health condition, cancer shows the largest number of beneficiaries of SANNA licenses (842 mothers and 630 fathers) with a ratio of usage between parents (1.48) similar to the aggregate (1.45) in terms of days of leave. It is interesting to note that in the case of end-stage or terminal illness, the same number of mothers and fathers accessed the program (57 and 56, respectively) probably due to the seriousness of the illness. However, still, the average number of leave days is larger for mothers (133 vs 95).

Figure 3: SANNA usage: monthly number of days of leave



Note: The figure shows the total number of leave days that mothers and fathers monthly take to care for a seriously ill child.

Because we want to study how families distribute the time allocated to child care, we restrict our sample to families where both parents are eligible for the paid leave at least once during the sample period.[12] Table 3 shows descriptive statistics for this restricted sample. From the total number of sick children in the complete sample, 835 are children for whom

[12] Given the nature of the eligibility criteria that is based on the number of months a worker has contributed to social security, parents' eligibility to the program changes monthly, and hence changes also during the sample period. Thus, we could also think of an alternative measures of eligibility such as parents being eligible at the time of the first paid leave. In this first version, we present results for the sample where both parents were eligible at least once during the sample period to maximize our sample.

both parents are eligible. In this sample, mothers take 179 days of leave and fathers take 116 days. By health condition, cancer shows the largest number of beneficiaries of SANNA licenses (617 mothers and 317 fathers) followed by transplants and end-stage of terminal illness. As in the complete sample, the case where parents share more equally their days of leave is end-stage or terminal illness probably due to the seriousness of the illness.

Table 3: Descriptive Statistics: restricted sample

(a) Mothers in the restricted sample

Variables	Obs.	Mean	Median	Std. Dev.	Min	Max
Days on leave	687	178.78	142	139.60	1	899
By health condition:						
Cancer:						
Days on leave	617	162.51	125	123.63	1	627
Transplants:						
Days on leave	110	143.65	100	98.51	14	441
End-stage or terminal illness:						
Days on leave	40	158.90	101.5	156.20	2	630
Serious accident:						
Days on leave	4	62.25	68	40.45	8	105

(b) Fathers in the restricted sample

Variables	Obs.	Mean	Median	Std. Dev.	Min	Max
Days on leave	360	116.09	90	103.46	6	652
By health condition:						
Cancer:						
Days on leave	317	105.55	88	88.61	3	599
Transplants:						
Days on leave	49	102.88	90	90.73	15	336
End-stage or terminal illness:						
Days on leave	35	91.71	45	108.69	1	582
Serious accident:						
Days on leave	2	18.5	18.5	4.95	15	22

Source: SANNA register and Registro Civil.

5 Empirical Framework

As mentioned earlier, the program provides each parent between 45 and 90 days per year depending on the type of illness, and parents can choose to take some or all of their entitled days or transfer them to the other parent. That is, parents have the flexibility to decide how many days to take and how to divide them. Our study focuses on analyzing the allocation of days between the mother and the father, using the fraction of days taken by the mother as our dependent variable. This fraction is calculated by dividing the number of days taken by the mother by the total number of days taken by both the mother and father.

We want to estimate the economic cost of adhering to traditional gender norms, so we leverage a key aspect of the SANNA program. As depicted in Figure 2, the program provides wage compensation up to a certain threshold, beyond which it remains fixed. This means that if one parent earns above the threshold and the other earns below it, it is economically efficient to transfer the benefit to the lower-earning parent. However, if families follow traditional gender norms and the mother assumes the caregiver role, they may allocate the days in a way that is not cost-effective. Our hypothesis is that mothers above the threshold are less likely to transfer the benefit compared to fathers above the threshold. By exploring this phenomenon, we can estimate the financial cost of traditional gender norms for families.

Let $\bar{w}_p$ be the average wage of parent p for the 12 months before the child becomes ill, with $p \in \{mother, father\}$. We divide families into four categories, according to which parent has the higher income, and whether mothers' income is above or below the threshold:

$$\text{Family type} = \begin{cases} 1 & \text{if } \bar{w}_{mother} < \bar{w}_{father} \text{ and } \bar{w}_{mother} < S_{max} \\ 2 & \text{if } \bar{w}_{mother} < \bar{w}_{father} \text{ and } \bar{w}_{mother} > S_{max} \\ 3 & \text{if } \bar{w}_{mother} > \bar{w}_{father} \text{ and } \bar{w}_{mother} < S_{max} \\ 4 & \text{if } \bar{w}_{mother} > \bar{w}_{father} \text{ and } \bar{w}_{mother} > S_{max} \end{cases}$$

where S_{max} is the threshold up to which parents get their salary fully compensated. If families maximize their income, we should observe that the mother takes all the leave in families type 2, while the father takes all the leave in families type 4 (in families type 3 all allocations are consistent with households maximizing income, while in families type 1 the efficient allocation depends on whether the father's income is above or below the threshold). However, if families follow traditional gender norms, we expect type 4 families to deviate from the income maximizing allocation.[13]

Let $\mathbf{1}\{\bar{w}_{mother} > S_{max}\}$ be a variable that takes the value of 1 when the mother's income is above the threshold. Also, let $\mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$ be a variable that takes the value of 1 when the mother has the highest income. Then, we can express the share of days allocated to the mother by each family type using the following equation:

$$\begin{aligned} ShareDaysMother &= \beta_0 + \beta_1 \mathbf{1}\{\bar{w}_{mother} > S_{max}\} + \beta_2 \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\} \\ &\quad + \beta_3 \mathbf{1}\{\bar{w}_{mother} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\} + \varepsilon \end{aligned} \tag{1}$$

We expect both economic incentives and gender norms to affect the share of days taken by the mother. In particular, gender norms and economics incentives should affect the magnitude of parameter β_3. When only economic incentives are considered, β_3 should be large and negative, since families for which we have $\mathbf{1}\{\bar{w}_{mother} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\} = 1$ should allocate all days to the father. However, if families follow traditional gender norms, this should moderate the magnitude of β_3. If gender norms are more important than economic incentives, we can even expect β_3 to be equal to zero.

We first estimate equation 1 for the whole sample of families where both parents are eligible. We include controls for parental age gap, parental average age, an indicator for younger siblings, total number of sick days, gender, age and type of illness of the child, and a proxy for divorced parents, which is constructed as an indicator for the presence of a younger half-sibling. We also include region and year fixed effects. The results of these estimations

[13] In Appendix B we develop a model where we consider six types of families.

will allow us to explore whether economic incentives are relevant in this setting.

To estimate the cost of following traditional gender norms, we need to compare the magnitude of β_3 in the absence of gender norms, and the magnitude of β_3 for families that follow traditional gender norms. We then compare the magnitude of β_3 in each sample. Potential proxies for gender norms in our data are average parental age (i.e., older parents should follow traditional gender norms), parental age gap (i.e., traditional families may have larger gaps being the father the older parent), and parents' employer economic sector (i.e., traditional families should work in traditional economic sectors). We have 3 proxies for traditional economic sectors: the share of women that work in the mother's sector, the share of women that work in the father's sector, and the difference between the share of women that work in the mother's and father's sectors.

To test if these variables are a good proxy for gender norms we explore the correlation between them and the share of days allocated to the mother. We expect traditional families to allocate a larger share of days to the mother, and thus, a significant correlation between our proxies of gender norms and our dependent variable. Table 4 shows the results of a regression of the dependent variable on our proxies for gender norms, and controls for labor market variables and the sick child variables. From our variables the only significant correlation with the share of days to the mother is parents average age. Hence, in what follows we will use parents age as our proxy for gender norms.

To divide the sample into families who follow traditional gender norms, and less traditional families we separate the sample by parental average age. As shown in Table 4, Panel A, we see that in older families, the percentage of days taken by the mother is larger. Also as we argue in the introduction, older families should be more likely to follow traditional gender norms than younger families. In particular, we split our sample of eligible families into families with average parental age above 38 years, and families whose average age is below or equal 38 years (we use 38 years to have a similar number of observations in each sample). As shown in Table 4, Panel B, the share of days taken by the mother in families

Table 4: Share of days taken by the mother and proxies for gender norms

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Continuous variables						
Parents sectors fem. share difference	0.001					0.001
	(0.001)					(0.001)
Father sector fem. share		-0.001				-
		(0.001)				-
Mother sector fem. share			0.000			-0.000
			(0.002)			(0.002)
Parents age gap				0.001		-0.001
				(0.002)		(0.003)
Parents average age					0.008***	0.008***
					(0.002)	(0.002)
Observations	815	815	815	835	835	815
	(1)	(2)	(3)	(4)	(5)	(6)
Panel B: Dummy variables						
Parents sectors fem. share difference	0.001					0.002
	(0.036)					(0.045)
Father sector fem. share		-0.024				-0.017
		(0.035)				(0.042)
Mother sector fem. share			-0.037			-0.046
			(0.058)			(0.061)
Parents age gap				-0.009		-0.017
				(0.028)		(0.028)
Parents average age					0.098***	0.100***
					(0.029)	(0.029)
Observations	835	835	835	835	835	835

Notes: Each column is a regression with the share of days taken by the mother as the dependent variable. In panel A the regressions include continuous variables, while in panel B they are included as dummy variables. The regression includes the following controls: Gender sick child, Sick child age, Pathology, Divorced parents, Younger sibling, Region, Year, Labor attachment$_{mother}$, Labor attachment$_{father}$, Months with last employer$_{mother}$, Months with last employer$_{father}$, Last economic sector$_{mother}$ and Last economic sector$_{father}$. ***, ** and * indicate statistical significance at the 99%, 95% and 90%, respectively.

with parental age above the median is 10 percentage points larger than for families with parental below the median.

The summary statistics for our variables in the restricted sample is presented in Table 5. The average wage of fathers is 25% larger than that of mothers (CLP 958,000 vs CLP 766,000), and 5.4% and 9.8% of the mothers and fathers have wages that are above the threshold more than half of the time in the 12 months before the child becomes ill, respectively.

Table 5: Descriptive statistics: restricted sample

	N	Mean	Std. dev.	Min.	Median	Max.
Labor market variables						
$\bar{w}_{mother}$	835	0.766	0.69	0.00	0.58	3.57
$\mathbf{1}\{\bar{w}_{mother} > S_{max}\}$	835	0.126	0.33	0.00	0.00	1.00
$\bar{w}_{father}$	835	0.958	0.82	0.00	0.72	5.67
$\mathbf{1}\{\bar{w}_{father} > S_{max}\}$	835	0.206	0.40	0.00	0.00	1.00
$\bar{w}_{mother} > \bar{w}_{father}$	835	0.394	0.49	0.00	0.00	1.00
Labor attachment$_{mother}$	835	0.792	0.25	0.00	0.91	1.00
Labor attachment$_{father}$	835	0.824	0.23	0.00	0.93	1.00
Months with last employer$_{mother}$	835	40.368	33.52	0.00	30.00	121.00
Months with last employer$_{father}$	835	38.359	35.58	0.00	26.00	121.00
Sick child variables						
Younger siblings indicator	835	0.659	0.47	0.00	1.00	1.00
Gender indicator	835	0.456	0.50	0.00	0.00	1.00
Sick child's age (years)	835	7.375	4.87	0.00	6.00	18.00
Divorced parents indicator	835	0.164	0.37	0.00	0.00	1.00
Total sick days	835	147.09	143.86	0.00	103.00	899.00
Parents characteristics						
Parental age gap (years)	835	2.559	4.99	-14.00	2.00	26.00
Parental average age (years)	835	37.831	6.95	20.00	37.50	59.50

Notes: $\bar{w}_p$ is the average wage of parent p for the 12 months before the child becomes ill in millions of Chilean pesos with $p \in \{mother, father\}$. $\mathbf{1}\{\bar{w}_{mother} > S_{max}\}$ takes the value of 1 when the mother's wage is above the threshold and $\mathbf{1}\{\bar{w}_{father} > S_{max}\}$ takes the value of 1 when the father's wage is above the threshold. Labor attachment measures parents' share of months with a regular job since the first observed regular job. Months with last employer is the number of months parents have worked in their last job. Younger siblings indicator takes the value of 1 if the sick child has siblings younger than 6 years old, gender indicator takes the value of 1 if the sick child is a girl, and divorced parents indicator takes the value of 1 if the sick child has a younger half-sibling. Parental age gap is computed as the fathers' age minus the mothers' age, and parental average age as the mean of the sick child parents' age.

To construct labor attachment we use data starting on January 2010 and calculate the percentage of months worked. Mothers in our sample have been employed 79.2% of the time span considered, while fathers have been employed 82.4%. Mothers have stayed 40.4 months with their last employer, while father have stayed 38.4 months. The average parental age gap, computed as the difference between the father's age and the mother's age, is 2.6 years, while the average age of the sick children is 7.4 years old and of their parents is 37.8 years.

In our sample, 45.6% of the sick children are girls, and 26.4% have their parents divorced. Finally, the average length of parents' total leave days is 147 or around 5 months.

In Table 6 we present summary statistics for families with average parental age above 38 years, and families whose average age is below or equal 38 years. These 2 groups of families have different characteristics for our selected labor market variables: both parents in older families have larger wages on average, and are more attached to the labor market (i.e. the share of months both parents have stayed in a regular job is larger). Mothers in the traditional group have also stayed longer in their last job. Parents in more traditional families have older children (in these families the average age of the sick child is larger, and the share of them that have a younger sibling is smaller), a smaller share of them are divorced, and the age gap between the mother and the father is larger.

6 Results

Table 7 shows the results of estimating equation 1 for the whole sample of families where both parents are eligible at least once during the sample period. In all equations we control for the indicator variables (and their interactions) defined in Section 5: $\mathbf{1}\{\bar{w}_{mother} > S_{max}\}$, $\mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$, and $\mathbf{1}\{\bar{w}_{mother} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$. We also control in all columns for region and year fixed effects. In column 2, we include our proxy for gender norms. In column 3 and 4, we add controls for observable characteristics of the sick child, for other siblings in the household, a proxy for divorced parents, and type of illness. In column 5 we also control for the length of the parental leave, and column 6 we add controls for parents' labor attachment and their employers economic sector. The results presented in Table 7 show that the coefficient for the mother's previous wage being above the threshold is positive. This indicates that in families with higher past incomes, the mother is taking a larger share of the parental leave. The large and positive coefficient for the mother's previous wage being above the threshold is somewhat surprising and cannot be explained by

Table 6: Comparing samples above vs. below age 38

	Avg. parental age ≤ 38	Avg. parental age > 38	T-test P-value
Labor market variables			
$\bar{w}_{father}$	0.84	1.10	0.000
$\bar{w}_{mother}$	0.64	0.92	0.000
Labor attachment$_{mother}$	0.76	0.83	0.000
Labor attachment$_{father}$	0.81	0.84	0.051
Months with last employer$_{mother}$	38.00	43.15	0.027
Months with last employer$_{father}$	36.90	40.07	0.201
Sick child variables			
Age	5.72	9.32	0.000
Gender	0.47	0.44	0.468
Younger sibling indicator	0.80	0.49	0.000
Pathology 1	0.00	0.01	0.529
Pathology 2	0.86	0.86	0.858
Pathology 3	0.14	0.13	0.745
Divorced parents indicator	0.20	0.13	0.009
Gender norms			
Parental age gap	1.78	3.47	0.000
Mother sector fem. share	0.45	0.45	0.889
Father sector fem. share	0.52	0.57	0.158
Parents sectors fem. share difference	0.53	0.50	0.317

Note: $\bar{w}_p$ is the average wage of parent p for the 12 months before the child becomes ill in millions of Chilean pesos with $p \in \{mother, father\}$. Labor attachment measures parents' share of months with a regular job since the first observed regular job. Months with last employer is the number of months parents have worked in their last job. Younger siblings indicator takes the value of 1 if the sick child has siblings younger than 6 years old, gender indicator takes the value of 1 if the sick child is a girl, and divorced parents indicator takes the value of 1 if the sick child has a younger half-sibling. Parental age gap is computed as the fathers' age minus the mothers' age, and parental average age as the mean of the sick child parents' age. Female share mother/father sector indicator takes the value of 1 if the share of female workers in the mother/father sector is larger than the median. Parents sectors difference is computed as the share of female workers in the mothers' sector minus fathers' sector.

the economic incentives generated by the threshold. This positive coefficient suggests that the sick leave could be seen by mothers as a luxury good. In particular, it is possible that mothers with high paying jobs also can afford the indirect costs of taking the parental leave (for example, costs in terms of future career development).

The results in Table 7 also show that when the mother has the highest past income, the

Table 7: Main results: Share of days taken by the mother

	(1)	(2)	(3)	(4)	(5)	(6)
$\mathbf{1}\{\bar{w}_{mother} > S_{max}\}$	0.312***	0.293***	0.314***	0.316***	0.296***	0.302***
	(0.057)	(0.057)	(0.055)	(0.055)	(0.060)	(0.066)
$\mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$	0.262***	0.262***	0.256***	0.257***	0.253***	0.203***
	(0.025)	(0.025)	(0.025)	(0.025)	(0.025)	(0.029)
$\mathbf{1}\{\bar{w}_{mother} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$	-0.464***	-0.483***	-0.496***	-0.501***	-0.476***	-0.490***
	(0.087)	(0.087)	(0.085)	(0.085)	(0.089)	(0.091)
Parental average age		0.008***	0.010***	0.010***	0.010***	0.009***
		(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Younger siblings			0.036	0.037	0.031	0.033
			(0.041)	(0.041)	(0.041)	(0.041)
Gender sick child			0.023	0.024	0.025	0.020
			(0.025)	(0.025)	(0.025)	(0.025)
Sick child (3 to 5 years)			-0.042	-0.039	-0.045	-0.049
			(0.040)	(0.040)	(0.040)	(0.041)
Sick child (6 to 13 years)			-0.017	-0.012	-0.022	-0.009
			(0.046)	(0.046)	(0.046)	(0.048)
Sick child (14 to 18 years)			-0.032	-0.025	-0.025	-0.007
			(0.060)	(0.060)	(0.060)	(0.062)
Divorced parents			0.143***	0.141***	0.153***	0.139***
			(0.040)	(0.040)	(0.040)	(0.041)
Other illnesses				-0.129	-0.158	-0.166
				(0.167)	(0.166)	(0.154)
More than one illness				-0.160	-0.236	-0.226
				(0.169)	(0.170)	(0.160)
Total sick days					0.000***	0.000*
					(0.000)	(0.000)
Labor attachment$_{father}$						-0.107*
						(0.063)
Labor attachment$_{mother}$						0.323***
						(0.060)
Months with last employer$_{father}$						-0.000
						(0.000)
Months with last employer$_{mother}$						-0.000
						(0.000)
Region	Yes	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes	Yes
Economic sector (last) mother	No	No	No	No	No	Yes
Economic sector (last) father	No	No	No	No	No	Yes
Adjusted R^2	0.116	0.138	0.155	0.154	0.161	0.202
Observations	835	835	835	835	835	835

Notes: The dependent variable is the share of days taken by the mother. $\bar{w}_p$ is the average wage of parent p for the 12 months before the child becomes ill in millions of CLP with $p \in \{mother, father\}$. Other controls are described in Table 5. Robust standard errors are presented in parentheses. ***, ** and * indicate statistical significance at the 99%, 95% and 90%, respectively.

percentage of days taken by the mother is higher than the percentage of days taken by the mother in families where the father has the highest past income. This result is somewhat puzzling because the difference in wages is used as a proxy of bargaining power. Thus, our results show that mothers with higher "bargaining power" take a larger percentage of the leave than mothers with lower bargaining power. In other words, mothers are not using this bargaining power to take a lower share of the parental leave. This again is consistent with the sick leave being seen by mothers as a luxury good.

Another interesting result shown in Table 7 is the effect of the proxy for divorced parents. In particular, we find that for families where the sick child has a younger half-sibling, the percentage of days taken by the mother is larger than the percentage of days in families with no younger siblings, or where the younger sibling share both the same mother and father.

Regarding our proxy for gender norms, we find that the average parental age has a positive effect on the percentage of days taken by the mother. In particular, the share of days taken by the mother is larger for older families, which is consistent with the hypothesis that older families follow more traditional gender roles.

Finally, our main variable of interest is the interaction between the indicator for the mother having the highest past income, and the mother's past income to be above the threshold. When this interaction takes the value of 1, families are incurring in a cost in terms of income for assigning days to the mother. Thus, if families care for economic incentives in this setting, then we should expect a negative coefficient for this interaction. Consistent with our hypothesis, we see a negative and statistically significant effect for this interaction in all columns, suggesting that economic incentives are still important in this setting, where families are deciding who takes care of a seriously ill child.

Table 8, Panel A and B, show the results of estimating the same regression for families whose average parental age is above 38 years and for families with an average parental age less than or equal 38 years respectively. The results for these two sub samples are consistent with the results for the whole sample of families. However, there are some important differences in the magnitude of the coefficients between the two sub samples. Our main variable of interest, the interaction between the indicator for the mother having the highest past income and the

Table 8: Results by parents' average age

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Average age > 38						
$\mathbf{1}\{\bar{w}_{mother} > S_{max}\}$	0.186**	0.204**	0.218**	0.219**	0.199**	0.226**
	(0.089)	(0.087)	(0.085)	(0.086)	(0.087)	(0.109)
$\mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$	0.213***	0.214***	0.207***	0.208***	0.205***	0.142***
	(0.035)	(0.034)	(0.035)	(0.035)	(0.035)	(0.041)
$\mathbf{1}\{\bar{w}_{mother} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$	-0.371***	-0.391***	-0.398***	-0.396***	-0.370***	-0.403***
	(0.118)	(0.116)	(0.116)	(0.117)	(0.120)	(0.126)
Observations	384	384	384	384	384	384
Panel B: Average age ≤ 38						
$\mathbf{1}\{\bar{w}_{mother} > S_{max}\}$	0.469***	0.440***	0.459***	0.459***	0.486***	0.469***
	(0.034)	(0.041)	(0.044)	(0.045)	(0.047)	(0.069)
$\mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$	0.302***	0.300***	0.292***	0.291***	0.287***	0.230***
	(0.035)	(0.035)	(0.035)	(0.035)	(0.035)	(0.043)
$\mathbf{1}\{\bar{w}_{mother} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$	-0.604***	-0.584***	-0.585***	-0.605***	-0.652***	-0.691***
	(0.140)	(0.142)	(0.139)	(0.139)	(0.138)	(0.162)
Observations	451	451	451	451	451	451
Panel C: Coefficient difference						
Difference	0.233	0.193	0.188	0.194	0.273	0.288
P-value	0.193	0.283	0.287	0.272	0.122	0.125

Notes: The dependent variable is the share of days taken by the mother. $\bar{w}_p$ is the average wage of parent p for the 12 months before the child becomes ill in millions of CLP with $p \in \{mother, father\}$. Controls are the same of Table 7. Robust standard errors are presented in parentheses. ***, ** and * indicate statistical significance at the 99%, 95% and 90%, respectively.

mother's past income being above the threshold, is smaller in magnitude in all columns for the sample of older families. This is consistent with the hypothesis that traditional gender norms are moderating the effect of economic incentives. Thus, this indicates that older families are incurring in a cost for following traditional gender norms.

Table 8, Panel C, shows both the difference in magnitude between our coefficient of interest for our two samples, and the p-value for the hypothesis that both coefficients are equal. As discussed above, the difference between younger and older families is always negative, indicating that the effect of economic incentives is larger for less traditional families. However, the p-value indicates that we cannot reject the hypothesis that the effect of economic incentives is different for both families. Even though the difference is not statistically significant at the conventional levels, the difference between both coefficients is rather large. The effect for older families is 57 to 68% of the effect for younger families. Or alternatively, traditional gender norms reduces the effect of economic incentives between 32 and 42%.

Finally, we do a back of the envelope calculation to measure the cost of following traditional gender norms. We calculate this cost as follows: for families where the mother's past wage is above the threshold and she has higher past wages than the father, the monthly cost of taking the parental leave is $\bar{w}_{mother} - S_{max}$ if the father's wage is below the threshold, and $\bar{w}_{mother} - \bar{w}_{father}$ if the father's wage is above the threshold. The average monthly cost in our sample is equal to CLP 486,557. We then multiply this number by the difference in the coefficients between older and younger families. In our preferred specification, this difference is equal to 0.288. Thus, the difference in monthly cost for the two types of families is equal to CLP 140,128.42. Finally, the parental leave length for families where the mother's past wage is above the threshold and her past wages are higher than the father past wages, is 200 days. Thus, the total average cost of following traditional gender norms is equal to CLP 947,268 (USD 1,200). This number is about one third of the average monthly income for these women.

7 Final Remarks

Gender identity norms can have real impacts on different labor market outcomes. In this paper, we explore the role of gender norms on families' decisions to specialize in market versus home production. In particular, we explore how families allocate a new paid parental leave to care for a seriously ill child in Chile. We first explore whether families follow economic incentives in this setting.

For this, we exploit the fact that the parental leave is designed to cover the wage of the beneficiary up to a threshold; and its fixed above this amount. Therefore, if one of the parents is above this threshold and the other is below, it is economically efficient to transfer the benefit to the other parent. Our results show that this economic incentive is relevant. For example, in families where the mother has the highest past income and her past wages, the share of days taken by the mother decreases in 46-50pp.

We then explore whether gender norms moderate the effect of economic incentives. For this, we exploit the fact that younger generations of parents are less likely to follow traditional gender norms. We compare the magnitude of the effects for families from younger and older cohorts to estimate the cost of following traditional gender norms. We find that the coefficient for younger families is 46 to 65% the coefficient for older families. Although we do not have sufficient power for this difference to be statistically significant at conventional levels, our results are consistent with older families incurring on a cost for following traditional gender norms.

Our results suggest that both economic incentives and gender norms affect labor market decisions, even when these decisions are as delicate as taking care of a seriously ill child. Thus, both elements should be considered in the design of such policies.

Finally, our empirical strategy is static and does not consider dynamic effects such as career concerns. More research is needed to explore how this intertemporal consideration may affect our estimates of the costs associated with gender norms.

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Appendix

A Additional figures and tables

Figure A.1: Gender roles towards mothers' paid work and the division of labor

Source: Encuesta Bicentenario UC, 2017.

B Six types of families

B.1 Empirical framework

In this appendix we develop an alternative empirical framework where we divide families into six categories, according to which parent has the higher income, and whether the income is above or below the threshold up to which the wage of the beneficiary is covered.

Let $\bar{w}_p$ be the average wage of parent p for the 12 months before the child becomes ill, with $p \in \{m, f\}$, where m denotes the mother and f denotes the father, families can be divided into six groups:

$$\text{Family type} = \begin{cases} 1 & \text{if } \bar{w}_{mother} < \bar{w}_{father} < S_{max} \\ 2 & \text{if } \bar{w}_{mother} < S_{max} < \bar{w}_{father} \\ 3 & \text{if } S_{max} < \bar{w}_{mother} < \bar{w}_{father} \\ 4 & \text{if } \bar{w}_{father} < \bar{w}_{mother} < S_{max} \\ 5 & \text{if } \bar{w}_{father} < S_{max} < \bar{w}_{mother} \\ 6 & \text{if } S_{max} < \bar{w}_{father} < \bar{w}_{mother} \end{cases}$$

where S_{max} is the threshold up to which parents get their salary fully compensated. If families maximize their income, we should observe that the mother takes all the leave in families type 2 and 3, while the father takes all the leave in families type 5 and 6 (in families type 1 and 4, all allocations are consistent with households maximizing income). However, if families follow traditional gender norms, we expect type 5 and 6 families to deviate from the income maximizing allocation.

Let $\mathbf{1}\{\bar{w}_{mother} > S_{max}\}$ be a variable that takes the value of 1 when the mother's income is above the threshold and $\mathbf{1}\{\bar{w}_{father} > S_{max}\}$ be a variable that takes the value of 1 when the father's income is above the threshold. Also, let $\mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$ be a variable that takes the value of 1 when the mother has the highest income. Then, we can express the share of days allocated to the mother by each family type using the following equation:

$$\begin{aligned} ShareDaysMother \;=\;& \beta_0 + \beta_1\mathbf{1}\{\bar{w}_{mother} > S_{max}\} + \beta_2\mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\} \\ & +\beta_3\mathbf{1}\{\bar{w}_{mother} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\} + \beta_4\mathbf{1}\{\bar{w}_{father} > S_{max}\} \\ & +\beta_5\mathbf{1}\{\bar{w}_{father} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\} + \varepsilon \end{aligned} \tag{B.1}$$

We expect both economic incentives and gender norms to affect the share of days taken by the mother. In particular, gender norms and economics incentives should affect the magnitude of parameter β_3. When only economic incentives are considered, β_3 should be large and negative, since families for which we have $\mathbf{1}\{\bar{w}_{mother} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\} = 1$ should allocate all days to the father. However, if families follow traditional gender norms, this should moderate the magnitude of β_3. If gender norms are more important than economic incentives, we can even expect β_3 to be equal to zero.

As in the main text, we first estimate equation B.1 for the whole sample of families where both parents are eligible. We include controls for parental age gap, parental average age, an indicator for younger siblings, total number of sick days, gender, age and type of illness of the child, and a proxy for divorced parents, which is constructed as an indicator for the presence of a younger half-sibling. We also include region and year fixed effects. The results of these estimations will allow us to explore whether economic incentives are relevant in this setting.

To estimate the cost of following traditional gender norms, we compare the magnitude of β_3 in the absence of gender norms, and the magnitude of β_3 for families who follow traditional gender norms. To do this, we separate the sample by parental average age. As we argue in the introduction in the main text, older families should be more likely to follow traditional gender norms than younger families. In particular, we split our sample of eligible families into families with average parental age above 38 years, and families whose average age is below 38 years (we use 38 years to have a similar number of observations in each sample). We then compare the magnitude of β_3 in each sample.

B.2 Estimation results

Table B.1 shows the results of estimating equation 1 for the whole sample of families where both parents are eligible at least once during the sample period. In all equations we control for the indicator variables (and their interactions) defined in Section 5: $\mathbf{1}\{\bar{w}_{mother} > S_{max}\}$, $\mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$, $\mathbf{1}\{\bar{w}_{father} > S_{max}\}$, $\mathbf{1}\{\bar{w}_{mother} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$ and $\mathbf{1}\{\bar{w}_{father} > S_{max}\} \times \mathbf{1}\{\bar{w}_{mother} > \bar{w}_{father}\}$. We also control in all columns for region and year fixed effects. In column 2, we add controls for observable characteristics of the sick child.

In column 3 we add controls for other siblings in the household and a proxy for divorced parents. In column 4, we include proxies for more traditional families. Finally, in column 5 we also control for the length of the parental leave.

The results presented in Table B.1 show that the coefficients for both maternal and paternal previous wage above the threshold are positive. This indicates that in families with higher past incomes, the mother is taking a larger share of the parental leave. This positive coefficient was expected for the indicator for paternal previous wage above the threshold: if paternal previous wage is above the threshold and the father's wage is higher than the mother's wage, it is economically efficient for the mother to take the whole parental leave. The large and positive coefficient for maternal previous wage above the threshold is somewhat surprising and cannot be explained by the economic incentives generated by the threshold. This positive coefficient suggests that parental leave could be seen by mothers as a luxury good. In particular, it is possible that mothers with high paying jobs also can afford the indirect costs of taking the parental leave (for example, costs in terms of future career development).

The results in Table B.1 also show that when the mother has the highest past income, the percentage of days taken by the mother is higher than the percentage of days taken by the mother in families where the father has the highest past income. This result is somewhat puzzling because the difference in wages is used as a proxy of bargaining power. Thus, our results show that mothers with higher "bargaining power" take a larger percentage of the leave than mothers with lower bargaining power. In other words, mothers are not using this bargaining power to take a lower share of the parental leave. This again is consistent with parental leave being seen by mothers as a luxury good.

Another interesting result shown in Table B.1 is the effect of the proxy for divorced parents. In particular, we find that for families where the sick child has a younger half-sibling, the percentage of days taken by the mother is larger than the percentage of days in families with no younger siblings, or where the younger sibling share both the same mother and father.

Regarding our two proxies for traditional families, we find that only the average parental age has a positive effect on the percentage of days taken by the mother. In particular, the share of days taken by the mother is larger for older families, which is consistent with the hypothesis that older families follow more traditional gender roles. The parental age gap does not have a significant effect on the share of days taken by the mother.

Finally, our main variable of interest is the interaction between the indicator for the mother having the highest past income, and the mother's past income to be above the threshold. When this interaction takes the value of 1, families are incurring in a cost in terms of income for assigning days to the mother. Thus, if families care for economic incentives in this setting, then we should expect a negative coefficient for this interaction. Consistent with our hypothesis, we see a negative and statistically significant effect for this interaction in all columns, suggesting that economic incentives are still important in this setting, where families are deciding who takes care of a seriously ill child.

Tables B.2 and B.3 show the results of estimating the same regression for families whose average parental age is more than 38 years and for families with an average parental age less than 38 years respectively. In general, the results for these two sub samples are consistent with the results for the whole sample of families. However, there are some important differences in the magnitude of the coefficients between the two sub samples. Our main variable of interest is smaller in magnitude in all columns for the sample of older families. This is consistent with the hypothesis that traditional gender norms are moderating the effect of economic incentives. Thus, this indicates that older families are incurring on a cost of following traditional gender norms.

Table B.4 shows both the difference in magnitude between our coefficient of interest for our two samples, and the p-value for the hypothesis that both coefficients are equal. As discussed above, the difference is always negative, indicating that the effect of economic incentives is larger for less traditional families. However, the p-value indicates that we cannot reject the hypothesis that the effect of economic incentives is different for both families. Even though the difference is not statistically significant at the conventional levels, the difference between both coefficients is rather large. The effect for older families is 46 to 65% of the effect for older families. Or alternatively, traditional gender norms reduces the effect of economic incentives between 35 and 54%.

Table B.1: Share of days taken by the mother

	(1)	(2)	(3)	(4)	(5)
$\mathbf{1}\{\bar{w}_f > S_{max}\}$	0.153***	0.160***	0.155***	0.117**	0.108*
	(0.055)	(0.055)	(0.056)	(0.057)	(0.056)
$\mathbf{1}\{\bar{w}_m > S_{max}\}$	0.188**	0.207***	0.224***	0.225***	0.213***
	(0.074)	(0.073)	(0.073)	(0.071)	(0.074)
$\mathbf{1}\{\bar{w}_m > \bar{w}_f\}$	0.280***	0.286***	0.279***	0.271***	0.266***
	(0.026)	(0.026)	(0.026)	(0.026)	(0.026)
$\mathbf{1}\{\bar{w}_m > S_{max}\} \times \mathbf{1}\{\bar{w}_m > \bar{w}_f\}$	-0.321***	-0.355***	-0.355***	-0.398***	-0.382***
	(0.109)	(0.106)	(0.105)	(0.105)	(0.107)
$\mathbf{1}\{\bar{w}_f > S_{max}\} \times \mathbf{1}\{\bar{w}_m > \bar{w}_f\}$	-0.207	-0.193	-0.193	-0.144	-0.135
	(0.153)	(0.153)	(0.151)	(0.152)	(0.154)
Gender sick child		0.028	0.028	0.026	0.027
		(0.026)	(0.025)	(0.025)	(0.025)
Sick child age (3 to 5 years)		-0.029	-0.031	-0.036	-0.042
		(0.040)	(0.041)	(0.040)	(0.040)
Sick child age (6 to 13 years)		0.040	0.019	-0.002	-0.012
		(0.037)	(0.047)	(0.047)	(0.047)
Sick child age (14 to 18 years)		0.082*	0.059	-0.013	-0.014
		(0.044)	(0.060)	(0.061)	(0.061)
Other illnesses		-0.178	-0.182	-0.139	-0.167
		(0.156)	(0.156)	(0.166)	(0.165)
More than one illness		-0.217	-0.216	-0.168	-0.240
		(0.159)	(0.159)	(0.168)	(0.169)
Younger siblings			0.003	0.038	0.033
			(0.041)	(0.041)	(0.041)
Divorced parents			0.116***	0.134***	0.147***
			(0.040)	(0.041)	(0.041)
Parental age gap				-0.001	-0.000
				(0.003)	(0.003)
Parental average age				0.009***	0.009***
				(0.002)	(0.002)
Total sick days					0.000***
					(0.000)
Region	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.124	0.131	0.140	0.157	0.163
Observations	835	835	835	835	835

Standard errors in parentheses

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

Table B.2: Share of days taken by the mother: average parental age > 38

	(1)	(2)	(3)	(4)
$\mathbf{1}\{\bar{w}_f > S_{max}\}$	0.091	0.082	0.074	0.070
	(0.069)	(0.070)	(0.070)	(0.071)
$\mathbf{1}\{\bar{w}_m > S_{max}\}$	0.123	0.143	0.152	0.137
	(0.104)	(0.099)	(0.100)	(0.100)
$\mathbf{1}\{\bar{w}_m > \bar{w}_f\}$	0.229***	0.227***	0.223***	0.220***
	(0.038)	(0.039)	(0.039)	(0.038)
$\mathbf{1}\{\bar{w}_m > S_{max}\} \times \mathbf{1}\{\bar{w}_m > \bar{w}_f\}$	-0.262*	-0.290**	-0.284**	-0.262*
	(0.137)	(0.134)	(0.136)	(0.136)
$\mathbf{1}\{\bar{w}_f > S_{max}\} \times \mathbf{1}\{\bar{w}_m > \bar{w}_f\}$	-0.227	-0.193	-0.188	-0.187
	(0.180)	(0.183)	(0.183)	(0.186)
Gender sick child		0.015	0.014	0.015
		(0.037)	(0.037)	(0.037)
Sick child age (3 to 5 years)		-0.123**	-0.116*	-0.114*
		(0.062)	(0.063)	(0.063)
Sick child age (6 to 13 years)		-0.103*	-0.139**	-0.141**
		(0.054)	(0.069)	(0.069)
Sick child age (14 to 18 years)		-0.096	-0.141*	-0.140*
		(0.059)	(0.084)	(0.084)
Other illnesses		0.064	0.063	0.048
		(0.257)	(0.257)	(0.257)
More than one illness		0.048	0.049	0.013
		(0.260)	(0.260)	(0.264)
Younger siblings			-0.038	-0.041
			(0.061)	(0.061)
Divorced parents			0.108*	0.117*
			(0.060)	(0.062)
Total sick days				0.000
				(0.000)
Region	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes
Adjusted R^2	0.101	0.096	0.098	0.097
Observations	384	384	384	384

Standard errors in parentheses

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

Table B.3: Share of days taken by the mother: average parental age ≤ 38

	(1)	(2)	(3)	(4)
$\mathbf{1}\{\bar{w}_f > S_{max}\}$	0.197*	0.240**	0.224**	0.204**
	(0.101)	(0.102)	(0.103)	(0.102)
$\mathbf{1}\{\bar{w}_m > S_{max}\}$	0.285***	0.259***	0.280***	0.322***
	(0.098)	(0.099)	(0.100)	(0.102)
$\mathbf{1}\{\bar{w}_m > \bar{w}_f\}$	0.317***	0.325***	0.315***	0.308***
	(0.036)	(0.036)	(0.036)	(0.036)
$\mathbf{1}\{\bar{w}_m > S_{max}\} \times \mathbf{1}\{\bar{w}_m > \bar{w}_f\}$	-0.500**	-0.446**	-0.485**	-0.568***
	(0.215)	(0.218)	(0.211)	(0.204)
$\mathbf{1}\{\bar{w}_f > S_{max}\} \times \mathbf{1}\{\bar{w}_m > \bar{w}_f\}$	-0.008	-0.061	-0.006	0.050
	(0.267)	(0.274)	(0.263)	(0.250)
Gender sick child		0.048	0.045	0.046
		(0.036)	(0.036)	(0.035)
Sick child age (3 to 5 years)		-0.009	-0.016	-0.031
		(0.050)	(0.051)	(0.051)
Sick child age (6 to 13 years)		0.079	0.070	0.050
		(0.049)	(0.062)	(0.062)
Sick child age (14 to 18 years)		0.135*	0.091	0.082
		(0.073)	(0.088)	(0.088)
Other illnesses		-0.476***	-0.471***	-0.495***
		(0.061)	(0.063)	(0.073)
More than one illness		-0.522***	-0.508***	-0.601***
		(0.074)	(0.074)	(0.087)
Younger siblings			0.059	0.047
			(0.059)	(0.059)
Divorced parents			0.130**	0.145***
			(0.055)	(0.055)
Total sick days				0.000***
				(0.000)
Region	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes
Adjusted R^2	0.159	0.171	0.187	0.202
Observations	451	451	451	451

Standard errors in parentheses

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

Table B.4: Results equal coefficient test

	(1)	(2)	(3)	(4)
Average parental age ≤ 38	-0.500**	-0.446**	-0.485**	-0.568***
	(0.215)	(0.218)	(0.211)	(0.204)
Average parental age > 38	-0.262*	-0.290**	-0.284**	-0.262*
	(0.137)	(0.134)	(0.136)	(0.136)
Difference between coefficients	-0.238	-0.156	-0.200	-0.306
p-value equal coefficient test	0.339	0.530	0.409	0.197

Standard errors in parentheses

***, ** and * indicate statistical significance at the 99%, 95% and 90%, respectively.