The Effect of Paid Maternity Leave on Families' Living Arrangements

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Abstract

I analyze the effect of two maternity leave (ML) reforms on families' living arrangements using regression discontinuity and differences-in-differences methodology. The first reform is an extension of ML from 60 days to one year and offers 65% of income before childbirth. It has insignificant impact on families' living arrangements. The second reform is an extension of ML from one year to two years and offers 85% of income. It increases the probability of single-motherhood accompanied by a decrease in the likelihood of getting married. The second's reform results support the independence hypothesis- an increase in women's income reduces the need to pool resources and makes household work specialization less advantageous leading to higher probability of single motherhood.

JEL Classification: J08,J1,J13

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1 Introduction

Most developed countries offer paid maternity leave for mothers after childbirth. These policies differ significantly across countries in terms of duration, benefits, job protection, and eligibility. They are designed to increase women's labor force attachment, improve health outcomes for children and mothers, and balance work and family responsibilities. To better understand the impacts of maternity leave (ML), I analyze how two extensions of ML affect family living arrangements (e.g., married, single, divorced) in Romania. The first policy extended ML from 60 days to one year (i.e., the child's first birthday), and it was implemented as a presidential decree on the 19^{th} of January 1990 in Romania. The second policy, enacted in 1997, repealed the 1990 ML reform. The 1997 policy extended ML from one year to two years and increased benefits. Each child's birth date was used to determine which ML policy applied to the mother. Mothers who gave birth after each policy reform was implemented were eligible for the extended ML.

I use regression discontinuity and differences-in-differences methodologies to estimate the reforms' causal impact on family living arrangements. The data allows me to distinguish divorced parents from mothers who never married (i.e., single motherhood), and the married category includes cohabiting parents. In addition, I analyze how maternity leave policy impacts the probability of grandparents living in the household. The 1990 maternity leave policy had no significant impact on the probability of parents being married or divorced or the mother being single at childbirth (i.e., single motherhood) two years after the implementation date. In contrast, the 1997 maternity leave policy decreased parents' probability of being married by 1.1 percentage points five years after the reform. This effect is accompanied by an increase of 1.2 percentage points in the probability of the mother being single. There are no significant differences in the ML effect on family living arrangements based on the child's gender or birth order.

The research on maternity leave impacts on women's labor market outcomes and maternal health outcomes is extensive. A growing literature provides evidence that long paid leave tend to hurt women's labor force participation (Schönberg and Ludsteck (2014), Lalive and Zweimüller (2009)) and wages (Ruhm (1998)), while short leaves increases it (Baum (2003), Baker, Gruber, and Milligan (2008), Bergemann and R. Riphahn (2015), Rossin-Slater, R. J. Ruhm, and Waldfogel (2013)).

The research literature has extensively examined ML impacts on women's labor market outcomes and maternal health outcomes. A growing body of literature provides evidence that longer paid leave tends to adversely affect women's labor force participation (Schönberg et al. (2014), Lalive et al. (2009)) and wages (Ruhm (1998)), while shorter paid leave tends to increase their labor force participation and wages (Baum (2003), Baker et al. (2008), Bergemann et al. (2015), Rossin-Slater et al. (2013)). Previous studies have shown that paid leave improves a range of health outcomes, including body mass index (BMI), blood pressure, pain, and mental health (Bütikofer, Riise, and Skira (2021)), and depressive symptoms and reduces the number of outpatient visits (Chatterji and Markowitz (2005)). However, Baker and Milligan (2008) found that maternity leave had no effect on self-reported maternal health in Canada.

Less is known about the effect of paid leave on marital stability, and few studies have analyzed the effect of paternity leave (i.e., leave offered only to fathers) on family living arrangements. For example, Avdic and Karimi (2018) examine the "daddy-month" reform, implemented in 1995 in Sweden, which offered wage replacement and could not be transferred to mothers. The authors found that, for married couples, if the father was eligible for paternity leave, these couples had a higher probability of separation than couples in which the fathers were ineligible. Cools, Fiva, and Kirkebøen (2015) that one-month paternal leave, enacted in 1993 in Norway, had no significant effect on the probability of parents being married 14 years after the reform was implemented.

Research on the impact of parental leave (i.e., leave offered to either parent) on family living arrangements yields ambiguous results. Dahl, Løken, Mogstad, and Salvanes (2016) analyze the effect of six parental leave policies, which extended leave from 18 to 35 weeks in Norway, on long-term marital stability. These expansions did not affect marriage or divorce, except for the 1992 reform, which increased the probability of parents being married 14 years after childbirth, conditional on not being married the year before the reform. However, Kluve and Schmitz (2018) show that the 2007 parental leave, which replaced a fixed benefit program that provided 67% pre-childbirth labor earnings and introduced two months of paternal leave, decreased the probability of marriage. Their research was extended by Cygan-Rehm, Kuehnle, and R. T. Riphahn (2018), who show that parents are more likely to cohabit and that the results are driven by a decrease in single motherhood among higher-income women. These results are consistent with the hypothesis that improved financial situation for households -reducing intra-household conflicts leads to lower single motherhood-, and father involvement hypothesis (Morgan, Lye, and Condran (1988))- that increased paternal involvement in childcare increases marital stability.

This paper contributes to the research literature in three ways. First, it considers two different paid leaves (i.e., maternity leave) to address the knowledge gap regarding the effects of various ML policies on family living arrangements. Second, it contributes to the literature focusing on children's gender and living arrangements. Finally, it provides evidence of ML impact on family's living arrangements for an Eastern-European country with a different institutional and cultural background than Western-European countries.

The remainder of the paper proceeds as follows. Section 2 presents the literature review, and Section 3 focuses on the institutional background. Data are described in Section 4, while Section 5 explains the empirical strategy. Results are explained in Section 6, and conclusions are provided in Section 7.

2 Theoretical framework

This section describes the mechanisms through which maternity leave might affect family living arrangements. I focus on the effect of the ML policy on three mutually exclusive outcomes: the probability of a child living with married parents, single-motherhood, and divorced parents.

Family living arrangements may be affected by changes in income. Becker (1981) argues that the increase in women's labor force participation increased their opportunity cost of childcare and household activities. Women's earnings make household work specialization less advantageous between couples, leading to lower rates of marriage. The economic independence hypothesis states that increasing women's income may reduce the need to pool resources, increasing the probability of single motherhood. To test the economic independence hypothesis, Cancian and Meyer (2014) use data that captured randomly assigned differences in child support for single mothers who receive Temporary Assistance for Needy Families (TANF) benefits. In most states, the government retains child support if women receive TANF. However, Wisconsin implemented child support uniquely: one group received $50 from the child support in addition to TANF benefits, while the other group received the whole child support amount. The authors find that the increase in income led to lower cohabitation rates between mothers and their partner who is not the biological father, supporting the economic independence hypothesis. Simultaneously, there were no significant effects on the probability of being married or cohabiting with the biological father. On the other hand, an increase in income may have eased financial hardship in a household, thereby improving household welfare by mitigating intra-household conflicts that emerge from such hardship (Cancian et al. (2014)). This mechanism may reduce single motherhood probability, leading to an ambiguous effect of an increase in income on family living arrangements.

The standard theoretical framework for analyzing consumption behavior and labor supply assumes that individuals have common consumption preferences. Thus, individuals maximize the utility using the total family income (Becker (1981)). Nevertheless, the marriage theoretical framework rejects the income pooling assumption, and the relative wage of parents is used. An increase in a mother's relative wage is seen as improving women's bargaining power (Lundberg and Pollak (1996)). Women tend to spend a higher share of their income on children, which is seen as a family investment decreasing single-motherhood probability (Lundberg, McLanahan, and Rose (2007)).

Lastly, the ML may generate a differential effect on family living arrangements based on the child's sex. Giuliano (2007) suggests two hypotheses. First, fathers spend more time with their sons compared to their daughters, which increases not only the father's happiness (Mammen (2011), Choi, Joesch, and Lundberg (2008)) but also the mother's satisfaction because she will perceive her husband as a better father. These effects help improve marital stability. Second, girls and boys have different needs, so women may perceive that their sons, more than daughters, may need a father's presence, which may reduce divorce probability for couples with sons. Giuliano (2007) finds support for both hypotheses by showing that mothers with sons reported greater marital satisfaction, positive views of their husbands, and greater involvement of fathers in childcare, thereby reducing the probability of divorce.

Dahl and Moretti (2008) show that a first-born daughter is less likely to live with her father than a first-born son. This result is driven by the fact that women with first-born daughters are less likely to marry, more likely to divorce, and more likely to obtain custody of daughters than sons. Lundberg et al. (2007) find that sons born to unmarried couples are more likely than daughters to receive a father's surname, and there is greater paternal involvement in childcare around the time of birth. Still, these do not affect family living arrangements. However, sons born to married parents are more likely to live in an intact family than daughters one year after birth.

3 Institutional Background

3.1 Maternity Leave Reform of 1990

Decree Number 31/1990 extended the length of maternity leave from 112 days to one year (i.e., the child's first birthday) in Romania. This extension aimed to facilitate the bond between mothers and their new children, provide parents with the flexibility to balance work and family life, and encourage women to remain in the labor force.

The financial support provided by ML was 65% of mothers' pre -childbirth monthly income. To qualify, mothers needed to be employed at the onset of pregnancy, and they needed to contribute to the social security system.

The reform was implemented less than one month after the fall of the Communist regime in the country. It was signed by the president on the 18^{th} of January 1990 and enacted as a decree the following day. The policy was not discussed or debated prior to implementation because it was enacted shortly after the democratic government came to power. Mothers with children born after $19^{th}$of January 1990 were eligible for the new ML policy, while those with children born before this date did not qualify for the ML extension.

After the fall of the Communist regime, there were extensive structural and cultural changes in the country. These policies were not implemented simultaneously with the maternity leave reform. The most crucial policy related to family living arrangements was Law no.85/1992, which allowed houses owned by the government (i.e., approximately 70% of total housing) to be purchased by their current tenants (World Bank, Report No.: 106856 (2015)). The government feared that liberalization would increase prices for goods and services, including property, so it set fixed prices for state-owned homes based on the number of bedrooms and location. The nominal average wage increased rapidly due to inflation, while apartment prices remained unchanged, which contributed to Romania having one of the highest home-ownership rates in Europe. These changes did not occur overnight, so they are unlikely to affect family living arrangements.

3.2 Maternity Leave Reform of 1997

The decree-law 31/1990 regarding maternity leave was rescinded in 1997 when Law 120/1997 was enacted. The time and benefits allocated exclusively to mothers immediately before and after childbirth (112 days) remained unchanged. However, the extension of the maternity reform underwent essential changes.

It increased the monetary benefits and the duration of ML without changing the eligibility conditions. Mothers received 85

Although Romanian citizens did not anticipate the 1997 reform, it did not materialize as quickly as the 1990 reform. The latter reform was first adopted by the Senate in May 1997 and sent to the president in July 1997. The official law was established on the 11th of July 1997. The short legislative process did not threaten the identification strategy because mothers could not plan or change their delivery date in response to this policy.

Another identification concern is whether any policies were implemented around the same time, which may affect the outcomes of interest, family living arrangements (e.g., single, divorced, married, and grandparents present in the household). Analyzing the official records, I did not identify any policies that might affect my outcomes. However, essential changes in the public pension system and other social security rights occurred in 2000, three years after the maternity leave reform.

4 Data

This study's data are drawn from a 15% sample of the 1992 Romanian census provided by the Inter-university Consortium for Political and Social Research. The principal investigator for this dataset is the United Nations Economic Commission for Europe (2013) mainly to assure that Romania has a reliable and valuable methodology for collecting data, given that only recently, before the survey, it had become a democratic country. The second dataset is a 10% sample of the 2002 Romanian census provided by the Minnesota Population Center (2020).

4.1 Descriptive Statistics

I match mothers and children using a unique identification number. The observations without the mother ID numbers or data on family living arrangements are dropped from the sample. I further refine the sample to children for whom I can identify the date of birth.

Table 1 provides descriptive statistics for the 1992 Romanian census sample. Panel A shows the mean of the outcome variables: 35% of children have their grandparents present in the household, 95.4% of mothers are married, 2.9% of single mothers (never married), and 1.7% of mothers are divorced. Characteristics for children and mothers are presented in Panel B and Panel C, respectively. A small percentage of mothers completed a university degree (3.8%).

Table 2 presents descriptive statistics for the 2002 Romanian census sample. There are some changes in family living arrangements compared to the previous census. Only 25.2% of children have their grandparents living in the same household, and there is a decrease in the marriage rate from 95.4% in 1992 to 79.7% in 2002. Rates of single motherhood and divorce were higher in 2002. There is an increase in mothers' educational attainment; the relative frequency of tertiary education increased to 5.6

5 Empirical Strategy

5.1 Methodology

In contrast to the 2002 Romanian census, which only includes month and year of birth, the 1992 Romanian census consists of the day of birth, which allows me to identify mothers who were eligible for maternity leave. I restrict the sample to mothers who had a child 60 days before the policy was implemented (i.e., control group: no maternity leave) and 60 days after (i.e., treatment group: maternity leave). I use a regression discontinuity design to study the family living arrangements two years after implementing the reform.

I estimate the following equation:

$$Y_i = \beta_0 + \beta_1 T_i + \beta_2 f(d) + \beta_3 f(d) * T_i + \beta_4 X_i + \epsilon_i \tag{1}$$

where Y_i is an indicator for various family outcomes: married, single-motherhood, divorced. In addition, I include as an outcome the probability of grandparents living in the household two years after the reform. T_i is a dummy variable that is assigned a value of 1 if the reference child was born within 60 days after reform and 0 if born within 60 days before. The running variable was normalized to 0 for January 19^{th}, 1990 and f is a polynomial in the running variable d (day). I include maternal and time-related controls, such as mother's age and education, day of the week fixed effects corresponding to child's birth, gender, and a dummy variable for multiple births. Standard errors are clustered at the day of birth level to avoid the issues of using a discrete running variable in the regression discontinuity framework (Lee and Card (2008)). Estimates are adjusted for mass points in the running variable for the local polynomial regression (Calonico, Cattaneo, and Titiunik (2014)).

The 2002 Romanian census includes only the month and year of the birthday and not the specific day of birth. The new maternity leave reform was published on the 11^{th} of July 1997, so in my analyses, I drop all children born in July because I cannot distinguish mothers who had access to the old ML from those with access to the new ML. Moreover, regression discontinuity does not perform well using only a few mass points in the running variable (in this case, the month of birth). Thus, I estimate a differences-in-differences specification, in which I restrict the sample to mothers who gave birth within 90 days before or after the reform was implemented as the treatment group, while for the control group, I use mothers who gave birth in the same months, but one year earlier.

$$Y_i = \alpha_0 + \alpha_1 Treat_i + \alpha_2 After_i + \alpha_3 Treat_i * After_i + \alpha_4 X_i + \epsilon_i \qquad (2)$$

where Y_i is an indicator for various family outcomes: married, single-motherhood, divorced, or grandparents present in the household, five years after the reform. $Treat_i$ is an indicator assigned a value of 1 if the mother gave birth within 90 days before or after the reform, and 0 if mothers gave birth in the same months but a year earlier.

$After_i$ is an indicator variable equal to 1 if the child is born in the months "after" the reform (August to October) in the treatment year (i.e., 1997) and control year (i.e., 1996), and assigned a value 0 if the child is born "before" the reform (April to June) during the same years. $Treat_i * After_i$ is an interaction term between the previous two variables. X_i is a vector that includes mother's age, an indicator for the maternal level of education, dummies for multiple births and child's sex, and month fixed effects are corresponding to child's birth. The coefficient of interest is α_3, which refers to children born between August to October (1997), and identifies the Intention to Treat effect of ML on family living arrangements.

5.2 Identification Assumptions

The identification assumption for regression discontinuity is that potential outcomes are continuous as a function of the running variable. This means that the only factor which causes the outcome to change at the threshold is the treatment. This assumption would be violated if mothers manipulated the day of childbirth to qualify for a ML reform. Figure 1, I plot the density of birth at the day level around the reform's implementation date. It shows regular fluctuations in the density of birth, with no evident jumps at the threshold. I empirically test the discontinuity at the threshold using the McCrary local linear estimator test and plot it in Figure 2. The solid black lines show the histogram's two local linear smoothings with the corresponding 95% confidence intervals. The confidence intervals overlap at the threshold which means that the test fail to reject the null hypothesis that there is a discontinuity in the number of birth when the policy was implemented.

Furthermore, I run a balancing check between the two groups to ensure that they are not different. In Table 3, I present the coefficients of the reform impact on covariates that could influence family living arrangements using a linear polynomial regression discontinuity. Panel A shows a child's characteristics: gender and being a first-born child, while Panel B shows mother's characteristics: age, a dummy variable for college-educated, high skill occupation, and employment. None of these six outcomes are statistically significant, suggesting no differences in observables between the treatment and control groups.

The crucial assumption for the internal validity of the differences-in-differences strategy is the parallel trend assumption. It requires that the difference in the outcome of interest for the two groups is constant over time in the absence of treatment. Because it involves counterfactuals, it is not possible to test this assumption. An indirect test of the parallel trend assumption is to check the pre-treatment balance between control and treatment groups. The treatment group represents mothers who had a child around the 1997 reform, while the control group represents mothers who had a child during the same months, but one year before, in 1996. Given that my groups are selected for a different period, and outcomes were reported in 2002, I cannot visually inspect data on their family living arrangements before the maternity leave policy was implemented. Instead, I validate the causal effect of a differences-in-differences methodology using two placebo reforms: one in 1996 and one in 1998. The results are reported in Table 9. The estimates using the placebo reforms are statistically insignificant.

6 Results

6.1 Results for the 1990 reform

Table 4 presents the effect of the maternity leave enacted in 1990 on family living arrangements. Each coefficient is obtained from a separate regression using estimation strategy 1. The header of the table represents the outcome variables. The first three outcomes are mutually exclusive: the probability of a child living with married parents, single mothers, or divorced parents. The last outcome of interest is whether grandparents live in the household. The intuition would be that mothers who were not eligible for ML would seek help caring for the child from grandparents, leading to them moving into the household. For each outcome, the first column shows the estimates from a linear polynomial regression, while the second column shows the estimates from a local linear polynomial regression using the bandwidth selection algorithm proposed by Calonico et al. (2014).

The 1990 reform had no significant effect on the probability of being married, single, or divorced. The 0.9 percentage point increase in the likelihood of being married came from a decrease in single motherhood by 0.3 percentage points and 0.6 percentage points in divorce. It seems that there is no significant difference in the likelihood of grandparents being present in the household between mothers who were eligible for ML and those who were ineligible. The results are represented visually in Figure 3. The solid lines depict a linear fit, on each side of the cutoff, on the probability of being married, single, or divorced or of grandparents living in the same household. The circles represent the percentage average bin of these outcomes two years after the reform.[1] Shaded areas represent 95% confidence intervals.

I modify Equation 1 to analyze the differential effect of the reform on family living arrangements based on the child's gender or birth order. The terms added are: a group indicator (for whether the child is female or a firstborn child), an interaction term of the group indicator and one for being eligible for maternity leave, as well as the interaction between the group indicator and the running variable (day of birth).

$$Y_i = \beta_0 + \beta_1 T_i + \beta_2 f(d) + \beta_3 f(d) * T_i + \beta_4 T_i * Group_i + \beta_5 f(d) * Group_i + \beta_6 X_i + \epsilon_i \quad (3)$$

[1] Due to the small window around the reform, the average outcome is computed for four-day bins.

The results in Panel A of Table 5 do not align with finding in previous studies (i.e.,Dahl et al. (2008), Choi et al. (2008)) that suggested that daughters increase the likelihood of single motherhood or divorce. In Panel B, I test if there are significant changes for families with first-born children because there is a higher probability of parents getting married (Cygan-Rehm et al. (2018)). The results confirm a significant increase of 2.4 percentage points in the probability of being married for women who were eligible for maternity leave. This goes along with a decrease in the probability of divorce of 1.2 percentage points and 1 percentage point in the probability of single motherhood, respectively.

I perform a series of robustness checks. First, I explore the sensitivity of the estimates to a different window and polynomial. I restrict the sample to a smaller window (i.e., 30 days before and after the policy implementation date), then I run a second-order polynomial regression on the original sample. Second, I simulate two placebo reforms: (a) one year before the policy was implemented and (b) one year after implementation. Finally, I use a different estimation strategy (i.e., a differences-in-differences regression) with various window sizes ranging from two months to four months. All coefficients are statistically insignificant, but they are similar in magnitudes. The results are presented in Table 6.

6.2 Results for the 1997 reform

Table 7 reports the estimates on the effect of maternity leave policy implemented in 1997 on family living arrangements using differences-in-differences regression. The header represents dummy variables for the outcomes of interest: married, single, divorced, and grandparents living in the household. Because the first three outcomes are mutually exclusive, I perform a multinomial logit regression as a robustness check.

Mothers who had access to ML are 1.2 percentage points more likely to be single five years after the reform. This result supports the economic independence hypothesis, which states that an increase in women's income reduces the need to pool resources, thus increasing the probability of single-motherhood.

The 1.2 percentage point increase in the probability of single motherhood is consistent with a 1.1 percentage point decrease in the probability of being married for the treatment group (i.e., women with maternity leave). Women's earnings make household work specialization less advantageous between couples, leading to lower marriage rates (Becker (1981)). The estimates are statistically significant at 0.001 level.

Analyzing ML reform in Germany, Cygan-Rehm et al. (2018) find a 3.1 percentage points decrease in single-motherhood combined with a 4.3 percentage points increase in the probability of cohabitation. Cygan-Rehm et al. (2018) separate married couples from cohabiting couples, and the single mothers' category included divorced parents. My outcomes of interest are slightly different: I can distinguish divorced mothers from mothers who never married, while the married category includes cohabiting parents. Thus, it is difficult to compare my results with the Cygan-Rehm et al. (2018) findings.

To study heterogeneity effects, I add to Equation 2, a group indicator (for whether the child is a female or firstborn child) and an interaction term of the group indicator and the dummy variable for a group, as well as the interaction between the group indicator and the indicator for time period, and the triple interaction between the previous three variables.

$$\begin{aligned} Y_i = &\alpha_0 + \alpha_1 Treat_i + \alpha_2 After_i + \alpha_3 Treat_i * After_i + \alpha_4 Treat_i * Group_i \\ &+ \alpha_5 After_i * Group_i + \alpha_6 Treat_i * After_i * Group_i + \alpha_4 X_i + \epsilon_i \end{aligned} \tag{4}$$

The effect of ML reform on family living arrangements based on gender and birth order is presented in Table 8. Previous studies have shown that women with daughters are less likely to marry (Dahl et al. (2008)), and that daughters are less likely than sons to live with their fathers (Lundberg et al. (2007)). The results in Panel A of Table 8 show no significant differential effect based on the child's gender. Among the families with daughters, the probability of single motherhood is 1.1 percentage points higher for women eligible for ML than those ineligible, accompanied by a 0.8 percentage point decrease in the likelihood of being married.

Panel B of Table 8 shows the heterogeneity effects of being the firstborn child. One would expect to have stronger results for the firstborn child because changes in family living arrangements are less likely to occur when families have more children. The standard errors are large compared to the magnitude of the coefficients; thus, I cannot reject the hypothesis that maternity leave has a stronger effect for the firstborn child.

In Table 9, I perform various sensitivity and robustness checks to confirm the main results. First, I perform the main regression without controls, and then I vary the window used before and after the reform to check for potential seasonality effects. I use a multinomial logistic regression with differences-in-differences estimation to account for the fact the outcomes are mutually exclusive. The estimates from these regressions are similar to those in the main results. Second, I test for two placebo reforms (i.e., one year before and one year after reform implementation date), and the estimates are statistically insignificant.

The two maternity leave impacts on family living arrangements should not be compared because these reforms differ in many aspects. The 1997 ML reform offered mothers 85% of their previous income for two years, while the 1990 reform offered mothers 65% of their income before childbirth for only one year. The 1997 ML reform effects on family living arrangements are probably generated by the larger increases in the amount of money mothers received during the leave compared to the previous reform. Another difference in my analysis is that I observe family living arrangements two years after the 1990 reform and five years after the 1997 reform.

7 Conclusion

I investigate the causal effect of two different maternity leave policies in Romania. The first reform, which extended ML to one year and provided mothers with 65% of pre-childbirth income, was introduced in 1990; the second reform, which extended ML to two years and provided mothers with 85% of income before childbirth, was introduced in 1997. To identify the effect of the policies on family living arrangements, I use regression discontinuity and differences-in-differences and Romanian census data from 1992 and 2002.

In this study, I investigate the effect of paid maternity leave in an understudied area: family living arrangements. The empirical research results suggest that the 1990 maternity leave policy (one year) had no significant effect on family living arrangements. However, the 1997 maternity leave policy (two years) led to a 1.2 percentage point increase in the probability of single motherhood, which supports the economic independence hypothesis and improvement in women's bargaining power.

These results are relevant to policymakers working on paid maternity leave policies. Maternity leave's main goal is to offer employment and income security after childbirth, promoting gender equality and women's attachment to the labor force. Nevertheless, it is crucial to take into consideration its effect on family living arrangements. For instance, maternity leave may increase women's labor force participation due to increased single motherhood probability, making the household worse-off. To assess if the net effect of a policy is positive, policymakers need research on various outcomes. As my results show, a shorter period of maternity leave does not affect family living arrangements, while more extended maternity leave increases the probability of single motherhood.

Table 1: Descriptive Statistics. 1992 Romanian Census

Variables	Mean	St. Dev.	Min	Max
A. Living arrangements				
Grandparent in household	0.353	0.478	0	1
Married couple	0.954	0.210	0	1
Single mother	0.029	0.168	0	1
Divorced	0.017	0.129	0	1
B. Child's characteristics				
Frequency of being male	0.504	-	0	1
Frequency of being first born	0.355	-	0	1
C. Mother's characteristics				
Age	27.304	6.064	14	53
Frequency of education degree				
High School or less	0.962	-	0	1
College	0.038	-	0	1
Observations		16,532		

Note: Data represent 15% sample from 1992 Romanian census. Sample is restricted to mothers who had a child two months before or after January 19^{th}, 1990.

Table 2: Descriptive Statistics. 2002 Romanian Census

Variables	Mean	St. Dev.	Min	Max
A. Living arrangements				
Grandparent in household	0.252	0.434	0	1
Married couple	0.797	0.402	0	1
Single mother	0.110	0.312	0	1
Divorced	0.039	0.195	0	1
B. Child's characteristics				
Frequency of being male	0.512	-	0	1
Frequency of being first born	0.507	-	0	1
C. Mother's characteristics				
Age	29.735	5.208	14	59
Frequency of education degree				
High School or less	0.944	-	0	1
College	0.056	-	0	1
Observations		20,777		

Note: Data represent 10% sample from 2002 Romanian census. Sample is restricted to mothers who had a child three months before or after July 1997 (July month being excluded).

Table 3: Covariates balance check

Variables	Coefficient	St.Error	P-value
A. Child's characteristics			
Male	0.010	(0.016)	0.532
First born	-0.005	(0.016)	0.737
B. Mother's characteristics			
Age	0.173	(0.205)	0.399
College educated	0.002	(0.006)	0.738
High Skill Occupation	0.004	(0.014)	0.754
Employment	0.003	(0.016)	0.871
Observations		16,532	

Note: The table presents estimates from the main equation: $Y_i = \beta_0 + \beta_1 T_i + \beta_2 f(d) + \beta_3 f(d) * T_i + \beta_4 X_i + \epsilon_i$, where $f(d)$ is the value of a linear polynomial in the running variable d (day). Each coefficient comes from a different regression. Robust standard errors are clustered at the day of birth. Significance levels: *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$

Table 4: Paid Maternity Leave (1990 reform) Effect on Families' Living Arrangements

	Married		Single		Divorced		Grandparents in HH	
	Linear	CCT	Linear	CCT	Linear	CCT	Linear	CCT
RRD Effect	0.009	0.001	-0.003	-0.008	-0.006	-0.004	0.004	0.036
	(0.007)	(0.013)	(0.006)	(0.013)	(0.004)	(0.006)	(0.015)	(0.037)
Bandwith		19.69		19.77		14.71		14.87
Observations	16,532	16,532	16,532	16,532	16,532	16,532	16,532	16,532

Note: Each coefficient comes from a different regression. The sample includes mothers who had a child two months before or after the policy was enacted. The column header indicates the outcomes and if the regression is a linear polynomial or a local linear polynomial. Controls not shown include mother's age and education, day of the week fixed effects corresponding to child's birth, gender, and indicator for multiple births. Robust standard errors are clustered at the day of the birth level. Data: 1992 Romanian census. Significance levels: *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$

Table 5: Heterogeneity by Child's Gender and Birth Order (1990 reform)

Dependent Variables	Married	Single	Divorced	Grandparents in HH
Panel A: Gender				
Treatment	0.012	-0.006	-0.005	0.001
	(0.011)	(0.009)	(0.007)	(0.020)
Treatment * Female	-0.005	0.005	-0.001	0.006
	(0.015)	(0.010)	(0.010)	(0.029)
Observations	16,532	16,532	16,532	16,456
Panel B: Birth Order				
Treatment	0.022	0.002	0.000	-0.002
	(0.019)	(0.007)	(0.006)	(0.004)
Treatment * Firstborn Child	0.024*	-0.01	-0.012	-0.0589*
	(0.013)	(0.011)	(0.009)	(0.032)
Observations	16,532	16,532	16,532	16,456

Note: Each coefficient comes from a different regression. The sample includes mothers who had a child two months before or after the policy was enacted. The column header indicates the outcomes. Controls not shown include mother's age and education, day of the week fixed effects corresponding to child's birthdate, gender, and indicator for multiple births. The interaction terms between the subgroup indicator (female or firstborn child) with the running variable (day of birth), and the interaction between treatment and running variables are not reported. Robust standard errors are clustered at the day of the birth level. Data: 1992 Romanian census.
Significance levels: *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$

Table 6: Robustness checks for the 1990 maternity leave reform

	Married	Single	Divorced	Grandparents in HH
R1: baseline (N=16,532)	0.009	-0.003	-0.006	0.004
	(0.007)	(0.006)	(0.004)	(0.015)
R2: RD 30 days window (N=8,235)	0.009	-0.009	-0.004	0.003
	(0.009)	(0.007)	(0.005)	(0.020)
R3: RD 2nd order polynomial (N=16,532)	-0.001	-0.005	0.006	0.005
	(0.010)	(0.008)	(0.005)	(0.023)
R4: RD placebo 1989 (N=15,958)	0.001	-0.001	-0.001	0.016
	(0.004)	(0.004)	(0.001)	(0.012)
R5: RD placebo 1991 (N=11,504)	0.005	-0.003	-0.002	0.015
	(0.005)	(0.004)	(0.002)	(0.018)
R6: DID 2 months (N=33,221)	-0.003	0.004	-0.001	0.007
	(0.005)	(0.004)	(0.003)	(0.010)
R7: DID 3 months (N=50,193)	0.001	0.003	-0.003	0.007
	(0.004)	(0.003)	(0.002)	(0.008)
R8: DID 4 months (N=58,931)	0.001	0.002	-0.003	0.003
	(0.004)	(0.003)	(0.002)	(0.008)

Note: Each coefficient comes from a different regression. The column header indicates the outcomes. Controls not shown include mother's age and education, day of the week fixed effects corresponding to child's birthdate, gender, and indicator for multiple births. Robust standard errors are clustered at the day of the birth level. Data: 1992 Romanian census. Significance levels: *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$

Table 7: Paid Maternity Leave (1997 reform) Effect on Families' Living Arrangements

Dependent Variable	Married	Single	Divorced	Grandparents in HH
Treatment	0.004**	-0.001	-0.002	-0.003
	(0.001)	(0.001)	(0.002)	(0.004)
After	0.001	0.003*	0.003	0.007
	(0.002)	(0.002)	(0.002)	(0.008)
Treatment * After	-0.011***	0.012***	-0.003	-0.012*
	(0.003)	(0.002)	(0.003)	(0.006)
Observations	20,777	20,777	20,777	20,777

Each coefficient comes from a different regression. The sample includes mothers who had a child three months before or after July 1997 (July month is excluded). Controls include mother's age, an indicator for the maternal level of education, dummies for multiple births and child's sex, and month fixed effects corresponding to the child's birth. Data: 2002 Romanian census. Significance levels: *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$

Table 8: Heterogeneity by Child's Gender and Birth Order (1997 reform)

Dependent Variables	Married	Single	Divorced	Grandparents in HH
Panel A: Gender				
Treatment * After	-0.007	0.007	-0.002	-0.003
	(0.005)	(0.007)	(0.006)	(0.008)
Treatment * After * Female	-0.008	0.011	-0.002	-0.019
	(0.009)	(0.012)	(0.013)	(0.020)
Observations	20,777	20,777	20,777	20,777
Panel B: Birth Order				
Treatment * After	-0.001*	0.012**	-0.005	-0.025***
	(0.005)	(0.005)	(0.004)	(0.006)
Treatment * After * Firstborn Child	-0.002	0.000	0.002	0.020**
	(0.017)	(0.016)	(0.006)	(0.009)
Observations	20,777	20,777	20,777	20,777

Each coefficient comes from a different regression. The sample includes mothers who had a child three months before or after July 1997 (July month is excluded). Panel A represents heterogeneity based on the child's sex, and panel B for being a first-born child. The interaction terms between subgroup indicators (female or first born) with *treatment*, and *after* variables are not reported. Controls include mother's age, an indicator for the maternal level of education, dummies for multiple births and child's sex, and month fixed effects corresponding to the child's birth. Data: 2002 Romanian census. Significance levels: *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$

Table 9: Robustness Checks for the Paid Maternity Leave (1997 reform)

	Married	Single	Divorced	Grandparents in HH
R1: baseline (N=20,777)	-0.011***	0.012***	-0.003	-0.012*
	(0.003)	(0.002)	(0.003)	(0.006)
R2: no controls (N=20,777)	-0.019*	0.015*	-0.004	-0.011
	(0.009)	(0.008)	(0.004)	(0.008)
R3: DID 4 months (N=26,945)	-0.008***	0.009***	-0.003	-0.0120**
	(0.002)	(0.003)	(0.003)	(0.005)
R4: DID 5 months (N=33,120)	-0.008***	0.007**	-0.002	-0.009**
	(0.002)	(0.003)	(0.002)	(0.004)
R5: Multinomial logit (N=20,777)	-0.008**	0.012***	-0.004	-
	(0.004)	(0.004)	(0.003)	
R6: placebo reform 1996 (N=17,208)	0.001	-0.004	0.006	-0.020
	(0.014)	(0.011)	(0.008)	(0.016)
R7: placebo reform 1998 (N=17,464)	0.003	-0.006	0.001	0.015
	(0.014)	(0.012)	(0.007)	(0.017)

Note: Each coefficient comes from a different regression. Controls include mother's age, an indicator for the maternal level of education, dummies for multiple births and child's sex, and month fixed effects corresponding to the child's birth. Data: 2002 Romanian census. Significance levels: *** $p < 0.01$ ** $p < 0.05$ * $p < 0.1$

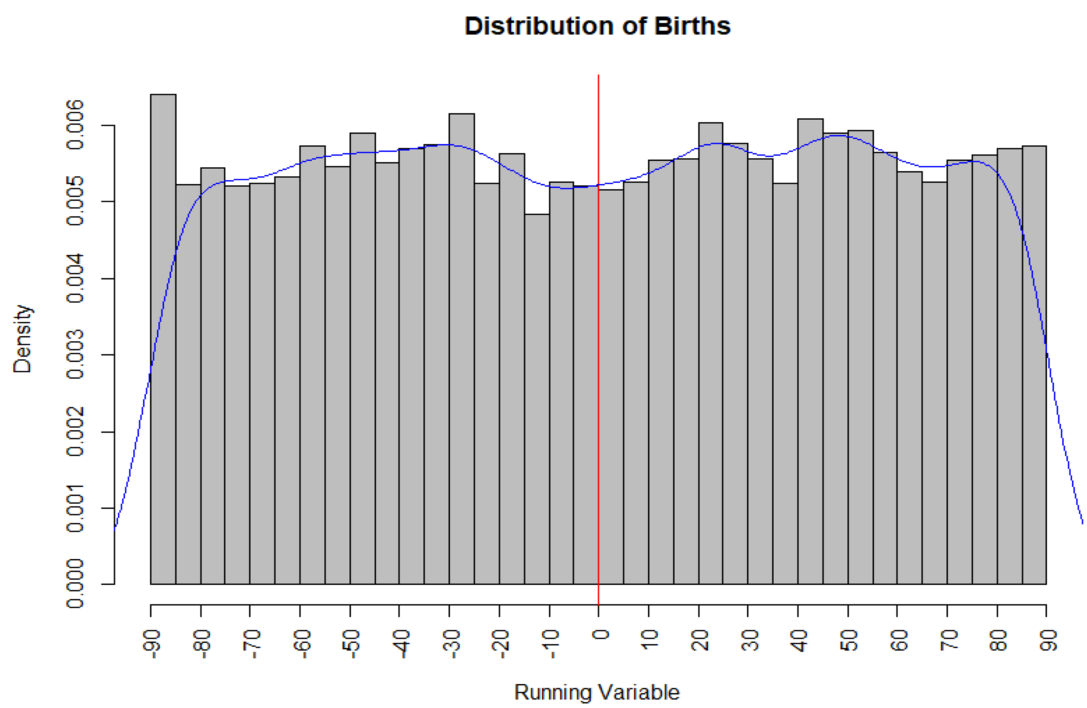

Figure 1: Histogram of number of births

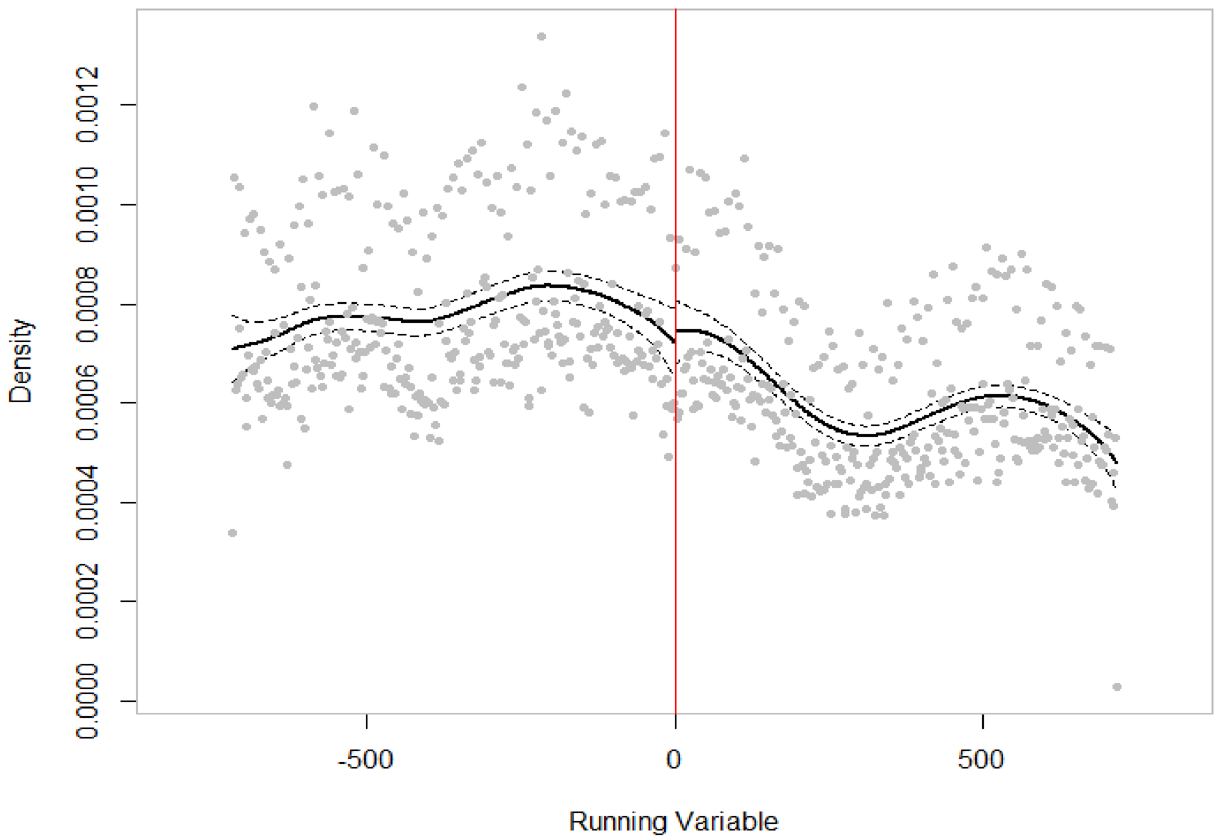

Figure 2: Local Linear Estimator- McCrary (2008). It computes equally spaced bins of the running variable (day of birth) and frequency counts are calculated within those bins. Then, these frequency counts are used as dependent variable and midpoint of the bins as independent variable, in a local linear smoothing of the histogram, which is conducted separately on each side of the threshold. Each gray dot represents the number of births in a bin, and the solid black lines show the two local linear smoothing of the histogram with the corresponding 95% confidence intervals.

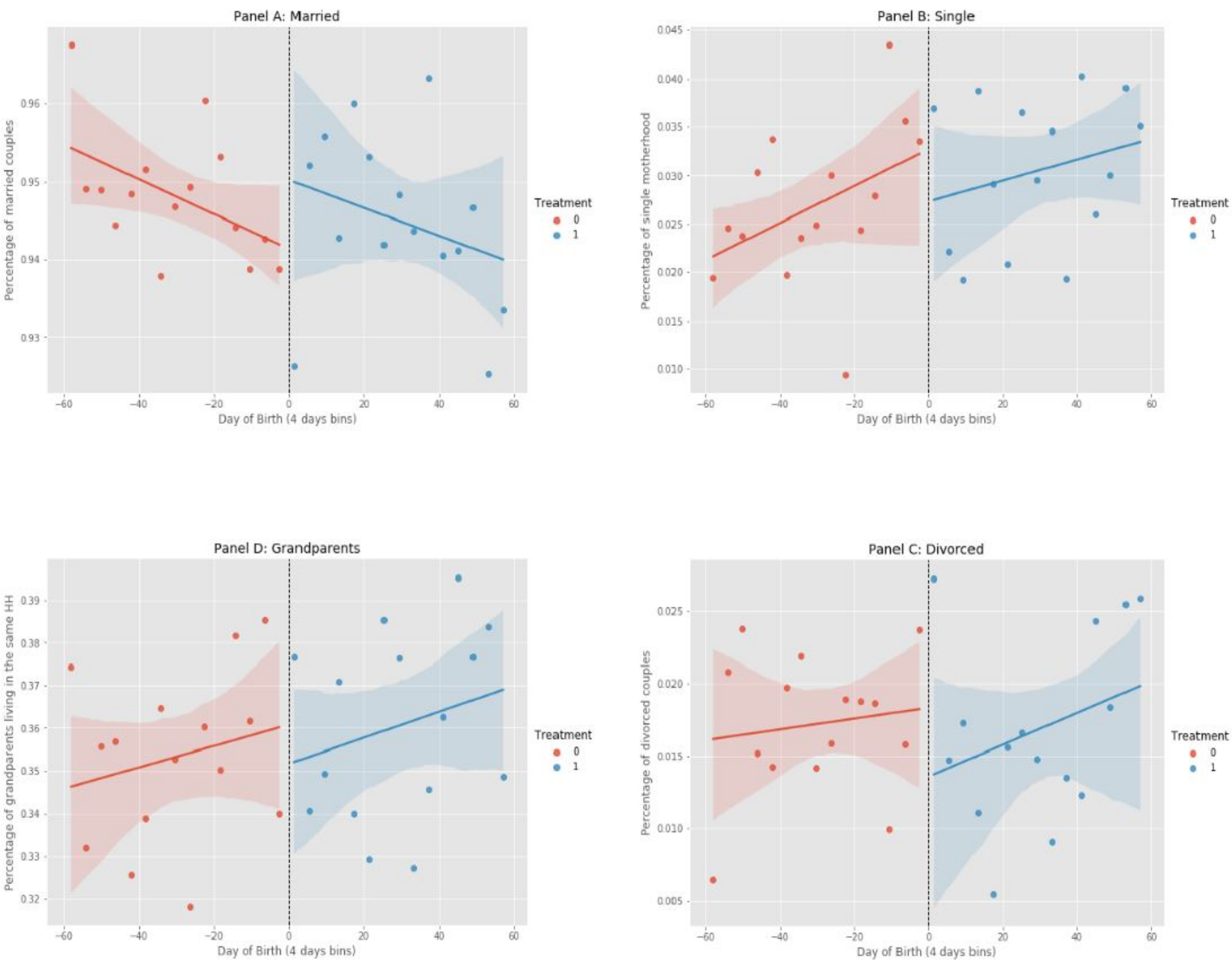

Figure 3: The effect of maternity leave on families' living arrangements

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