The inter-generational fertility effect of an abortion ban

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Abstract

This study examines the extent to which banning women from having abortions affected the fertility of their children, who did not face a similar legal constraint. Using multiple censuses from Romania, I follow men and women born around the time Romania banned abortion in the mid-1960s to investigate the demand for children over their life cycle. The empirical approach combines elements of regression discontinuity design and the Heckman selection model. The results indicate that individuals whose mothers were affected by the ban had significantly lower demand for children than those who were not. One-third of the decline is explained by inherited socio-economic status.

Keywords Intergenerational fertility transmission · Fertility preferences · Romania · Abortion ban

1 Introduction

In 2015, 42% of countries in the world had active policies aimed at reducing fertility rates, whereas 28% had implemented measures targeting the opposite result.[1] The high prevalence of population policies worldwide reflects the generally accepted view that a society's birth rate is a fundamental determinant of its economic well-being. A comprehensive understanding of fertility determinants and the socio-economic consequences of policies designed to influence them has enormous implications. However, most studies on birth control policies are limited to their contemporaneous effect, ignoring the consequences that such policies may have on the

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fertility of future generations. For example, Bailey (2010) examined how the birth control pill accelerated the post-1960 US fertility decline. Miller (2010) analyzed the role that *ProFamilia*, a program that provided IUD devices to married women, had on the Colombian demographic transition, Gertler and Molyneaux (1994) assessed the contribution of contraceptives to the Indonesian fertility decline during the 1980s, and Levine et al. (1999) studied the impact of abortion legalization on fertility rates in the USA. Nonetheless, a few papers in the literature focus on the long-term consequences of abortion policies (Ananat et al. 2009; Donohue and Levitt 2001).

The current paper complements this literature by presenting evidence that population policies can have long-lasting effects extending beyond one generation. Specifically, this paper shows that the abortion ban implemented in Romania in 1966 not only affected the fertility of women who were directly constrained by the law but also shaped the next generation's demand for children long after the abortion ban had ceased. Additionally, the paper uses this natural experiment to explore the mechanisms through which the fertility behaviors of parents pass on to their children.

Measuring the intergenerational transmission of fertility and its determinants is essential to understanding several dynamic aspects of a population and its material well-being. For example, couples with many children may invest relatively little per child (Becker and Lewis 1973), thus negatively affecting the future living standards of their offspring. Moreover, if children inherit the reproductive behavior of their parents, then the decline in living standards is likely exacerbated. Notably, the empirical literature on economic mobility has widely ignored fertility as a transmission channel. For example, a review by Black and Devereux (2010) does not mention any papers on this topic. Additionally, the persistence of fertility across generations may shed light on the patterns and speed of demographic transitions.

In 1966, Romania banned abortion and other forms of fertility control methods (Dethier et al. 1994; Pop-Eleches 2006). The abrupt implementation of this policy surprised a group of pregnant women who would have opted for abortion had the anti-abortion decree not existed. These women and their partners had different characteristics compared with couples who would not have aborted even if they have had the chance to do it. Thus, the abortion ban significantly changed the composition of families having children and, through this phenomenon, the background of second-generation individuals.[2]

The role played by background characteristics in fertility can be measured by comparing the reproductive behavior of second-generation men and women born around the time the policy was implemented. This approach has strong similarities with regression discontinuity design (RDD), where the running variable is the date of conception and the cutoff is the moment when the anti-abortion decree was enforced. However, it is conceptually different. Whereas first-generation pregnant women are plausibly similar around the policy cutoff, the abortion ban substantially distorted the selection into childbearing.

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[2]In this paper, the first generation refers to the individuals directly affected by the 1966 policy, while the second generation comprises their children. The meaning of these terms should not be confused with those used in the migration literature.

First-generation individuals likely affect the fertility of the second generation through various channels. One of them is socio-economic status. I use the 1977 Romanian census data (IPUMS-International 2015) to statistically condition the analysis on observable characteristics of first-generation fathers and mothers. The data contain a rich set of variables, including housing characteristics and locality of residence as well as the education, industry, and occupation of each household member. Although assuring that these variables are a sufficient statistic is not possible, it is worth noting that Romania was under a socialist regime. Little dispersion, if any, in living standards is expected after controlling for the observables just mentioned. The 1960 Romanian mechanisms to allocate labor and determine wages described in Section 4 support this statement.

Apart from changing the composition of families, the abortion ban may have directly affected second-generation individuals by altering their preferences and the nurturing behavior of their parents. For example, "unwanted" children may be less loved and neglected. I discuss these and other mechanisms in the paper.

The data used in this paper are the 1977, 1992, 2002, and 2011 Romanian censuses. Joint analyses of different datasets permit the study of second-generation individuals born around the policy cutoff at different points of their reproductive life (9, 24, 34, and 44 years old, respectively). Results indicate that individuals whose mothers were affected by the ban had significantly lower demand for children. Inherited socio-economic status explains one-third of this decline. The results are robust to different econometric specifications and across datasets that span women's entire reproductive life. The anti-abortion policy also appears to affect second-generation men as evidenced in their reproductive behavior.

This paper builds on previous work by Pop-Eleches (2006), which examines the impact of the abortion ban on the socio-economic outcomes of the second generation, and that of Pop-Eleches (2010), which examines the direct effect of the Romanian pro-natalist regime and its removal on the fertility of the first generation. This paper complements those studies by examining the fertility behavior of second-generation individuals born around the time of the abortion ban. Additionally, I present a new method that combines the Heckman (1979) sample selection model with RDD.

This paper directly relates to at least three areas of research. First, it helps understand the impact and effectiveness of fertility policies. As previously mentioned, the bulk of studies focus on the fertility impact that contraceptives potentially have on the generation of individuals directly affected by them Bailey (2010), Miller (2010), Ngo (2020), Gertler and Molyneaux (1994), and Pop-Eleches (2010). Papers that analyze the effect of contraceptives on future generations typically target different outcomes such as children's health (Joshi and Schultz 2013), their socio-economic status (Bailey 2013; Pop-Eleches 2006), and their propensity to commit crimes (Donohue and Levitt 2001). Only Ananat and Hungerman (2012) discussed the impact of the contraceptive pill on the fertility of second-generation individuals in the USA. In that case, the pill had negligible long-term effects on the demand for children.

Second, the current paper is related to the literature that focuses on the statistical association of fertility across successive generations (Danziger and Neuman 1989; Murphy 1999; Kolk 2014; Murphy 2012; Murphy and Knudsen 2002). Although

these papers provide valuable information, they rarely attempt to disentangle the role played by inherited wealth from other determinants.[3]

Third, this paper contributes to the *cultural transmission* literature by studying the intergenerational transmission of attitudes toward childbearing. Papers in this literature separately identify preferences from other fertility determinants by analyzing the behavior of US immigrants (Fernandez and Fogli 2009; Blau et al. 2013; Guinnane et al. 2006). Considering that preferences, but not economic and institutional constraints, are portable when people migrate, the statistical association between the fertility of second-generation immigrants and the fertility rate in the country where their parents were born is arguably explained exclusively by culture.

The "cultural transmission" literature uses an ingenious identification strategy and finds relevant results. However, it also has some limitations. First, migrants are likely not representative of the population in the country of origin. Thus, how much of the observed correlation is due to cultural transmission and that of migrants' self-selection is unclear. Second, most of these studies use second-generation migrants.[4] Then, the impact of the country of origin culture is diminished by the degree of assimilation to US culture. Third, given that the identification of the cultural transmission of preference relies on comparing immigrants from different countries of origin, then, the influence of the socio-economic status of immigrants' parents cannot be eliminated. It requires the difficult task of making cross-country comparisons of standards of living.[5] Despite the limitations that may bias the magnitude of the relationship of interest, the hypothesis that culture affects economic outcomes is credibly tested in this literature.

The rest of the paper is organized as follows. Section 2 describes the potential mechanisms driving the intergenerational effect of the abortion ban. Section 3 describes the anti-abortion policy implemented in Romania in 1966. Section 4 explains how the central government determined the allocation of labor and wages and suggests that the variables available in Censuses are plausibly sufficient statistics for the socio-economic status of households. Section 5 presents a theoretical framework that guides the empirical methodology. Section 6 presents the data. Sections 7 and 8 show fertility patterns across generations and describe the characteristics of first-generation women who self-selected to abort their pregnancies in light of the model from a previous section. Section 9 empirically assesses intergenerational fertility transmission. Section 10 analyzes the demand for children of second-generation men. Given that the total number of children ever fathered is unavailable, the section relies on the number of own children living with the father and the children ever born to the wives of men living with their spouses. Finally, Section 11 summarizes and concludes.

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[3]Some studies attempt to control for parent's economic conditions (e.g., Danziger and Neuman (1989)), but the information used is scarce and likely insufficient to isolate preferences from confounders.

[4]Second-generation migrants are defined here as people born and raised in the USA with foreign-born parents.

[5]An exception is Blau et al. (2013) who analyzed the correlation of outcomes between first and second generations of immigrants in the USA. However, the paper does not attempt to separate the influence of inherited preferences from inherited wealth.

2 Why banning first-generation individuals from aborting changes the fertility of the second generation

Theoretical papers emphasize that economic constraints and preferences determine the demand for children (Becker and Lewis 1973; Willis 1973). Prohibiting the first generation (i.e., the parents) from aborting may affect the distribution of these two determinants among second-generation individuals (i.e., the children) through a variety of channels.

The first of such channels is selection, or stated differently, changes in the composition of families having children. Women who decided to terminate their pregnancies before the issuance of the decree that banned abortion were likely to have a relatively low "taste" for children and possibly a different socio-economic status than women who carried their pregnancies to term. Thus, banning abortion plausibly increased the proportion of individuals in the second generation whose parents had preferences for small families and changed the average economic conditions of birth cohorts. Both wealth and preferences may pass from parents to children, affecting the average fertility of the second generation.

Second, banning abortion may affect the optimal quantity of children and the timing of births. In both cases, the resources allocated to each child may change. As Becker and Lewis (1973) and Willis (1973) indicated in their quantity-quality trade-off fertility theory, larger families likely allocate fewer resources per child than smaller ones. Additionally, young parents tend to have lower incomes than their older counterparts.

Third, if banning abortion substantially increases the size of a cohort, as is this case, then, a crowding effect may occur (Pop-Eleches 2006). Some public resources, such as schools, may have experienced congestion, compromising the formation of human capital.

Identifying the crowding effect is not possible with the available data. Thus, this paper inevitably disregards the general equilibrium effect associated with the potential congestion of public goods, as is the case in other studies in the literature.

Finally, children born after unplanned pregnancies may receive less care and have worse outcomes, thereby modifying their propensity to have children. Although no strong empirical evidence exists to support this statement (see Section 9.4), the model in Section 5 contemplates such possibility.

3 Romanian fertility policies and intuition of the research design

On September 25, 1957, the government of Romania issued Decree 463, allowing women to abort during their first trimester of pregnancy (Dethier et al. 1994). Abortion centers began operating in medium and large hospitals as well as in outpatient facilities close to industrial plants (David and Wright 1971). In most cases, trained physicians performed abortions within a week of women's requests. The intervention fee was less than three US dollars.

Abortion slowly became a socially acceptable birth control method in Romania. In 1958, a total of 112,068 legal abortions were reported in this country (David and

Wright 1971). By the mid-1960s, the number of abortions reached 1.1 million per year (Horga et al. 2013). In 1966, the last year of Decree 463, the total fertility rate (TFR) was 1.88 (below replacement level), and the crude birth rate (CBR) was 14.1, the lowest in the world (Dethier et al. 1994). Overall, population growth fell by fifty percent during the period 1957–1966, reaching 0.61% per year.

Concerned about low fertility levels, the government of Romania issued Decree 770 in October 1966. The new policy banned abortion and established severe penalties for women and physicians who violated the official mandate. Simultaneously, the Romanian leader Nicolae Ceauescu issued a series of population policies. He ended the importation of contraceptives, reduced the income tax for families with three or more children by 30%, and imposed a "childlessness" tax for men and women over the age of twenty-six.

The impact of these pro-birth policies became evident within a year. The crude birth rate increased from 14.1 in 1966 to 27.4 in 1967. In the same period, the TFR jumped from 1.88 to 3.7, and the population growth rate accelerated from 0.61% to 1.8% (Dethier et al. 1994; Pop-Eleches 2006; 2010).

Figure 1 shows the number of people by birth cohort in a 10% random sample extracted from the 1977 Romanian Census (IPUMS-International 2015). The cohort size increased to more than 150% between February and July 1967, approximately 9 months after the decree was enforced.

After reaching a peak in the third trimester of 1967, fertility started falling. By 1983, the crude birth rate reached 14.1, the same level it had in 1966, before the implementation of the anti-abortion decree. Nonetheless, part of this decline was the result of a relatively small cohort of women in their reproductive age. The TFR also declined in Romania. However, it never returned to the pre-policy levels during Ceauescu's leadership (Dethier et al. 1994).
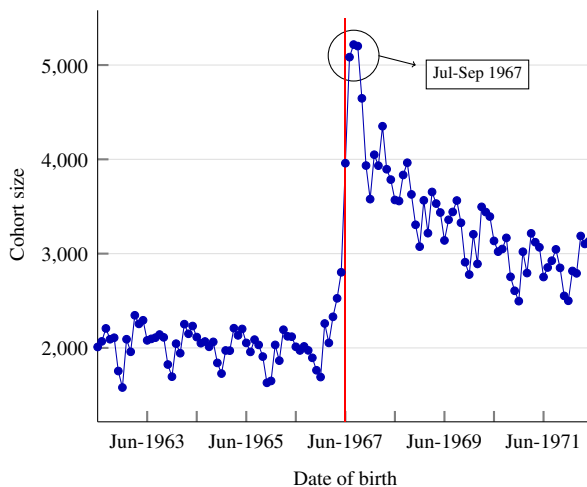

Fig. 1 Cohort size in 1977 Romanian census (by month of birth)

In December 1989, the communist regime ended. Within weeks, abortion became legal once again in Romania. In 1990, the first year of the policy change, one million abortions were reported (Pop-Eleches 2010). Modern contraceptives became widely available through international donations. By the mid-1990s, family planning clinics and trained physicians were accessible to most women (Horga et al. 2013).

After the initial spike in 1990, abortion rates started to decline in Romania. However, the TFR reached stable levels suggesting that women began using other birth control methods. In 2007, the contraceptive prevalence was above 70%. Modern contraceptives accounted for 61.1%. The most frequently used methods were condoms (27.6%), pills (22.8%), and intrauterine devices (7.0%) (Horga et al. 2013).

Policy change and empirical approach The empirical strategy in this paper involves a comparison of the reproductive behavior of second-generation individuals born before and after the policy cutoff observed in the summer of 1967, 9 months after Decree 770 was enforced (see Fig. 1). These two groups of individuals faced the same institutional and aggregate economic constraints over their lives. However, on average, they were born and raised in families with different preferences and socio-economic characteristics.

If the policy implementation was truly unanticipated, then, a group of women were already pregnant when the decree was enforced. These women, particularly those who were in their first trimester of pregnancy, had no option to abort and had not previously foreseen when they became pregnant and thus prevented them from terminating their pregnancies. Thus, the selection into pregnancy was identical for women who gave birth around the policy cutoff, but the selection into childbearing discontinuously changed because of the abortion ban. Assuming that the policy was unanticipated simplifies the interpretation. However, violation of this assumption does not invalidate the analysis. Below, I discuss the empirical consequences.

The pro-birth policy implemented in 1966 had a limited direct impact on second-generation individuals. It persisted until 1989 when the Romanian socialist regime ended. People born in 1967 (i.e., sons and daughters of first-generation women who became pregnant around the implementation of the anti-abortion decree) were 22 years old, i.e., they were early in their reproductive life when they gained full availability of contraceptive methods.

Moreover, the direct impact of the abortion ban implemented in 1966 on the fertility of second-generation individuals is identical around the policy cutoff. Thus, the empirical strategy in this paper identifies the indirect impact of the anti-abortion policy on the reproductive behavior of second-generation individuals through changes in the composition of families and parents' attitudes.

4 Labor allocation and wage determination in centrally planned Romania

In the 1960s, Romania had a centrally planned economy. Anyone willing to work could obtain a job provided by the government. Workers were not involved in any job search and had no freedom to choose the industry, occupation, or geographical

location. The Romanian system guaranteed stable employment and income for 40 years to most of the working-age population. The government allocated workers to enterprises at a young age. In some cases, this assignment was provided to persons aged 14 and still attending school (Dethier et al. 1994).

The absence of a decentralized labor market totally disabled any process of wage negotiation at the individual level. The central government specified a set of rules to determine workers' pay. These rules classified workers into pay categories defined by easily observable characteristics, such as industry, education, and experience. Notably, differences in individual performance or productivity were not components of wages.

The empirical analysis in Section 9 estimates the extent to which the observable socio-economic characteristics of first-generation parents explain the fertility behavior of second-generation individuals. The dataset used in this paper contains no information about individual earnings. However, it includes comprehensive details regarding the age, education, industry, sector, and occupation of each household member. Given the centrally planned labor system just described, these variables are plausibly sufficient to account for income differences in Romania during the 1960s. Moreover, they likely approximate permanent living standards better than current income. The data also provide comprehensive information on dwelling characteristics, including location, thus completing an accurate measure of the socio-economic status of the household.[6]

5 Conceptual framework

This section jointly models fertility decisions and resource allocation to guide the empirical strategy. The following Bellman's equation represents a woman's intertemporal decision process. The role of other household members is non-essential and hence ignored.

$$
V(n, A, D) = \max_{\substack{x, q \in R^+ \\ a \in \{0,1\} \\ \tilde{A} \in R}} \left\{ u(x, q, n; D, \xi) + \beta E \left[a\, V(n, \tilde{A}, \tilde{D}) + (1-a)\, V(n+1, \tilde{A}, \tilde{D}) \right] \right\}
\tag{1}
$$

The left-hand side of Eq. 1 is the value function, which depends on demographic characteristics D (e.g., age and ethnicity), the number of children n she has, and a stock of assets A (i.e., the state variables). The right-hand side of Eq. 1 indicates that her choices affect both the current utility $u(.)$ and the expectation of future utilities $E[.]$. The parameter $\beta \in [0, 1)$ is a time discount factor. In the current period, this woman optimally chooses her consumption level x, the assets $\tilde{A}$ she wants to keep

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[6]Wage differential across workers' categories was low. In 1989, before the revolution, earnings of "specialists" were only 14 percent higher than those of "regular workers" (Dethier et al. 1994). Then, a possibly reasonable assumption is as follows: accounting for the variables that determined workers' categories eliminates any significant dispersion in workers' compensation.

for the next period, and the consumption level for each of her children q. The variable q is the average investment in children's quality.

Apart from x and q, this woman derives utility $u(.)$ from her children n. She can increase the size of her family in future periods by getting pregnant and not having an abortion ($a = 0$). If she implements a family planning method ($a = 1$), the number of children remains unchanged. The preference parameter ξ partially determines all choices. For simplicity, I call this parameter the "taste" for children. In the utility function $u(.)$, demographic characteristics D operate as taste shifters.

The next period's expected value function $V(.)$ (i.e., the second term on the right-hand side) depends on future demographics $\tilde{D}$ (i.e., one year older), future assets $\tilde{A}$, and the number of children resulting from family planning decisions.

$$p_x x + p_q qn + \tilde{A} + p_a a = y(n) + A(1 + r) \tag{2}$$

The right-hand side of the woman's budget constraint (2) is the sum of capitalized assets $A(1+r)$, where r is the interest rate and a stochastic labor income $y(n)$ realized at the beginning of the period. The number of children n may influence the moments of the labor income distribution. For example, women with many children may accumulate relatively less human capital or work part time, resulting in below-average earnings.

The left-hand side of the budget constraint shows that the representative woman purchases a consumption bundle x in which the price index is p_x, goods for her children at a price p_q, assets for the next period $\tilde{A}$, and family planning procedures in which price is p_a in the applicable case.[7] The interaction term qn, which generates non-linearities in the budget constraint, is standard in the fertility literature (Becker and Lewis 1973). Thus, increasing the average quantity of goods consumed by children in one unit requires that the mother purchases n goods (e.g., one pair of shoes for each child).

Abortion decisions around 1966 (first-generation women) The problem described in Eqs. 1 and 2 can be used to explain women's propensity to abort in 1966 before the implementation of the Romanian pro-natalist policy.

Computing the inter-temporal utility maximization (1)–(2) is considered in two stages. First, women optimize with respect to x, q, and $\tilde{A}$ for each value of a to obtain conditional indirect utility functions. The difference in these two conditional indirect utilities defined as $a* \equiv V(.|a = 1) - V(.|a = 0)$, is a function of state variables n^f, A^f, D^f, realized labor income $y(n)^f$, and preferences ξ^f as Eq. 3 shows. The superscript f indicates that the variables correspond to first-generation women. The second stage in the maximization process, Eq. 4, takes the resulting difference in

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[7]Although the health system covered the direct cost of abortions during the socialist regime, women may have faced other costs such as transportation and the opportunity cost of time. The magnitude of p_a is irrelevant for the empirical analysis.

indirect utilities and indicates whether having an abortion is the optimal choice.

$$a* \equiv V(.|a=1) - V(.|a=0) = g(n^f, A^f, D^f, y(n)^f, \xi^f) \tag{3}$$

$$a = \begin{cases} 1 & \text{if } a* > 0, \\ 0 & \text{if } a* \leq 0. \end{cases} \tag{4}$$

$$where \quad \frac{\partial a*}{\partial n^f} > 0, \frac{\partial a*}{\partial A^f} \gtrless 0, \frac{\partial a*}{\partial D^f} > 0, \frac{\partial a*}{\partial y^f} \gtrless 0, \frac{\partial a*}{\partial \xi^f} < 0$$

Assuming standard properties of the per-period utility function $u(.)$, the partial derivatives of the propensity to abort, $g(.)$ in Eq. 3, indicate that older women and those who had previously given birth to relatively many children were likely to abort. The role of income and assets is theoretically ambiguous. The interaction term qn in the budget constraint, together with a relatively high income-elasticity of q, may create a negative association between the desired number of children and the socio-economic status (Becker and Lewis 1973; Willis 1973). Finally and more importantly, all else equal, women with a relatively low taste for children were unambiguously likely to abort.

The unexpected implementation of the pro-natalist decree in Romania forced many women, who would otherwise have had an abortion, to give birth. Thus, the usual selection process given by expressions (3) and (4) was suddenly interrupted. As a result and given the derivatives of the propensity to abort $g(.)$, the model predicts that first-generation women who gave birth after the policy cutoff should be older, with more children, with a relatively low taste for children, and probably a higher socio-economic status.

Equations 3 and 4 also indicate that conditioning on observables $X^f \equiv (n^f, A^f, D^f, y^f)$, the proportion of first-generation women who did not abort when it was legal was exclusively determined by their taste for children relative to a threshold $\bar{\xi}_x$ as indicated in Eq. 5.

$$\begin{aligned} P(a=0|X^f) &= P(a^* \leq 0|X^f) \\ &= P(g(X^f, \xi^f) \leq 0|X^f) \\ &= P(\xi^f > \bar{\xi}_x|X^f) \end{aligned} \tag{5}$$

The left-hand side of expression (5) is the conditional probability of not having an abortion when it was legal. The right-hand side of expression (5) results from inserting equality (3) and using the selection mechanism (4). Given that $g(X^f, \xi)$ in Eq. 3 is strictly negatively sloping in ξ^f (i.e., high fertility preferences constantly imply a lower propensity to abort), then, it can be inverted. Consequently, the threshold is given by $\bar{\xi}_x = g^{-1}(X^f)$.[8]

After the implementation of the pro-natalist policy, the probability of having an abortion ceased to be driven by Eq. 5, implying that the distribution of the taste for children abruptly changed among first-generation mothers. This discontinuity of the selection rule constitutes the basis of the empirical identification. An explanation follows below.

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[8]If the taste for children is additively separable $g(X^f, \xi) = \ddot{g}(X^f) - \xi$, then, the threshold is simply $\bar{\xi}_x = \ddot{g}(X^f)$ in Eq. 5.

Fertility decisions of women born around the anti-abortion decree (second-generation women) The inter-temporal fertility problem (1)–(2) also characterizes the behavior of second-generation women (i.e., the daughters of women who gave birth around the policy cutoff).

$$n(y_0^s, A_0^s, D^s, \xi^s, \mu) \tag{6}$$

Equation 6 is the demand for children. It indicates that the optimal number of children ever born to a second-generation woman at age D^s depends on her initial socio-economic status, given by income y_0^s and assets A_0^s at the beginning of her reproductive life and preferences for children ξ^s. Noticeably, only the socio-economic status in childhood is exogenous. The income and asset profiles over the adult years depend on the decisions made in each period, including fertility choices. Nonetheless, the empirical Section 9 analyzes adult socio-economic variables as potential mediators.

Expression (6) adds the variable μ not present in the model, to capture the potential psychological impact of being an "unwanted" child and any crowding effect on the demand for children (e.g., lower quality of education; see Section 2 for a discussion).

Define the vector of observable variables $X^s = (y_0^s, A_0^s, D^s)$ - where the superscript s denotes the second generation. Assume that the taste for children and the "unwantedness" effect μ to be additively separable. Then, Eq. 6 becomes

$$n = h(X^s) + \xi^s + \mu \tag{7}$$

After separately taking the conditional expectation of Eq. 7 for cohorts born before ($pre = 1$) and after ($pre = 0$) the policy cutoff marked in Fig. 1, the resulting average demands for children are as follows:

$$E(n|X^f, X^s, pre = 1) = h(X^s) + E(\xi^s|X^f, X^s, \xi^f > \tilde{\xi}_x) \tag{8}$$
$$E(n|X^f, X^s, pre = 0) \quad = h(X^s) + E(\xi^s|X^f, X^s) + \mu \tag{9}$$

Equation 7 indicates that the variation in the number of children ever born after conditioning on age and the initial socio-economic status of second-generation mothers (X^s) is given by the heterogeneity in preference ξ^s and the direct "unwantedness" effect μ. As shown in Eqs. 8 and 9, these effects are only relevant for individuals born after the policy implementation.[9]

Equations 8 and 9 also indicate that the anti-abortion decree likely changed the distribution of such preferences across birth cohorts. Before June 1967 ($pre = 1$), only the preferences ξ^s of second-generation women not aborted are in the data. As Eq. 5 indicates, women who were not aborted had mothers with relatively high taste for children ($\xi^f > \bar{\xi}_x$). By contrast, Eq. 9 indicates that the anti-abortion decree eliminated the possibility of abortion for all first-generation women, regardless of whether they had high or low fertility preferences. Thus, if full compliance of the anti-abortion decree occurred, then, $\xi^f > \bar{\xi}_x$ should be eliminated from the conditioning set in this second case.[10]

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[9]Unwanted children were reported before the abortion ban. However, what matters for the empirical analysis is the increase in such proportion. Assuming $\mu = 0$ in pre-policy periods is innocuous.

[10]The empirical sections discuss the case of incomplete compliance.

The difference between Eqs. 8 and 9 is driven entirely by $E(\xi^s|X^f, X^s, \xi^f > \tilde{\xi}_x) - E(\xi^s|X^f, X^s) - \mu$, which, if different than zero, is a sufficient condition to conclude that either preferences across generation are correlated ($corr(\xi^f, \xi^s) \neq 0$), and/or the direct impact of being unwanted (psychological and crowning effects) played a major role in the demand for children. The precise empirical approach is explained in Section 9.1 after presenting the data structure. Section 9.4 details the empirical strategy to estimate $corr(\xi^f, \xi^s) \neq 0$.

6 Data and descriptive statistics

The data used in this paper come from the Romanian Censuses carried out in the years 1977, 1992, 2002, and 2011. For each of these Censuses, IPUMS-International (2015) made microdata publicly available for a 10% random sample. Census 1977 excludes the residents of Alba and Arad counties.[11] For consistency across years, I drop people born in these two counties from other census samples. The sub-population of second-generation women and men in this study includes individuals born between 1962 and 1972, covering five years before and five years after 1967, which is the birth year of the first cohort who were affected by the anti-abortion policy.

Tables 6 and 7 in Appendix A present the descriptive statistics from the 1977 Census, when second-generation individuals born at the policy cutoff (i.e., June 1967) were 9 years old and living with their parents. This Census provides essential information about the socio-economic status of families needed to analyze the abortion decision of first-generation women and the fertility decisions of second-generation women at different points in their reproductive life (i.e., the information in vectors X^f and X^s in Section 5).

Using income or consumption to measure standards of living is not possible because these variables are absent in the Census data. However, a rich set of information exists about housing characteristics, which indicate asset usage as well as variables associated with the earning capacity of each household member such as education, employment status, industry, sector, and occupation. As discussed in Section 4, Romania was under a socialist regime in the 1960s and 1970s. Little dispersion in standards of living likely persists after conditioning on these and other variables such as age, county of residence, and whether the household resides in an urban area.

Table 8 in Appendix A shows the descriptive statistics for relevant variables in Censuses 1992, 2002, and 2011. I use these Census data to analyze the fertility choices of second-generation individuals at different points in their lives (ages 24, 34, and 44). I perform a separate analysis for men and women. The information available for women is more accurate than men. Only women report the number of children ever born to them and the age when they were first married. The data limitation for

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[11]Only 3.8% of the Romanian population reside in these counties.

second-generation men implies that the study of their reproductive behavior is some-how restrictive. I proxy the number of children that men ever had with the number of own children living in their house, and the children ever born to their female partner. I discuss the statistical consequences of this data limitation in Section 10.

7 Anti-abortion decree and the fertility of second-generation women

This section graphically analyzes the extent to which the fertility of second-generation women discontinuously changed across birth cohorts due to the anti-abortion decree issued in 1966. This pro-birth policy directly affected the parents of these women; however, since it ended in 1989, nearly no constraint was imposed on the availability of family planning methods for second-generation individuals.[12]

The upper portion of Fig. 2 shows the average children ever born to women in each month-year birth cohort. Each graph contains a cross-section of women ranging from 15 to 55 years old, covering all reproductive ages and beyond. The red vertical line cuts the horizontal axis in June 1967, the month when first-generation women, who were first affected by the pro-birth policy, gave birth. The shaded areas cover 11 years around June 1967, from January 1962 to December 1972. The cohorts born in these years form the sample included in the empirical analysis.

In 1992, second-generation women born in June 1967 were 24 years old and had, on average, one child. Figure 2a shows no apparent break around this policy cutoff. This fact may lead to the conclusion that the anti-abortion policy had no conse-quences on the fertility decisions of second-generation individuals. However, women in their mid-twenties are far from completing their reproductive cycle. Fertility dif-ferences between women born before and after June 1967 may not be evident at that age.

Figure 2b shows the cross-sectional life cycle fertility in 2002 when women born in June 1967 were 34 years old. Contrary to Census 1992, the trend break becomes evident. This break in the trend is substantially additionally apparent in Census 2011 when women born in 1967 approached the end of their reproductive life.

Figure 3 highlights the discontinuity in fertility trends around the cohort born in June 1967. The three graphs zoom in the same series plotted in Fig. 2 but only for women born between January 1962 and December 1972 (the shaded areas in Fig. 2).

The fertility patterns observed in Figs. 2 and 3 reveal that the anti-abortion policy that affected first-generation parents changed the reproductive behavior of second-generation individuals. However, these figures cannot tell the mechanisms underlying such association.

A large proportion of first-generation women who would have aborted in the absence of the decree gave birth in the third trimester of 1967. These women and their partners plausibly had a relatively low taste for children. The daughters of these

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[12]Second-generation individuals were 22 years old when the pro-birth policies ended (See Section 3). Although the abortion ban directly affected the initial years of their reproductive lives, the research design that exploits the discontinuity across birth cohorts is robust to this fact.
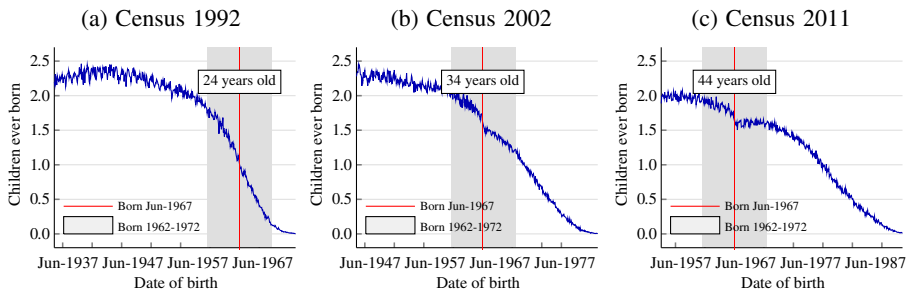
Fig. 2 Life cycle fertility, children ever born to women 15 to 55 years old by Census year

couples may have inherited or learned such preferences through socialization. Additionally, the nurturing of "unwanted" children may be different, thereby affecting their preferences for childbearing.

By contrast, couples who would have aborted had the abortion ban not existed, possibly belonged to a different socio-economic status. As long as the economic condition is inherited and the wealth elasticity of the demand for children is not zero, the inheritance of wealth may be part of the explanation underneath the patterns in Figs. 2 and 3.

The inheritance of wealth competes with the intergenerational transmission of preferences and the direct psychological effect of being "unwanted" to explain the decline in the average children born to second-generation women born after June 1967.

8 Anti-abortion decree and the self-selection of first-generation women into motherhood

The anti-abortion policy implemented by the government of Romania changed the usual mechanism through which women self-selected into motherhood. Women who would have opted to abort in the last trimester of 1966 had no choice but to keep the baby and give birth by mid-1967.
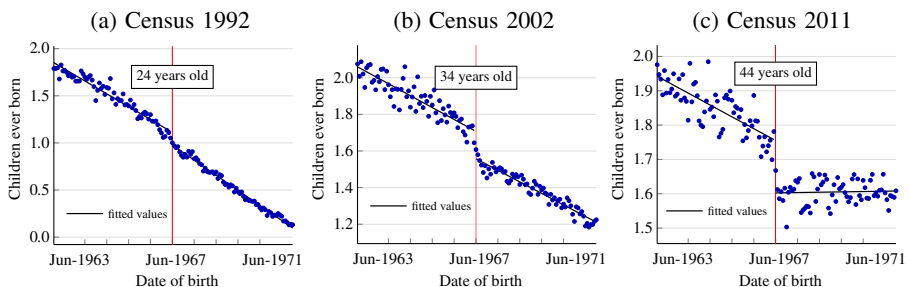
Fig. 3 Children ever born to women in cohorts Jan-1962 to Dec-1972 by Census year

Equation 3 indicates how each relevant variable affects the likelihood of interrupting a pregnancy when legal. According to this relationship, older women with relatively many children and possibly higher socio-economic status were likely to abort. Given that the pro-birth policy "forced" these women to give birth, then, the mothers of children born by mid-1967 are plausibly different in these dimensions.

Figure 4a shows the average age of the mother as a function of her child's date of birth. Consistent with the model in Section 5, mothers who gave birth after June

Fig. 4 Selected covariates (Census 1977)

1967 were 1 year older. Figure 4b shows that the child's average number of older siblings is higher for children born after the policy implementation. Although this figure is consistent with the theoretical model, the magnitude of the "jump" seems small. A plausible explanation is that the predictions of the model in Eq. 3 are conditional on 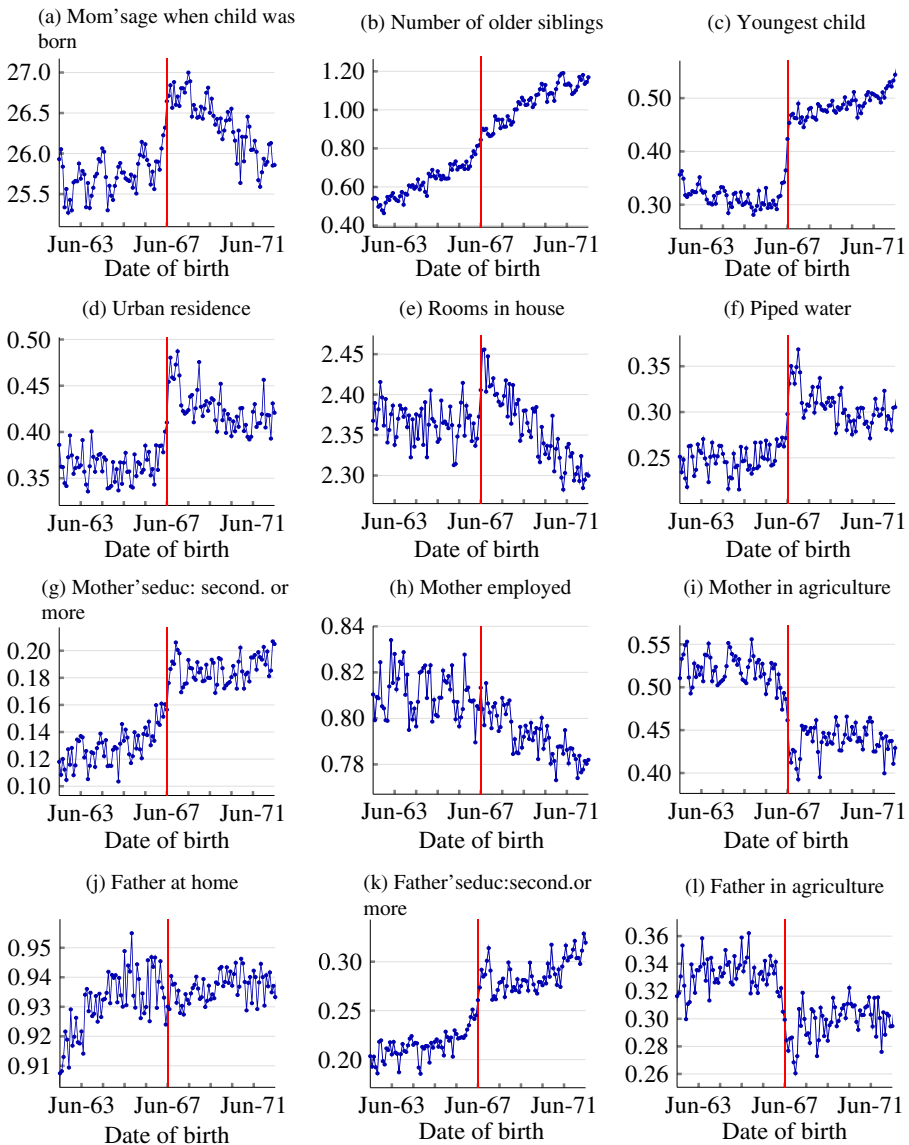other determinants to be fixed (i.e., a *ceteris paribus* analysis obtained from derivatives), but Fig. 4b shows the statistically unconditional relationship. If this assumption is the case, a better indicator of completed fertility is whether women decided to have more children after the policy shock.

Figure 4c shows the probability that the child was the youngest in the household in 1977. The "jump" is clear; 30% of kids born before the decree were the youngest at home but 46% after it, indicating that at least 35% of these children were born to mothers who had earlier reached the optimal number of children or were considering having only one extra child.[13]

The variables associated with the socio-economic status of women are highly relevant, not only because they plausibly have a significant role in the abortion decision of expecting mothers, but also because the SES likely passes from one generation to another. In census data, I measure the socio-economic status of a household using housing characteristics (i.e., assets) and the education and labor outcomes of adults.

Figure 4e and f depict the profile of two of the many house characteristics available in Census 1977 across birth cohorts. Children born in the third quarter of 1967 lived in houses with extra rooms and better access to utilities, such as piped water. In part, the anti-abortion decree increasing the proportion of mothers living in urban areas is a consequence as shown in Fig. 4d.

Figure 4g to l present the profile of variables associated with the earnings capacity of household members. Pregnant women who gave birth after the policy cutoff and who would have aborted in the absence of the decree were more educated and less likely to work in agriculture than those before. Nonetheless, no discontinuity of female employment seems to exist around June 1967. Father's education and labor outcomes show similar patterns. Importantly, the proportion of children living with their fathers was not significantly affected by the anti-abortion policy (See Fig. 4j), suggesting that men did not leave the household as a result of an unexpected child.

The graphs in Fig. 4 are all consistent with predictions of the selection process (3)–(4). However, they show statistically unconditional relationships. I discuss the identification and estimation of conditional relationship (3) in Section 9.4.

9 Empirical analysis of the intergenerational transmission of fertility

The first part of this section analyzes the extent to which socio-economic characteristics explain the fertility profiles in Figs. 2 and 3. The second part deals with the point estimation of the intergenerational transmission of fertility preference.

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[13]The excess number of youngest children after the policy was $(46 - 30)/46 = 0.348$.

9.1 Role of inherited socio-economic status in the demand for children

After imposing a linear functional form approximation, Eqs. 8 and 9 lead naturally to the following regressions:

$$chborn_{ict} = \gamma_{0t} + \gamma_{1t}\, pre_c + \eta_{1t} W_c\, pre_c + \eta_{2t} W_c\, (1 - pre_c) + \bar{X}_c^f\, \beta_{ft} + \bar{X}_c^s\, \beta_{st} + \epsilon_{ict} \tag{10}$$

The dependent variable is the number of children ever born to a second-generation woman i belonging to month-year birth cohort c observed in Census year $t \in \{1992, 2002, 2011\}$. The variable pre_c is an indicator of whether her mother had the option to abort while she was in the womb. Following the cutoff used by Pop-Eleches (2006), $pre_c = 1$ if the woman was born before June 1967; otherwise, 0 (see Fig. 1).

The variable W_c is the distance in months between the second-generation woman's date of birth and the cutoff date in June 1967. Thus, the third and fourth terms in regression Eq. 10 fit the linear trends on both sides of the cutoff, where the slopes on each side are allowed to be different. As indicated in Section 5, variables $\bar{X}_c^s$ and $\bar{X}_c^f$ are fertility determinants of the woman and her mother. The coefficient of interest is γ_{1t}. It measures the fertility differences of women born on each side of the cutoff and arbitrarily close to it.

Graphically, Fig. 3 contains the results of estimating regression 3, ignoring covariates $\bar{X}_c^s$ and $\bar{X}_c^f$. Each dot in the scatter plot is the average children ever born for a given birth cohort. The estimated coefficient, $\widehat{\gamma}_{1t}$, is the discontinuity in the regression predictions (fitted values) at the cutoff in June 1967.

Specification Eq. 10 resembles the regression discontinuity design (RD), but it is conceptually different. The standard RD assumes that individuals who are arbitrarily close to the cutoff on each side are similar in both observables and unobservables (i.e., no selection in the limit). Thus, the discontinuity is entirely the consequence of the treatment analyzed. Conversely, as shown in Section 8, women born on each side of the cutoff are significantly different regarding observables (socio-economic status) and plausibly unobservables (preferences). Accordingly, the role of covariates is different in the standard RD than in regression Eq. 10. While in the first case, covariates only improve the statistical efficiency of the estimator (Calonico et al. 2016); here, they purge the influence of SES variables.[14]

The standard RD estimation techniques apply here despite the difference in the interpretation of results, given that the objective is to compare averages on both sides of a cutoff. Therefore, the procedures to estimate (10) include OLS (global fit) and local polynomial regression with triangular kernels, as is common practice in the RDD literature.

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[14]Covariates only affect the efficiency but not the consistency of the RD estimator in the standard case because the first moment of these variables is assumed to be smooth across the cutoff (Calonico et al. 2016). In regression Eq. 10, the conditional mean of the covariates is discontinuous at the cutoff. The bandwidth selection procedure in Calonico et al. (2014) and Calonico et al. (2016) is no longer applicable here.

Conditional on covariates, the coefficients γ_{1t} in regression (10) identify the following expression obtained from Eqs. 8 and 9.

$$\gamma_{1t} = E(\xi^s | X^f, X^s, \xi^f > \tilde{\xi}_x) - E(\xi^s | X^f, X^s) - \mu \qquad (11)$$

Testing the hypothesis that $\gamma_{1t} = 0$ is equivalent to testing whether the preferences of second-generation women (ξ^s) were, on average, related to the preferences of their mothers (ξ^f) or their preferences changed significantly as a result of being "unwanted" children (μ).[15]

For simplicity, I will refer to expression (11) as the aggregate role of preferences on estimates.[16] It contains the residual effect of the anti-abortion decree after eliminating the impact of socio-economic variables.[17]

Regression (10) includes not only fertility determinants of second-generation women X^s but also fertility determinants of their mothers X^f given that these variables affect the threshold $\tilde{\xi}_x$ in expression (5). If X^f was omitted in regression (10), then, γ_{1t} would confound the role preferences in expression (11) with the parent–daughter correlation in the socio-economic status.

The lack of longitudinal data complicates the analysis, including the fact that 1992, 2002, and 2011 Census samples used to compute (10) contain no retrospective information about the parents of women born around the implementation of the anti-abortion decree, i.e., the vector X^f, nor the socio-economic variables at the beginning of their reproductive life, X^s. Nonetheless, much of this information is available in the 1977 Census, when women born around June 1967 were 9 years old.

I move information across censuses by aggregating it at the cohort level. That is, the variables $\bar{X}_c^f$ and $\bar{X}_c^s$ contain average family characteristics observed in 1977 by month-year birth cohort and county of birth. For example, the father's education of a woman born in Covasna County in February 1967 is unavailable in Census data 2002 when she was 34 years old. Then, I assign her the average level of education of men in Census 1977 who had a girl child born in Covasna County in February 1967.

Considering that the anti-abortion decree varies only across but not within cohorts, including average values in regression (10) is as good for identification purposes (i.e., no bias added) as including variables at the individual level.

Notably, Eq. 10 identifies local effects. That is, it measures the impact of the abortion ban on second-generation women born around June 1967. The method is silent regarding the impact of Decree 770 on women born several years after this cutoff.

9.2 Results

Table 1 presents the results of estimating regression (10) for various specifications. It consists of three panels that show the regression outcomes for women belonging to

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[15] Notably, a non-zero $E(\xi^s | X^f, X^s, \xi^f > \tilde{\xi}_x) - E(\xi^s | X^f, X^s)$ is sufficient to conclude that mothers and daughters' preferences were correlated.

[16] γ_{1t} aggregates the two roles played by preferences: the parent-to-child transmission of "tastes" for children and the direct psychological impact of the policy.

[17] As discussed in Section 5, the variable μ likely contains a crowding effect. In this case, γ_{1t} not only reflects preferences but a general equilibrium effect. I discuss the role of this component below.

Table 1 Fertility of second-generation women as a function of their mothers' exposure to the policy

	OLS all month-years 1962–1972 (1)	OLS excluding 2nd quar. 1967 (2)	RDD band width 24 months (3)	RDD band width 36 months (4)
	Census 1992 (24 years old[a])			
Born pre Jun 1967 (a)	0.139***	0.150***	0.072*	0.082***
	(0.016)	(0.017)	(0.028)	(0.021)
Born pre Jun 1967 (b)	0.110***	0.122***	0.060**	0.061***
	(0.015)	(0.015)	(0.023)	(0.018)
Born pre Jun 1967 (c)	0.107***	0.120***	0.062**	0.059**
	(0.015)	(0.015)	(0.024)	(0.019)
Expl. by covariates [(a)-(b)]/(a)	0.21	0.19	0.16	0.25
Expl. by covariates [(a)-(c)]/(a)	0.23	0.20	0.14	0.27
Mean pre-policy dep. variable	1.16	1.16	1.16	1.16
Obs.[b]	164,209	159,894	159,894	159,894
	Census 2002 (34 years old[a])			
Born pre Jun 1967 (a)	0.152***	0.161***	0.136***	0.147***
	(0.015)	(0.015)	(0.032)	(0.023)
Born pre Jun 1967 (b)	0.115***	0.123***	0.095***	0.110***
	(0.014)	(0.014)	(0.028)	(0.020)
Born pre Jun 1967 (c)	0.115***	0.123***	0.096***	0.111***
	(0.014)	(0.014)	(0.026)	(0.020)
Expl. by covariates [(a)-(b)]/(a)	0.24	0.24	0.30	0.26
Expl. by covariates [(a)-(c)]/(a)	0.24	0.23	0.29	0.25
Mean pre-policy dep. variable	1.73	1.73	1.73	1.73
Obs.[b]	157,412	153,316	153,316	153,316
	Census 2011 (44 years old[a])			
Born pre Jun 1967 (a)	0.154***	0.160***	0.123***	0.139***
	(0.014)	(0.014)	(0.019)	(0.018)
Born pre Jun 1967 (b)	0.116***	0.121***	0.063**	0.088***
	(0.013)	(0.014)	(0.022)	(0.018)
Born pre Jun 1967 (c)	0.119***	0.125***	0.070**	0.094***
	(0.014)	(0.015)	(0.026)	(0.021)
Expl. by covariates [(a)-(b)]/(a)	0.25	0.24	0.49	0.37
Expl. by covariates [(a)-(c)]/(a)	0.23	0.22	0.43	0.33

Table 1 (continued)

Mean pre-policy dep. variable	1.77	1.77	1.77	1.77
Obs.[b]	142,712	138,953	138,953	138,953

Clustered standard errors in parentheses (cohort level)

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Sample of women born between Jan 1962 and Dec 1972. The dependent variable is "children ever born." Each cell is from a separate regression. The columns indicate the estimation method, and the rows indicate the covariates included in the specification. RDD uses a triangular kernel. Rows (a) have no covariates. Rows (b) include dwelling (house ownership, age of the structure, rooms, area, kitchen, baths, electricity, stories in building, construction materials, and fuel used for heating) and father characteristics (education, industry, sector, and occupation) as defined in Appendix B plus urban residence in 1977, county of birth fixed effects and quarter of birth fixed effects. Rows (c) adds mother characteristics (same as those included for the father) to the covariates included in the previous case. All regressions include independent linear trends on each side of the cutoff as indicated in Eq. 10

[a] Age of individuals born in June 1967

[b] RDD methods include the total number of observations, not the effective observations after the bandwidth selection

the cohorts of interest at different points of their reproductive life. In Census 1992, women born in June 1967, i.e., at the policy cutoff, were 24 years old. In Census 2002, members of the same birth cohort were 34 years old. In Census 2011, the same women were 44 years old and reaching the end of their reproductive lives.

The cells in the table report the point estimates and associated standard errors of coefficient γ_{1t} in Eq. 10. Each cell is the result of a separate regression that combines a method (column) and a set of covariates (rows). Row (a) in each panel includes no covariates. That is, it shows $\widehat{\gamma}_{1t}$ when Eq. 10 is estimated excluding $\bar{X}_c^f$ and $\bar{X}_c^s$ from the specification.

Columns (1) and (2) show the results of estimating (10) using ordinary least squares. Figure 3 illustrates the statistically unconditional relationship between children ever born to second-generation women and whether their mothers (first-generation women) had the chance to abort when they were in the womb. The discontinuities in the fitted values in June 1967 correspond to the intersection of column (1) and row (a) in Table 1. These results suggest that the daughters of first-generation women who could not abort in the mid-1960s had between 0.14 and 0.15 fewer children, which represents a 10% reduction, than the daughters of first-generation women who had the chance to abort legally.

Column (2) in Table 1 shows the same specification as column (1) but excluding women born between April and June 1967 from the sample. Figure 1 suggests that first-generation women who gave birth in the second quarter of 1967 were only partially affected by the anti-abortion policy. Results in column (2) are very similar to those in column (1).

Columns (3) and (4) follow a standard regression discontinuity design (RDD) approach. Equation 10 is computed using local linear regressions with triangular

kernels. Although this process is not a standard RDD (see previous sections for a discussion), the computational method is appropriate to compare cohorts born around June 1967.

The selection of the bandwidth in the RDD approach implies a trade-off between variance and bias. A relatively small bandwidth minimizes the bias by comparing individuals in a small neighborhood around the policy cutoff. However, it includes relatively few observations, which increase the variance of the estimator. By contrast, a large bandwidth reduces the variance of the estimator at the expense of including individuals relatively "far" from the policy cutoff.

The results of OLS and RDD are similar in Censuses 2002 and 2011. However, the RDD coefficients in Census 1992 are half of those obtained via OLS. A possible explanation for these differences is that the OLS method puts added weight on observations located away from the cutoff, with this approach more sensitive to non-linearities in the life cycle fertility trend (see Fig. 2). The coefficients from the RDD methods appear more consistent with Fig. 3, where the discontinuity in 1992 seems smaller, than those in the other two Censuses.[18]

The fact that women born after the implementation of the anti-abortion decree had fewer children may be the result of having a lower "taste" for children, a higher initial socio-economic status, or a combination of both. Rows (b) in Table 1 show the results of estimating (10) with dwelling characteristics and father's earnings capacity variables—age, education, industry, sector of work, and occupation (see Appendix B for the comprehensive list of regressors). Moreover, this specification includes a set of county-of-birth fixed effects and an urban residence indicator. These variables, measured in 1977 when second-generation individuals were 9 years old, plausibly accurately control for the initial socio-economic status of second-generation women. As discussed in Section 4, Romania was under a socialist regime in the 1960s. Remunerations were centrally determined. Little or no dispersion of earnings is expected after conditioning on the covariates.

The coefficient of interest declines from rows (a) to (b) (bottom lines in each panel) but less than a third in most of the cases. If the covariates in the regression accurately eliminate socio-economic status differences, then, the heterogeneity in preferences is the only fertility determinant excluded among regressors.

Rows (c) in Table 1 add the characteristics of the mother associated with her earning capacity, i.e., the same type of variables as those included for fathers, to the covariates included in rows (b). Some of these characteristics are at risk of being partially endogenously determined by the policy. For example, a first-generation woman who unexpectedly had a baby as a result of the anti-abortion decree may have decided to stop working to care for the child. Although knowing the initial response to an unexpected baby is not possible, Fig. 4h suggests that none in Census 1977, ten years after the decree, was issued. It shows no differential employment rate among first-generation women affected by the anti-abortion policy.

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[18]The visual inspection of Fig. 2 shows no apparent discontinuity in 1992. However, the "jump" seems smaller because of the slope of the curve as the regression analysis shows.

Rows (c) in Table 1 show that the percentage decline in the discontinuity remained virtually unchanged from rows (b) to (c).

Birth order and sibship size   The odd-numbered columns in Table 2 add two important regressors. One of them is the number of older siblings of second-generation women. The model in Section 5 indicates that the birth parity was an important determinant of first-generation women's propensity to abort. Additionally, the literature documents that the birth order of a child (i.e., the number of older siblings plus one) is associated with his/her adult socio-economic outcomes (Black et al. 2005).

The second variable included in the regression is the total number of brothers and sisters of second-generation women. Two reasons explain why this variable is relevant. First, the fertility quantity–quality literature (Becker and Lewis 1973; Willis 1973) stresses that larger families invest relatively little on each child. Given that the anti-abortion decree increased the size of the family, women born after the implementation of the policy may have had lower earnings capacities affecting their demand for children. Second, being born in larger families may affect the subjective optimal number of children. For example, second-generation women who had many siblings may be likely to have many children as an imitation of parent's behavior.

Results in Table 2 show that including birth order and sibship size mildly contributes to explaining the discontinuity. Comparing results from rows (c) in Table 1 to those in the odd columns of Table 2 indicates the similarity in the estimated values for the coefficient of interest.

Education and marital status of second-generation women   Figure 4 shows that women who gave birth after the implementation of the anti-abortion policy were, on average, more educated than those before. The "taste" for education may also pass from parents to children. In such a case, the tendency of second-generation women to stay in school for more years may induce them to postpone marriage and childbearing.

The even columns in Table 2 show the results when the level of education, marital status, and the age of first marriage of second-generation women are part of the regressors. These variables are certainly endogenously determined. Their inclusion in the regression likely dwarfs the coefficient of interest. Notably, although a relatively high taste for education may lower the demand for children, a relatively low "taste" for children may also give extra time for women to study further.

Despite the potential endogeneity of education and marital outcomes of second-generation women, the inclusion of these variables in the regression provides valuable information. If the observed lower fertility among women born after the policy implementation is entirely the result of their desire for further study and postponement of marriage, then, conditioning on these variables should reduce the coefficient of interest to zero.

Table 2 shows that the indicator variable for being born before June 1967 halved after conditioning on education and marital outcomes. However, the fact that this coefficient remains significant suggests that the average taste for children of second-generation women was significantly affected by the policy. This result can be

Table 2 Fertility of second-generation women as a function of their mothers' exposure to the policy (additional regressors)

| | Census 1992 | | | | Census 2002 | | | | Census 2011 | | | |
| | OLS | | RDD | | OLS | | RDD | | OLS | | RDD | |
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Born pre Jun 1967	0.094***	0.063**	0.054**	0.010	0.111***	0.074***	0.102***	0.047**	0.116***	0.061***	0.092***	0.053*
	(0.015)	(0.022)	(0.019)	(0.015)	(0.015)	(0.019)	(0.022)	(0.018)	(0.015)	(0.016)	(0.023)	(0.024)
Tot. num. of siblings	0.065***	0.065***	0.021	0.023	0.039**	0.039**	0.035	0.036	0.039*	0.039*	0.020	0.019
	(0.011)	(0.011)	(0.019)	(0.019)	(0.013)	(0.013)	(0.027)	(0.026)	(0.015)	(0.015)	(0.026)	(0.026)
Older siblings	−0.165***	−0.165***	−0.026	−0.020	−0.061*	−0.061*	−0.039	−0.037	−0.038	−0.034	0.012	0.015
	(0.018)	(0.018)	(0.032)	(0.032)	(0.027)	(0.026)	(0.040)	(0.040)	(0.029)	(0.029)	(0.055)	(0.056)
Less than primary		0.750		3.013***		1.867**		3.685***		1.215*		0.267
		(0.652)		(0.606)		(0.559)		(0.709)		(0.543)		(0.682)
Primary educ.		0.683***		0.214		0.373		0.240		0.012		−0.164
		(0.193)		(0.207)		(0.246)		(0.274)		(0.252)		(0.415)
College		0.934**		−0.467		−0.637		0.614		0.075		−0.067
		(0.347)		(0.339)		(0.348)		(0.589)		(0.267)		(0.398)
Never married		−1.907**		−2.486**		−1.723**		−2.192***		−2.704***		−1.802*
		(0.725)		(0.848)		(0.527)		(0.585)		(0.460)		(0.740)
Age when married		−0.096*		−0.093*		−0.070***		−0.077***		−0.086***		−0.048
		(0.039)		(0.044)		(0.020)		(0.022)		(0.019)		(0.030)
House charact.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Father charact.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mother charact.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
County FE charact.	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Table 2 (continued)

	Census 1992				Census 2002				Census 2011			
	OLS		RDD		OLS		RDD		OLS		RDD	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Birth quarter FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Expl. by covariates	0.38	0.58	0.34	0.88	0.31	0.54	0.31	0.68	0.27	0.62	0.34	0.62
Pre-policy dep. variable	1.16	1.16	1.16	1.16	1.73	1.73	1.73	1.73	1.77	1.77	1.77	1.77
Obs.[a]	159,894	159,894	159,894	159,894	153,316	153,316	153,316	153,316	138,953	138,953	138,953	138,953

Clustered standard errors in parentheses (cohort level)

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Sample of women born between Jan 1962 and Dec 1972. The dependent variable is "children ever born." Women born in the second quarter of 1967 are excluded from the sample. Triangular kernel with 36-month band width for RDD

[a] RDD methods include the total number of observations, not the effective observations after the bandwidth selection

explained either by a positive correlation in the taste for children across generations or by the direct change in preferences resulting from being an "unwanted" child.

Section 2 suggests the possibility that the variable μ in expression (11) contains a crowding effect. That is, a disproportionately larger cohort may have congested publicly provided goods. Pop-Eleches (2006) reported a significant decline in the years of education due to the Romanian anti-abortion policy.

If the crowding effect only affected the *quantity* of education, then, it is not a concern for the empirical approach given that the regression results in Table 2 account for it. However, if the quality of education was affected, then, the covariates in Table 2 are insufficient to eliminate the crowding effect. Whether the quality of education can affect women's fertility in a way that is unrelated to the "taste" for children is unclear. However, if such an effect existed, then, being disentangled from the aggregate role of preferences in Eq. 11 is not possible.

Heterogeneity across educational groups As previously mentioned, the level of education of second-generation women is potentially endogenous as indirectly influenced by Decree 770 issued in 1966. With this caveat, I classify second-generation women into three educational groups: women with primary education or less, women with secondary schooling completed, and women with college education.

Table 3 panel A measures the difference in fertility between women born before and after the policy cutoff. The empirical specification for the calculation of the coefficients is identical to the one used to compute the first row of Table 1. I only use censuses 2002 and 2011. In 1992, many women around the cutoff were still attending school.

The results in panel A indicate that the impact of the abortion ban was more pronounced for low-educated women. Among college-educated women, the magnitude of the discontinuity is small and statistically not different from zero.

Panel B in Table 3 shows the magnitudes of the discontinuities after conditioning on the full set of covariates. These values are methodologically comparable (i.e., same set of covariates) to those in the even-numbered columns of Table 2. Similar to previous results, the coefficients in Panel B are smaller than those in Panel A. However, the negative correlation between the level of education and the discontinuity jump persists.

Placebo policy This section shows the results of a placebo policy implementation. Abortion was legal in Romania from September 1957 (Decree 463) until October 1966 (Decree 770). During this period, no significant change occurred in population policies. Nonetheless, I assume that the government of Romania banned abortion in 1960 (a placebo discontinuity in June 1961) rather than in 1966 (observed treatment discontinuity in June 1967).

Table 4 presents the results of the placebo experiment. The coefficients in this table are methodologically comparable to those in Table 1. As expected, none of the results are statistically different from zero.

Table 3 Impact of the abortion ban on fertility by educational groups

	Census 2002		Census 2011	
	OLS	RDD	OLS	RDD
	(1)	(2)	(3)	(4)
	Panel A: No covariates			
Coefficient of "Born pre Jun 1967" for women with:				
Primary education or less	0.253***	0.132**	0.318***	0.172**
	(0.036)	(0.041)	(0.041)	(0.053)
Secondary education	0.142***	0.127***	0.140***	0.089***
	(0.012)	(0.018)	(0.015)	(0.016)
College education or more	0.051	0.021	0.035	0.047
	(0.029)	(0.044)	(0.020)	(0.030)
	Panel B: Full set of covariates			
Coefficient of "Born pre Jun-1967" for women with:				
Primary education or less	0.168***	0.105*	0.162***	0.143*
	(0.050)	(0.045)	(0.047)	(0.064)
Secondary education	0.046**	0.055*	0.065***	0.053*
	(0.017)	(0.024)	(0.018)	(0.027)
College education or more	0.044	−0.006	0.015	−0.017
	(0.037)	(0.044)	(0.030)	(0.041)

Each cell is from a separate regression. The columns indicate the estimation method and the rows the women included in the analysis. RDD uses a triangular kernel. All regressions include independent linear trends on each side of the cutoff as indicated in Eq. 10. The number of observations in 2002 is 44,448 for primary education or less, 93,295 for secondary education, and 15,489 for college education or more. In 2011, the number of observations is 33,089 for primary education or less, 79,528 for secondary education, and 26,336 for college or more. RDD methods include the total number of observations, not the effective observations after the bandwidth selection

9.3 Is international migration a concern for identification?

The communist regime' maintenance of power until 1990 restricted migration. Despite tight controls, 15,000 to 20,000 Romanian emigrated per year (Dethier et al. 1994). Romanian citizens were free to migrate after Nicolae Ceauescu's government ended in December 1989. Are the results presented earlier possibly biased due to migration?

The Romanian emigration rate is unknown. Nonetheless, the analysis in Fig. 2 and the associated regression results are biased only if (i) women born after June 1967 emigrated at a different rate than women born before that cutoff date, and (ii) women who migrated had different fertility preferences than non-migrants.

Table 4 Fertility of second-generation women placebo policy

	OLS all month-years 1957–1966 (1)	OLS excluding 2nd quar. 1961 (2)	RDD band width 24 months (3)	RDD band width 36 months (4)
Census 1992 (30 years old[a])				
Born pre Jun 1961	−0.012	−0.010	0.024	0.027
	(0.015)	(0.017)	(0.024)	(0.019)
Obs.	137,347	133,877	133,877	133,877
Census 2002 (40 years old[a])				
Born pre Jun 1961	0.007	0.002	−0.017	−0.008
	(0.014)	(0.015)	(0.017)	(0.017)
Obs.	131,249	127,965	127,965	127,965
Census 2011 (50 years old[a])				
Born pre Jun 1961	0.004	0.002	0.023	0.008
	(0.016)	(0.016)	(0.026)	(0.021)
Obs.	119,953	116,909	116,909	116,909

Clustered standard errors in parentheses

*$p < 0.05$, **$p < 0.01$, ***$p < 0.001$

Each cell is from a separate regression. The columns indicate the estimation method and the rows the covariates included in the specification. RDD uses a triangular kernel

[a] Age of individuals born in June 1961

[b] RDD methods include the total number of observations, not the effective observations after the bandwidth selection

Figure 5 shows the relative cohort size of women born between January 1962 and December 1972 over multiple censuses. The gap that separates the series is small in absolute and relative terms at all points. More relevant for the empirical approach, the vertical difference in the series across censuses is not different between the cohorts born before and those born after June 1967. This fact is sufficient to conclude that migration cannot explain the discontinuities in Figs. 2 and 3.

9.4 Measuring the intergenerational correlation of preferences in the absence of direct "unwantedness" effect

This section estimates the parent–child correlation of preferences under the somewhat strong assumption that the direct effect of being "unwanted" is negligible (i.e., the variable $\mu = 0$ in expression (6)).

Children born as a result of unwanted pregnancies may be less loved and likely to be neglected. However, determining causal evidence supporting this presumption is difficult. Although the medical and public health literature documents that
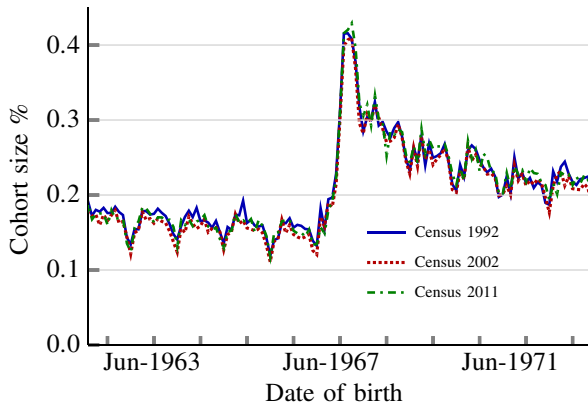
Fig. 5 Cohort size over census years

unwanted children perform worse in adult life (Forssman and Thuwe 1966; Dytrych et al. 1975), it fails to separate the role of "unwantedness" per se, which potentially affects the allocation of resources between wanted and unwanted children, from a socio-economic determinant across households.

Dytrych et al. (1975) studied 220 individuals born in Czechoslovakia during the period 1961–1963 whose mothers requested to abort them, but the government denied it. These individuals form the "treatment" group of unwanted children. Acknowledging the importance of determining a proper control group, researchers gathered an equal number of individuals whose parents were similar in observable characteristics but did not request an abortion.

The problem with Dytrych et al. (1975) study is that the self-selection of mothers to undergo an abortion is determined by their taste for children, consequently rendering the treatment and control group incomparable. Despite the limitation of this study, the authors stated that "The expectation that unwanted conceptions would lead inevitably to the children being unwanted proved **not** to be the case" p. 165 (emphasis in original paper).

Economics papers emphasize that women who are more likely to abort face less favorable socio-economic conditions (e.g., Donohue and Levitt (2001) and Pop-Eleches (2006)). Precisely, the exposure to this unfavorable environment arguably explains why "unwanted" children perform worse in adulthood. In the current paper, if the information available in Romanian Censuses is sufficient to account for the differences in the socio-economic status of households as argued in Section 4, then, this mechanism linking the mothers' desire to abort and the outcomes of their offspring is eliminated.

In summary, no causal evidence proves that mothers vary in nurturing their children. Instead, the literature indicates that children born after unplanned pregnancies live in less favorable socio-economic environments, which the empirical analysis of Section 9 contemplates. Nonetheless, the estimated correlations shown below should be interpreted with caution.

The procedure The expression for γ_{1t} obtained in the previous section contains information about the intergenerational transmission of preferences affecting the demand for children. Assuming that $\mu = 0$ in Eq. 11 and the distribution of preferences (ξ^f) among first-generation women is well-approximated by normal distribution, then, γ_{1t} in regression Eq. 10 can be written as follows:

$$\gamma_{1t} = E(\xi^s|X^s, \xi^f > \tilde{\xi}_x) - E(\xi^s|X^s) \tag{12}$$

$$= E(\xi^s - E(\xi^s)|X^s, \xi^f > \tilde{\xi}_x) \tag{13}$$

$$= \sigma^s \rho \, E(\xi^f|X^s, \xi^f > \tilde{\xi}_x) \tag{14}$$

$$= \sigma^s \rho \frac{\phi(\tilde{\xi}_x)}{1 - \Phi(\tilde{\xi}_x)} \tag{15}$$

The derivation (12)-(15) closely follows (Heckman 1979). As long as the fertility preferences of second-generation women are a linear function of their mother's preferences,

$$\xi^s - E(\xi^s) = \sigma^s \rho \, \xi^f + error \tag{16}$$

where σ^s is the standard deviation of ξ^s and $\rho = corr(\xi^f, \xi^s)$, then, expression (14) follows immediately from Eq. 13.[19] Furthermore, assuming that ξ^f is normally distributed with mean zero and standard deviation one, then, the conditional expectation in Eq. 14 is the well-known inverse mills ratio defined as the ratio of a standard normal density $\phi(.)$ and its cumulative distribution $1 - \Phi(.)$ evaluated at the truncation point $\tilde{\xi}_x$.

Replacing γ_{1t} in Eq. 10 with the right-hand side of Eq. 15 provides the following estimating equation:

$$chborn_{ict} = \beta_{0t} + \beta_\lambda \, pre_c \, \bar{\lambda}(\tilde{\xi}_x) + \eta_{1t} W_c \, pre_c + \eta_{2t} W_c \, (1 - pre_c) + \bar{X}^s_c \beta_{st} + \epsilon_{ict} \tag{17}$$
$$where$$
$$\beta_\lambda = \sigma^s \rho$$
$$\lambda(.) = \phi(.)/(1 - \Phi(.))$$

The coefficient β_λ is the correlation of preferences across generation scaled by the standard deviation of residuals. Regression (17) is identical to a Heckman (1979) sample selection model for pre-policy periods ($pre_c = 1$), where the inverse mills ratio $\lambda(.)$ is included to account for the fact that first-generation women self-selected motherhood. However, when $pre_c = 0$, the second term on the right-hand side disappears given that the anti-abortion decree eliminated the possibility of having an abortion.[20]

Compared with regression (10), X^f enters regression (17) only through the inverse mills ratio given that these variables are expected to affect $chborn_{ict}$ only by changing the mother's threshold $\tilde{\xi}_x$. Nonetheless, most of the variables in X^f are also in

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[19]The linearity of Eq. 16 imposes no constraint in practice given that it always holds for linear projections, which is the regression method used in the paper. Notably, the slope of the OLS/linear projection of ξ^s on ξ^f is $\frac{cov(\xi^f, \xi^s)}{var(\xi^f)} \equiv \sigma^s \rho$ considering that $var(\xi^f) = 1$.

[20]Notably, λ is aggregated at the cohort-geographic level due to data limitations. Accordingly, the direct "unwantedness" effect μ in Eq. 11 cannot be separately identified.

X^s because the socio-economic status variables affecting first-generation women's propensity to abort are the same variables that account for the initial socio-economic status of second-generation women.

Importantly, the presence of the indicator variable pre_c in regression (17) implies that β_λ is strongly identified even when the full set of variables in X^f is included in X^s. This finding is a remarkable advantage relative to the standard Heckman model, which relies on the functional form of $\lambda(.)$ in the absence of an exclusion restriction.

The estimation of the intergenerational transmission of preferences ρ via regression (17) has two complications. The first one is the computation of the inverse mills ratio $\lambda(.)$. The fact that the pregnancy status of women is unobservable renders the Heckman (1979) approach infeasible. The second complication is the estimation of σ^s, such that the correlation coefficient of preferences can be computed as $\widehat{\rho} = \widehat{\beta_\lambda}/\widehat{\sigma}^s$.

Considering that some of the regressors vary at the cohort level, then, the error term in Eq. 17 contains extra variability. The solution to these two complications requires a reformulation of the likelihood function and cross-dataset variance estimation. Owing to space limitation, these methodological issues are in Appendix A.

Results Figure 6 summarizes the results of estimating intergenerational fertility correlations and the portion explained by inherited preferences.[21]

Figure 6a shows the correlations under assumptions A1 (no anticipation) and A2 (full compliance; see Appendix A). In this case, the intergenerational fertility correlations when women were 34 years old (Census 2002) and 44 years old (Census 2011) are between 0.12 and 0.14. After conditioning on variables associated with the socio-economic status of first-generation parents and second-generation daughters, the correlations decline by a third, suggesting that the transmission of preferences across generations plays a significant role in parent–daughter fertility correlations.

Figure 6b shows the estimated correlations assuming that five were aborting for each birth in the pre-policy years (Dethier et al. 1994; see assumption B1 in Appendix A). The magnitudes of the correlations increase in all cases. However, the part attributed to inherited preferences remained robust at approximately two-thirds of the total fertility correlation.

The aggregate intergenerational fertility correlations found in this section are similar in magnitude to those found in other studies (Danziger and Neuman 1989; Murphy 1999; Kolk 2014; Murphy 2012; Murphy and Knudsen 2002).

10 Second-generation men

The previous sections analyzed the fertility behavior of second-generation women. However, the demand for children of the men born around the policy cutoff is also relevant. Men may also inherit the "taste" for children of their families and affect the choices that they and their female partners make about childbearing.

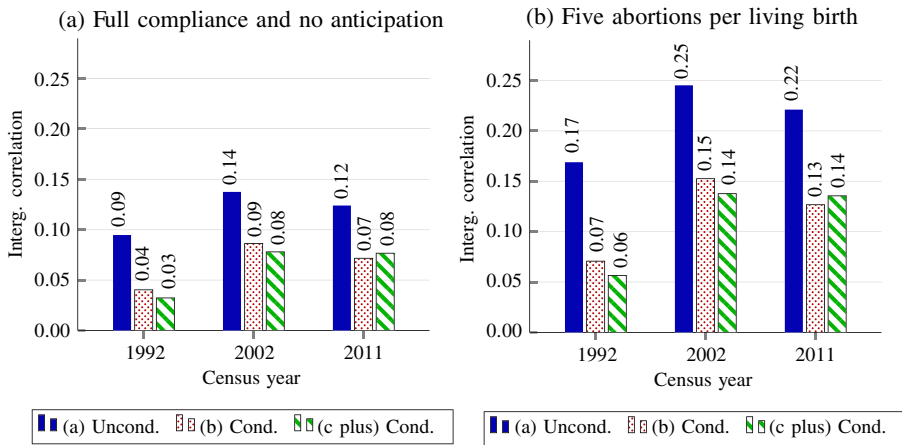[21]For a full set of results, see Tables A5 and A6 in online Appendix 4.

Fig. 6 **a** The intergenerational correlation without including any covariates. **b** Conditioning on dwelling and father characteristics as defined in Table 1 and Appendix B plus urban residence in 1977, county of birth fixed effects and quarter of birth fixed effects; (c plus) adds mother characteristics to the covariates included in the previous case plus total number of siblings and older siblings. All results were computed by RDD using a triangular kernel of 36-month band width

Censuses provide no information about the number of children ever fathered by men. However, they contain two closely related variables: (i) the number of own children living in the same household and (ii) the number of children ever born to their wives. These two variables have limitations. In the first case, children who left the household to form their own families are not in the sample. Then, the demand for children appears to be lower than what it is. In the second case, the analysis is possible only among men who decided to marry. Moreover, it cannot account for most of the births from previous marriages as women generally receive custody.

These variable limitations are problematic to the extent that they have different effects on men on each side of the policy cutoff, which seems not to be the case. Although twenty percent of second-generation men report living without a spouse, no discontinuity exists in the marital status of men around the policy cutoff, which rules out the possibility that selection into marriage affects the impact of the abortion ban on the demand for children.

Figure 7 shows the scatter plot for the number of own children in the household and wife's children ever born. These graphs are analogous to those in Fig. 3 but for men. All of them indicate that the demand for children is higher among men born before June 1967.

A concern in Fig. 7 is that the fertility behavior of the wife, not the husband, generates the discontinuity at the policy cutoff. If men born on one side of the policy cutoff were more likely to marry women on one of the two sides of the cutoff, then, distinguishing the role of men and women regarding fertility decisions would be impossible. However, it does not seem to be the case. Figure 8 shows the proportion of men married to women born before the anti-abortion decree (y-axis) as a function of men's birth cohort (x-axis). The graph shows no discontinuous jump at the cutoff value in June 1967.
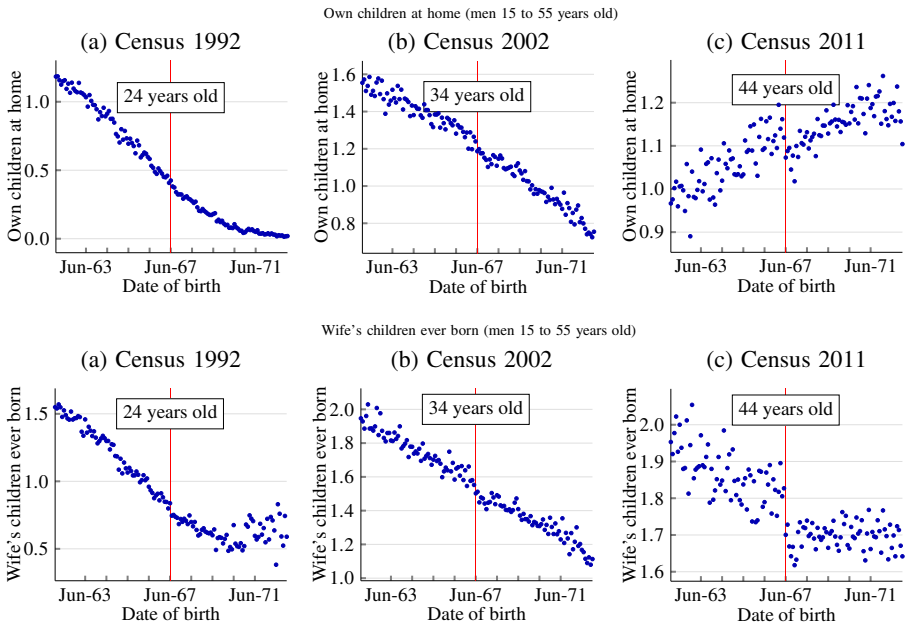
Own children at home (men 15 to 55 years old)

Fig. 7 Own children at home (men 15 to 55 years old) and Wife?s children ever born (men 15 to 55 years old)

Table 5 is similar to Table 1 but for the sample of second-generation men. When the dependent variable is the wife's children ever born, I condition the analysis on the wife's date of birth to eliminate the potential impact of the abortion ban channeled through the spouse.
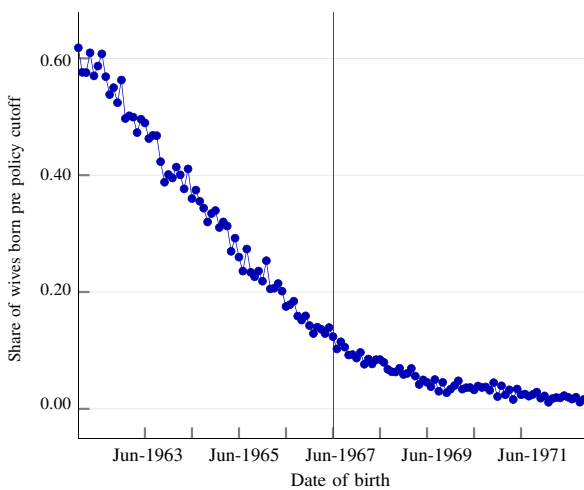

Fig. 8 Men's spouse born before June 1967 - Census 2011 (by men's month of birth)

Rows (a) in Table 5 show unconditional relations, which correspond to the magnitude of the discontinuities in Fig. 7. After conditioning on variables associated with the socio-economic status, the discontinuity declines in magnitude but does not disappear, particularly in Census 2011. This result suggests that men inherit wealth as well as preferences from their families, which affect the demand for children.

Table 5 Own children in household and wife's children ever born

	Census 2002 (34 years old[a])			
	Own children at home[b]		Wife's children ever born[c]	
	OLS	RDD	OLS	RDD
Born pre Jun 1967 (a)	0.046***	0.045*	0.059***	0.089***
	(0.013)	(0.019)	(0.015)	(0.021)
Born pre Jun 1967 (c)	0.020	0.024	0.029	0.063**
	(0.012)	(0.017)	(0.015)	(0.021)
Wife's date of birth	No	No	Yes	Yes
Expl. by covariates [(a)-(c)]/(a)	0.57	0.46	0.52	0.29
Obs.	153,219	153,219	119,043	119,043
	Census 2011 (44 years old[a])			
	Own children at home[b]		Wife's children ever born[c]	
	OLS	RDD	OLS	RDD
Born pre Jun 1967 (a)	0.056***	0.090***	0.097***	0.147***
	(0.014)	(0.020)	(0.019)	(0.028)
Born pre Jun 1967 (c)	0.043**	0.076***	0.059**	0.118***
	(0.014)	(0.018)	(0.018)	(0.021)
Wife's date of birth	No	No	Yes	Yes
Expl. by covariates [(a)-(c)]/(a)	0.23	0.15	0.39	0.20
Obs.[d]	142,267	142,267	106,986	106,986

Clustered standard errors in parentheses (cohort level)

*$p < 0.05$, **$p < 0.01$, ***$p < 0.001$

sample of men born between Jan 1962 and Dec 1972. Each cell is obtained from a separate regression. RDD uses a triangular kernel with 36 months band width. Rows (a) have no covariates. Rows (c), analogous to Table 1, include dwelling, father and mother, and characteristics as defined in Table 1 and Appendix B plus urban residence in 1977, county of birth fixed effects, and quarter of birth fixed effects. All regressions include independent linear trends on each side of the cutoff as indicated in Eq. 10

[a] Age of individuals born in June 1967

[b] Sample includes both married and unmarried men.

[c] Sample includes only men whose wife is at home

[d] RDD methods include the total number of observations, not the effective observations after the bandwidth selection

11 Summary and conclusions

In 1966, President Ceausescu issued a decree banning abortion and other fertility control methods. This pro-birth policy discontinuously affected women who have become pregnant in the last trimester of 1966 (the first generation) but had a limited direct effect on their daughters and sons (the second generation) since the abortion ban ended in 1989. The current research investigated the intergenerational fertility effect of this abortion ban and the mechanisms underneath the relationship of interest.

The understanding of intergenerational fertility effects has important implications. For example, it offers a better and more comprehensive view of the impact of population policies in the medium and the long run, presents valuable input to understand the pattern and speed of the demographic transition, and serves as an input to model economic mobility and income inequality persistence.

Results indicate that the intergenerational fertility correlation ranges from 0.15 to 0.25. Inherited socio-economic status explains one-third of these correlations. The intergenerational transmission of preferences presumably explains the rest. After conditioning on mediators such as the education of second-generation women and their age at first marriage, results remain strong, suggesting a significant role of the inherited "taste" for children. Conditioning on sibship size scarcely affects the magnitude of the findings. Second-generation individuals seem unable to replicate their parent's family structure. Instead, they inherit the values and norms of the previous generation.

Compliance with ethical standards

Conflict of interests The author declares that he/she has no conflict of interest.

Appendix: A

Computing the inverse mills ratio Regression (17) requires the computation of the inverse mills ratio $\lambda(.)$ to include it as a covariate. Following Heckman (1979), the inverse mills ratio can be obtained by estimating a probit model for the probability that a pregnant woman chose to give birth instead of aborting. The probit model can be written in latent variable form, which is derived from Eqs. 3 and 4 after assuming (i) the functional form $g(X^f, \xi^f) = X^f\alpha - \xi^f$, and consequently $\tilde{\xi}_x = X\alpha$ in regression (17), and ii) the distribution of fertility preferences ξ^f to be normal.

$$a*_i \quad = X_i^f\alpha - \xi_i^f \tag{18}$$

$$a_i = \begin{cases} 1 & \text{if } a*_i > 0, \\ 0 & \text{if } a*_i \leq 0. \end{cases} \tag{19}$$

Equations 18 and 19 give the standard expressions for the probability of giving birth $P(a = 0|X^f) = \Phi(X_i^f \alpha)$ and the inverse mills ratio $\lambda(X_i^f \alpha) = \phi(X_i^f \alpha)/1 - \Phi(X_i^f \alpha)$. The estimation of the model is not straightforward with the structure of the data available. The standard probit requires classifying women into those who had an abortion ($a = 1$) and those who gave birth ($a = 0$). The information contained in censuses is insufficient to make such classification. However, overcoming this problem is possible.

In any pre-policy period, knowing who carried the pregnancy to term is possible, because all the children are in the data but not who aborted.[22] Given the absence of the pregnancy status of women in the data, women who did not give birth in the months before June 1967 either aborted or were simply not pregnant. Conversely, none of the women who were in their first trimester of pregnancy had the option to abort after the decree was enforced. If compliance was high, then, the number of children born in July–September of 1967 should be the same as the number of pregnant women with pregnancy due date in that trimester (see Fig. 1). Thus, the unconditional probability of not having an abortion is the ratio of the cohort size born in July–Sept 1966 (i.e., the number of women who decided not to abort when abortion was legal) to the cohort size born in July–Sept 1967 (i.e., the total number of pregnant women). Formally, the variables b are defined as follows:

$$b_i = \begin{cases} 1 & \text{if woman } i \text{ gave birth in Jan–Mar 1967,} \\ 0 & \text{if woman } i \text{ gave birth in Jul–Sep 1967.} \end{cases} \quad (20)$$

Additionally, let s be the number of women with pregnancy due date in Jan–Mar 1967, and n the total number of women for which b is defined. Then, the unconditional probability of not having an abortion before the implementation of the pro-natalist policy is

$$P(a_i = 0) = \frac{\sum_{i=1}^{n} b_i}{s} \quad (21)$$

$$= \frac{\sum_{i}^{n} b_i}{\sum_{i=1}^{n}(1 - b_i)} \quad (22)$$

$$= \frac{\sum_{i}^{n} b_i}{n} \times \frac{n}{\sum_{i=1}^{n}(1 - b_i)} \quad (23)$$

$$= \frac{P(b_i = 1)}{P(b_i = 0)} \quad (24)$$

Equalities (21)–(24) prove the assumptions needed to compute the probit model. The right-hand side of Eq. 21 indicates that the probability of not having an abortion is, by definition, the ratio of births in Jan–Mar 1967 to the number of pregnant women who were supposed to give birth in that period had none of them aborted. The denominator s is not observed but can be replaced by a similar quantity under two assumptions.

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[22] Assuming for simplicity that infant death was low.

Assumption A1

The number of women with pregnancy due date in Jan–Mar 1967 was the same as the number of women with pregnancy due date in Jul–Sep 1967.

Figure 1 suggests that Assumption A1 is plausible. The cohort size before the pro-natalist policy was relatively constant over time, which is consistent with stable pregnancy and abortion rates in the absence of the policy.[23] More stringently, this assumption also says that the implementation of the decree occurred suddenly and prevented women with pregnancy due date in Jul–Sep 1967 from adjusting their behavior in anticipation of the policy. Assumption B1 relaxes this assumption.

Assumption A2 Full compliance with the anti-abortion decree in the short run.

Assumption A2 implies that the total number of pregnancies due in Jul–Sep 1967 effectively resulted in births up to a regular miscarriage rate. That is, no pregnant woman violated the anti-abortion decree (see assumption B1 below for a relaxation of A2).

Under assumptions A1 and A2, the denominator in Eq. 21 can be replaced by the number of births in Jul–Sep 1967, resulting in expression (22). Then, multiplying the numerator and the denominator by n in Eq. 23 provides the final result in Eq. 24.

Equalities (21)-(24) also hold when the probabilities are written conditional on covariates X^f. Then, they can be used in a likelihood function to compute the probit model (18)-(19).

$$L = \prod_{i=1}^{n} P(b_i = 1|X_i^f)^{b_i} \left(1 - P(b_i = 1|X_i^f)\right)^{1-b_i} \qquad (25)$$

Equation 25 is the likelihood function for a binary model with dependent variable b_i, as defined in Eq. 20, and regressors X_i^f. The relationship (26) between the conditional probabilities in Eq. 25 and the corresponding probabilities of not having an abortion are obtained from equalities (21)- (24).

$$P(b_i = 1|X_i^f) = \frac{P(a_i = 0|X_i^f)}{1 + P(a_i = 0|X_i^f)} \qquad (26)$$

$$= \frac{\Phi(X_i\alpha)}{1 + \Phi(X_i\alpha)} \qquad (27)$$

The specific functional form in Eq. 27 comes directly from the normality assumption of probit model (18)-(19). Replacing (27) into (25) gives the computationally

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[23] All computations in this paper were also performed assuming that the proportion of women with pregnancy due date in Jul–Sep 1967 was the same as that in (i) Jul–Sep 1966, which eliminates seasonality concerning assumption A1, and (ii) the average trimester pregnancies in the past five years. Results are nearly identical to these changes in assumption A1.

feasible likelihood function (28).

$$L(\alpha) = \prod_{i=1}^{n} \left(\frac{\Phi(X_i\alpha)}{1 + \Phi(X_i\alpha)} \right)^{b_i} \left(1 - \frac{\Phi(X_i\alpha)}{1 + \Phi(X_i\alpha)} \right)^{1-b_i} \tag{28}$$

Under assumptions A1 and A2, the coefficients $\widehat{\alpha}$ obtained from maximizing the function (28) are identical to those from the unfeasible standard probit for the probability of not aborting on a sample of pregnant women s.[24] After $\widehat{\alpha}$ is obtained, the computation of the inverse mills ratio to include it as a regressor in Eq. 17 is straightforward.

The selection probabilities from the maximization of Eq. 28 are downwardly biased if either some women anticipated the anti-abortion policy and took precautions not to get pregnant (violation of assumption A1), or some women who were already pregnant when the decree was issued aborted anyway despite the ban (violation of assumption A2). The World Bank country report describing the social and economic environment in socialist Romania mentions that five abortions were committed for every live birth in 1965, the year before the anti-abortion decree was issued (Dethier et al. 1994; p. 142). Assumption B1 takes this abortion rate as an alternative to assumptions A1 and A2.

Assumption B1: The ratio of abortions for every live birth prior to the anti-abortion decree was 5 to 1 (Dethier et al. 1994).

Under assumptions A1 and A2, Fig. 1 indicates only 1.5 abortions for every live birth, much lower than what assumption B1 states. Using assumption B1 instead of A1 and A2 requires "inflating" the probabilities of abortion. The modified likelihood function 30 does it by including an additional parameter ψ. The value of this new parameter is obtained after imposing that the unconditional probability of not having an abortion is 1/6 (i.e., 5 abortions per live birth). Results for assumptions A1 and A2, as well as for assumption B1 are computed below.[25]

$$L(\alpha, \psi) = \prod_{i=1}^{n} \left(\frac{\Phi(X_i\alpha + \psi)}{1 + \Phi(X_i\alpha + \psi)} \right)^{b_i} \left(1 - \frac{\Phi(X_i\alpha + \psi)}{1 + \Phi(X_i\alpha + \psi)} \right)^{1-b_i} \tag{30}$$

Computing the standard deviation of fertility preferences Regression (17) is the fertility equation of second-generation women. The dependent variable, the number of

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[24]The coefficients $\widehat{\alpha}$ from maximizing function (28) are the same as the coefficients of maximizing the function,

$$L(\alpha) = \prod_{i=1}^{s} (\Phi(X_i\alpha)^{1-a_i} (1 - \Phi(X_i\alpha))^{a_i} \tag{29}$$

,which is the likelihood function of the standard probit model for Eqs. 18- 19. As emphasized at the beginning of this section, this function cannot be computed because neither pregnant women in s - the observations to include in the maximization - nor the action of having an abortion a_i is observed.

[25]Under assumptions A1 and A2, the unconditional probability of not getting an abortion is 0.428 (approx. 1.5 abortions per live birth), which is obtained as $0.428 = \Phi^{-1}(-0.181)$ where the value -0.181 is the constant in the maximization of the likelihood function (28) without regressors. Then, the value of ψ used for modified likelihood function (30) is $\psi = -0.786$ since $P(a = 0) = 0.167 = \Phi^{-1}(-0.181 - 0.786)$. This probability of not aborting is the one stated in assumption B1.

children ever born, is observed in Censuses 1992, 2002, and 2011 when they were adults. On the other hand, the inverse mills ratio $\lambda(.)$ and all the information about the socio-economic status of second-generation women at the beginning of their reproductive life X^s are obtained from Census 1977, when these women were children. Because it is not possible to link individuals across census years, these covariates are included in the regression after being aggregated at the cohort level. This aggregation procedure, together with appropriate clustered standard errors, does not affect the estimation of regression coefficients. But, it affects the standard deviation of the residuals.

$$chborn_{ict} = \beta_0 + \beta_\lambda \, pre_c \, \lambda(\tilde{\xi}_{xi}) + \eta_1 W_c \, pre_c + \eta_2 W_c \, (1 - pre_c) + X_{ic}^s \, \beta_s + \xi_{ic}^s \,(31)$$

$$= \beta_0 + \beta_\lambda \, pre_c \, \bar{\lambda}(\tilde{\xi}_x) + \eta_1 W_c \, pre_c + \eta_2 W_c \, (1 - pre_c) + \bar{X}_c^s \, \beta_s + \ldots$$

$$\ldots \underbrace{\beta_\lambda \, pre_c \, (\lambda(\tilde{\xi}_{xi}) - \bar{\lambda}(\tilde{\xi}_x)) + (X_{ic}^s - \bar{X}_c^s) \, \beta_s + \xi_{ic}^s}_{\epsilon_{ict}} \quad (32)$$

The ideal fertility (31) contains regressors $\lambda(\tilde{\xi}_{xi})$ and X_{ic}^s at the individual level. Adding and subtracting cohort aggregates for these covariates gives the computationally feasible (32). The error term ϵ_{ict} contains not just preferences ξ_{ic}^s but deviations of individual regressors from cohort averages. Then, since $(X_{ic}^s - \bar{X}_c^s)$ is orthogonal to ξ_{ic}^s, the variance of ϵ_{ic} for those born in Jul–Sep 1967 (i.e., $pre_c = 0$, which eliminates the term $(\lambda(\tilde{\xi}_{xi}) - \bar{\lambda}(\tilde{\xi}_x))$ from the error ϵ_{ict}) is:

$$Var(\xi_{ic}^s) = Var(\epsilon_{ic}) - Var((X_{ic}^s - \bar{X}_c^s) \, \beta_s)$$

The estimator for $Var(\epsilon_{ic})$ is the sample variance of the residuals in Eq. 32 for women born in Jul–Sep 1967. The estimator for $Var((X_{ic}^s - \bar{X}_c^s) \, \beta_s$ is computed in two steps. First, $\widehat{\beta}_s$ is obtained from regressing (32) using Censuses 1992, 2002, and 2011. Second, this estimated vector of coefficients is used to build the single index $(X_{ic}^s - \bar{X}_c^s) \, \widehat{\beta}_s$ in Census 1977 and compute its sample variance. Then, the estimator for the intergenerational transmission of preferences specified in Eq. 17 is

$$\widehat{\rho} = \frac{\widehat{\beta}_\lambda}{\widehat{\sigma}^s} \quad (33)$$

$$s.t. \ \widehat{\sigma}^s = \sqrt{\widehat{Var}(\widehat{\epsilon}_{ic}) - \widehat{Var}((X_{ic}^s - \bar{X}_c^s) \, \widehat{\beta}_s)} \quad (34)$$

where the two sample variances on the right-hand side of expression (34) are computed as previously explained. Notice that using only women born Jul–Sep 1967 to compute $\widehat{\sigma}^s$ have the advantage not only of eliminating the inverse mills ratio from the residuals in Eq. 32, but also the usual heteroscedasticity that arises in standard selection models (Greene 1981).

Appendix: B

Tables 6 and 7 show the descriptive statistics from the 1977 Census, when second-generation individuals born at the policy cutoff, i.e., in June 1967, were 9 years old and living with their parents.

Table 6 Demographics and house characteristics (Census 1977 - 9 years old)[a]

| | Basic demograph. | |
	Mean	s.d.
Female	0.49	(0.50)
Mother tongue: Romanian	0.90	(0.31)
Num. of older siblings	0.88	(1.07)
Tot. brothers and sisters	2.29	(1.95)
Father at home	0.95	(0.22)
Urban residence	0.41	(0.49)
	House charact.	
House ownership	0.74	(0.44)
Age of structure	21.13	(21.22)
Rooms in dwelling	2.37	(0.96)
Live area - sq. meters	41.01	(16.84)
Kitchen facility	0.90	(0.30)
Bathing facility	0.27	(0.45)
Piped water	0.29	(0.46)
Water heater	0.21	(0.41)
Electricity	0.87	(0.33)
Stories in structure		
Less than 4 stories	0.81	(0.39)
4 stories	0.15	(0.35)
More than 4 stories	0.05	(0.21)
Const. material		
Concrete	0.29	(0.45)
Wood	0.38	(0.48)
Adobe or clay	0.34	(0.47)
Fuel for heating		
Solid fuel	0.72	(0.45)
Gas	0.13	(0.34)
Other	0.15	(0.36)
Obs.	327,897	
Birth counties	38	

[a] It corresponds to the age at the anti-abortion policy cutoff in 1977 of individuals born in June 1967 . The entire sub-sample in the table contains all individuals born between January 1962 and December 1972. The ages of individuals in this sub-sample range from 4 to 14 years old

Table 7 Characteristics of the father and the mother (Census 1977 - 9 years old)[a]

	Father charact.		Mother charact.	
	Mean	s.d.	Mean	s.d.
Age when child was born	29.98	(6.38)	26.14	(6.27)
School attendance	0.03	(0.17)	0.02	(0.14)
Illiterate	0.02	(0.12)	0.04	(0.19)
Employed	0.96	(0.19)	0.80	(0.40)
Education				
Less than primary completed	0.50	(0.50)	0.56	(0.50)
Primary completed	0.22	(0.41)	0.27	(0.45)
Secondary completed	0.26	(0.44)	0.15	(0.35)
University completed	0.03	(0.16)	0.02	(0.15)
Industry (if employed)				
Agriculture	0.34	(0.47)	0.58	(0.49)
Manufacturing	0.34	(0.47)	0.20	(0.40)
Other industry	0.32	(0.47)	0.22	(0.41)
Sector (if employed)				
Cooperative(	0.18	(0.38)	0.49	(0.50)
Public sector	0.80	(0.40)	0.41	(0.49)
Private sector	0.03	(0.17)	0.09	(0.29)
Occupation (if employed)				
Blue-collar worker	0.84	(0.36)	0.84	(0.37)
Obs.	311,523[b]	327,897		
Birth counties	38	38		

[a]It corresponds to the age at the anti-abortion policy cutoff in 1977of individuals born in June 1967. The entire sub-sample in the table contains all individuals born between January 1962 and December 1972. The ages of individuals in this sub-sample range from 4 to 14 years old

[b]5% of fathers not living in the household

Table 8 shows the descriptive statistics for relevant variables in Censuses 1992, 2002, and 2011. I use these Census data to analyze the fertility choices of second-generation individuals at different points in their lives (ages 24, 34, and 44). I separately study men and women because the information available for women is more accurate than those of men. Only women reported the number of children ever born and the age when they first got married.

Table 8 Own characteristics at different ages in adulthood

	Census 1992 (24 years old)[a]		Census 2002 (34 years old)[a]		Census 2011 (44 years old)[a]	
	Mean	s.d.	Mean	s.d.	Mean	s.d.
			Female sample			
Children ever born	0.93	(1.11)	1.59	(1.28)	1.70	(1.31)
Never married	0.39	(0.49)	0.27	(0.44)	0.27	(0.45)
Age at first marriage	19.15	(2.33)	20.94	(3.73)	22.63	(5.09)
Less than primary completed	0.03	(0.18)	0.04	(0.19)	0.03	(0.18)
Primary completed	0.30	(0.46)	0.25	(0.43)	0.21	(0.40)
Secondary completed	0.62	(0.48)	0.61	(0.49)	0.57	(0.49)
University completed	0.04	(0.19)	0.10	(0.30)	0.19	(0.39)
Obs.	163,531		157,326		142,712	
			Male sample			
Own children in household	0.42	(0.78)	1.16	(1.08)	1.12	(1.13)
Less than primary completed	0.03	(0.17)	0.04	(0.18)	0.03	(0.17)
Primary completed	0.25	(0.43)	0.19	(0.40)	0.17	(0.38)
Secondary completed	0.68	(0.47)	0.68	(0.47)	0.66	(0.47)
University completed	0.03	(0.18)	0.09	(0.29)	0.14	(0.35)
Obs.	165,229		157,232		146,072	

[a]They correspond to the ages at the anti-abortion policy cutoff of individuals born in June 1967 in each Census year. The entire sub-sample in the table contains individuals born between January 1962 and December 1972. The ages of individuals in this sub-sample range from 19 to 29 in Census 1992, 29 to 40 in Census 2002, and 38 to 49 in Census 2011

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