The Inter-generational Fertility Effect of an Abortion Ban*

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Abstract

This paper studies to what extent banning women from aborting affected their children's fertility, who did not face such legal constraint. Using multiple censuses from Romania, I follow men and women born around the mid-1960s Romanian abortion ban to study the demand for children over their life cycle. The empirical approach combines elements of the regressions discontinuity design and the Heckman's selection model. Results indicate that individuals whose mothers were affected by the ban had a significantly lower demand for children. One-third of such decline is explained by inherited socio-economic status.

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1 Introduction

In 2015, 42% of countries in the world had active policies to lower fertility rates, while 28% had implemented measures to achieve opposite results.[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 the policies designed to influence them has enormous implications. However, most of the research on birth control policies is limited to their contemporaneous effect, ignoring the consequences that such policies may have on the fertility of future generations. For example, Bailey (2010) studied how the birth control pill accelerated the post-1960 U.S. fertility decline. Miller (2010) analyzed the role that ProFamilia, a program that provided IUD devices to married women, had on the Colombian demographic transition, and Gertler and Molyneaux (1994) assessed the contribution of contraceptives to the Indonesian fertility decline during 1980s.

The current paper complements this literature by presenting evidence that population policies can have long-lasting effect that extend beyond one generation. Specifically, this paper shows that the abortion ban implemented in Romania in 1966 not only affected the fertility of women directly constrained by the legal measure but also shaped the demand for children of the next generation 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 to their children.

Measuring the intergenerational transmission of fertility and its determinants is important to understand 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)), negatively affecting the future living standards of their offspring. If, in addition, the children inherit the reproductive behavior of their parents, then the decline in standards of livings is likely exacerbated. Notably, the empirical literature on economic mobility has widely ignored fertility as a transmission channel. For example, the review by Black and Devereux (2010) mentions no paper on this topic. Additionally, the persistence of fertility across generations

[1] United Nations, World Population Policies Database. https://esa.un.org/poppolicy/about_database.aspx

may shed light on the patterns and speed of the demographic transitions.

In 1966, Romania banned abortion and other forms of fertility control methods (Dethier et al. (1994), Pop-Eleches (2006)). The sudden implementation of this policy took by surprise a group of pregnant women who would have aborted had the anti-abortion decree not been implemented. These women and their partners had different characteristics in relation to couples who would have not aborted even if they had the chance to do it. Thus, the abortion ban significantly changed the composition of families having children and, through this, the background of second-generation individuals.[2] The role played by these background characteristics on fertility can be obtained by comparing the reproductive behavior of second-generation men and women born around the policy implementation. This approach has strong similarities with the regressions discontinuity design (RDD), where the running variable is the date of conception and the cutoff is the moment when the anti-abortion decree was implemented. However it is conceptually different. While the selection into pregnancy of first-generation women is plausibly similar around the policy cutoff, the selection into childbearing was substantially distorted due the abortion ban.

First-generation individuals are likely to affect the fertility of the second generation through a variety of channels. One of them is the inherited 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 to quantify its relative importance. The data contain a rich set of variables associated with the socio-economic status of families, including housing characteristics, locality of residence as well as education, industry, and occupation of each household member. Although it is impossible to claim with certainty that these variables are sufficient to account for SEC differences, it worth mentioning that Romania was under a socialist regime. Little dispersion, if any, in living standards is expected after controlling for the observables just mentioned. The 1960s Romanian mechanisms to allocate labor and determine wages described in section 4 support this statement.

In addition to changing the composition of families having children and their socio-economic

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

status, the abortion ban may have directly affected second-generation individuals through altering their preferences and the nurturing behavior of their parents. For example, "unwanted" children may be less loved and neglected. These an other mechanisms are discussed throughout the paper.

The data used in this paper are the 1977, 1992, 2002 and 2011 Romanian censuses. The joint analysis of different datasets permits 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). The main results indicate that the fertility correlation across successive generations ranges from 0.15 to 0.25; of these values, one-third corresponds to the intergenerational transmission of observed socio-economic status. Results are robust to different econometric specifications and across datasets that span the entire reproductive life of women. Second-generation men also appear to be affected by the anti-abortion policy, which is evidenced in their reproductive behavior. While results for women appears to be larger in magnitude, it is not possible to know if the impact on daughters was stronger or if they had better intra-household bargaining power when they reach adulthood in relation the couple's demand for children.[3]

This paper directly relates to at least three areas of research. Firstly, It helps to better 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), Gertler and Molyneaux (1994), Pop-Eleches (2010)). Papers that analyze the effect of contraceptives on future generations usually 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 III and Levitt (2001)). Only Ananat and Hungerman (2012) discusses the impact of the contraceptive pill on the fertility of second-generation individuals in the U.S. In her case, the pill had negligible long-run effects on the demand for children.

Secondly, 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

[3]A couple's observed number of children depends on preferences and economic conditions of each member, and on the relative bargaining power of them. This paper cannot disentangle these determinants.

valuable information, they rarely attempt to disentangle the role played by inherited wealth from other determinants.[4]

Thirdly, this paper contributes to the *cultural transmission* literature. Since culture is defined as *"The integrated pattern of human knowledge, belief, and behavior that depends upon the capacity for learning and transmitting knowledge to succeeding generations"* (Merriam-Webster dictionary), the intergenerational transmission of attitudes towards childbearing is considered part of it.[5] Papers in this literature separately identify preferences from other fertility determinants by analyzing the behavior of U.S. immigrants (Fernandez and Fogli (2009), Blau et al. (2013), Guinnane et al. (2006)). Since preferences, but not economic and institutional constraints, are portable when people migrate, the statistical association between immigrants' fertility and the fertility rate in their country of origin is arguably explained exclusively by culture. That is, the fact that second-generation migrants from high-fertility countries have relatively many children at the destination is attributed to the influence of attitudes towards childbearing from the society where their parents were raised.

The cultural transmission literature uses a clever identification strategy and finds very interesting results. However, it also has some limitations: i) migrants are likely not representative of the population in the country of origin. Thus, it is not clear how much of the observed correlation is due to a cultural transmission and how much due to migrants' self-selection, ii) most of these studies use second-generation migrants.[6] Then, the impact of the country of origin culture is diminished by the degree of assimilation to the U.S. culture, and iii) since the identification of the cultural transmission of preference is obtained by comparing immigrants from different countries of origin, the influence of the socio-economic status of immigrants' parents cannot be completely eliminated because it requires the difficult task of making cross-country comparisons of standards of living.[7] It is worth emphasizing, however, that despite the limitations that may bias the magnitude of the relationship of interest, the hypothesis that culture

[4] 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.

[5] Fernandez and Fogli (2009) cite the same definition.

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

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

matters for key economic variables 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 determined wages. This section 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 the intergenerational fertility transmission. Section 10 analyzes the demand for children of second-generation men. Since the total number of children ever fathered is not available, 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.

2 Why banning first-generation individuals from aborting is expected to affect the fertility of the second generation

The economic theory of fertility (Becker and Lewis (1973), Willis (1973)) indicates that the demand for children is determined by both economic constraints and preferences. Prohibiting the first generation (i.e. the parents) to abort 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. Before abortion was banned, women who decided to terminate their pregnancies where likely to have a relatively low ‘taste’ for children and possibly a different socio-economic status that women who carried their pregnancies to term. Thus, banning abortion is expected to increase the proportion of individuals in the second generation whose parents had preferences for small families and change the average economic conditions in which second-generation individuals were raised. Both wealth and preference can be transmitted 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 are compromised. As Becker and Lewis (1973) and Willis (1973) indicate in their quantity-quality trade-off fertility theory, larger families likely allocate fewer resources per child. Additionally, young parents are in early stages of their life-cycle earnings profile being relatively more tightly constrained in the resource allocation.

Third, if banning abortion substantially increases the size of a cohort, as is this case, then there may be a crowding effect, Pop-Eleches (2006). Some public resources, such as schools, may have experienced congestion compromising the formation of human capital. The crowding effect is difficult to identify. Thus, the general equilibrium effect associated with the potential congestion of public goods is inevitably ignored, as is the case in other studies in the literature.

Finally, children born as a results of unwanted pregnancies may be less loved and more likely to be neglected. Although there is no strong empirical evidence supporting this statement (see section 9.3), the model in section 5 contemplates such possibility.

3 Romanian fertility policies and intuition of the research design

In 1957, Romania implemented the most liberal abortion policy in Europe (Dethier et al. (1994)). It allowed women to abort unconditionally and on demand during the first trimester of pregnancy. The public health care system provided the service at no cost (Pop-Eleches (2006)). Abortion soon became the main birth control method in the country.

By mid-1960s the Romanian population growth rate reached the lowest level among European countries at 0.61% per year (Dethier et al. (1994)). In response to the low fertility rate, President Ceausescu issued a decree in 1966 declaring abortion and other forms of family planning illegal. The decree was unexpected and imposed an immediate cessation of abortions. Pregnant women and physicians received severe penalties if they were caught violating the decree. To guarantee compliance, pregnant women were monitored every three months.

The result of this pro-natalist policy was a sharp increase in births. 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 more than 150% between February 1967 and July 1967, approximately nine month after the decree was implemented.

The empirical strategy in this paper consists of comparing the reproductive behavior of second-generation individuals born before and after the policy cutoff. These two groups of individuals faced the same institutional and aggregate economic constraints over their lives. But, on average, they were born and raised in families with different preferences and socio-economic characteristics.

It is worth emphasizing that if the policy implementation was truly unanticipated, then there was a group of women who was already pregnant when the decree was issued. These women, particularly those who were in their first trimester of pregnancy, did not have the option to abort *and* had not previously foreseen at the moment of getting pregnant that they would not be able to abort. Thus, the selection into pregnancy was identical for women who gave birth around the policy cutoff (see Figure 1), but the selection into childbearing discontinuously changed as a result of the abortion ban. Assuming that the policy was completely unanticipated simplifies the interpretation. However, a violation of this assumption does not invalidate the analysis. The empirical consequences are discussed below.

The pro-natalist policy implemented in 1966 had limited direct impact on second-generation individuals. It persisted until 1989 when the Romanian socialist regime ended. Individuals born in 1967 (i.e., sons and daughters of first-generation women who got pregnant around the implementation of the anti-abortion decree) were 22 years old, early in their reproductive life, when they gained full availability of contraceptive methods. The impact of the anti-abortion policy on the reproductive behavior of second-generation individuals was primarily indirect, which operated via changing the composition and behavior of parents.[8]

[8] Pop-Eleches (2010) shows that the end of the abortion ban had negative effect on fertility. This policy change does not threaten the identification strategy since it affected in the same way women born on both side of the 1967 policy cutoff.

4 Labor allocation and wage determination in centrally planned Romania

The empirical analysis in section 9 calculates to what extent the observable socio-economic characteristics of first-generation parents explain the fertility behavior of second-generation individuals. How accurately these characteristics account for differences in SEC? Although it is not possible to know with certainty, the answer depends on the mechanisms used to allocate labor and assign wages in Romania.[9]

In 1960s, Romania had a centrally planned economy. Anyone who was willing to work could find a job provided by the government. Workers were not involved in any job search and did not have the freedom to choose the industry, occupation or geographical location. Romanian system guaranteed most of the working age population a stable employment and income for 40 years. The allocation of workers to enterprises was determined at young ages. In some cases at age 14, before workers completed their education or training (Dethier et al. (1994)).

The absences of a decentralized labor market made impossible any process of wage negotiation at the individual level. Workers' pay was entirely determined by a set of rules specified by the central government. These rules classified workers into pay categories defined by easily observable characteristics, such as industry, education and experience. Importantly, differences in individual performance or productivity were not components of wages. However, there was a relationship between the aggregate productivity of the enterprise and its employees' wages. Each year, enterprises were committed to achieving output and quality targets. Failure to do so was penalized with a reduction in workers' compensations. As stated by Dethier et al. (1994), this system encouraged abuse by misreporting output.

The mechanisms of labor allocation and wage determination during the socialist Romania has been described as follows:

"Labor force participants [were] used to being allocated a job, not searching

[9] Ideally, the information available should be enough to form a sufficient statistic for the SEC. It is impossible to know if the observable variables satisfy this condition, but the mechanisms described suggest that they are good approximation.

for one, and having pay based on rules and seniority, not on performance or relative skill scarcities. Managers [were] also not used to deciding on employment and skill mix needed in the enterprise, negotiating with workers on working conditions and compensation, motivating staff, or rewarding workers for performance.”
p40 Dethier et al. (1994)

The dataset used in this paper, described in section 6, does not contain information about individual earnings. However, it contains detailed information about the age, education, industry, sector and occupation of each household member. Given the centrally planned labor system just described, these variables should be sufficient to account for income differences in Romania during 1960s. Additionally, the data provide extensive dwelling characteristics, including location, that complete an accurate measure of the socio-economic status of the household.[10]

5 Conceptual framework

This section jointly models fertility decisions and resource allocation with the objective of guiding the empirical strategy. The following Bellman's equation represents a woman's inter-temporal decision process. The role of other household members is non-essential at this point and hence ignored.

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

The left-hand side of equation (1) is the value function, which indicates that a woman's optimal inter-temporal utility at any point in time depends on demographic characteristics D (e.g. age, ethnicity), the number of children she has n and a stock of assets A (i.e., the state variables). The right-hand side of equation (1) indicates that her choices affect both the current utility $u(.)$ and the expectation of future utility $E[.]$. The parameter $\beta \in [0,1)$ is a time discount factor. During the current period, this woman optimally chooses her consumption level x, the amount

[10] Wage differential across workers categories was low. In 1989, before the revolution, ‘specialists’ earnings were only 14 per cent higher than those of ‘regular workers’ (Dethier et al. (1994)). Then, the assumption that accounting for the variables that determined workers' categories eliminates any significant dispersion in workers' compensation seems reasonable.

of assets $\tilde{A}$ she wants to keep for the next period, and the consumption level for each of her children q. The variable q is also referred as the average investment in children's 'quality'.

In addition to x and q, this woman derives utility $u(.)$ from the quantity of children n, which can be increased in future periods by getting pregnant and not having an abortion $(a = 0)$, or maintained at the same level by implementing a family planning method $(a = 1)$. All choices are partially determined by preferences ξ. For simplicity this preference parameter is referred as the taste for children, which is assumed to be malleable by social influences before the start of the reproductive life and heterogeneous in the population. In the utility function $u(.)$, demographic characteristics D operate as taste shifters.

Next period's expected value function $V(.)$, i.e. 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 formed by capitalized assets $A(1+r)$, where r is the interest rate, and a stochastic labor income $y(n)$, which is realized at the beginning of the period. The moments of the labor income distribution can be influenced by n. For example, women with many children may have accumulated relatively less human capital or currently work part-time resulting in below-average earnings.

The left-hand side of the budget constraint shows that resources can be used to pay for consumption goods, which price index is p_x, goods for her children, priced at p_q, assets for next period $\tilde{A}$ and family planning procedures, which are priced at p_a in the applicable case.[11] The interaction term qn, which generates non-linearities in the budget constraint, is standard in the fertility literature (Becker and Lewis (1973)). It indicates that in order to increase the average quantity of goods consumed by children in one unit, the mother must purchase n goods (e.g. one pair of shoes for each child).

[11] Even though 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.

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

The inter-temporal utility maximization (1)-(2) can be done 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 equation (3) shows. The superscript f indicates that the variables correspond to first-generation women. The second stage in the maximization process, equation (4), takes the resulting difference in indirect utilities and indicates whether having an abortion is the optimal choice.

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

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

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

Assuming standard properties of the per-period utility function $u(.)$, the partial derivatives of the propensity to abort, $g(.)$ in (3), indicate that older women and those who had previously given birth to relatively many children were more likely to abort - these two variables made women more likely to have completed the desired number of children. 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 as the quantity-quality trade-off theory suggests, (Becker and Lewis (1973), Willis (1973)). However, it is theoretically possible that the demand for children increases, and consequently the propensity to abort decreases, with income despite this interaction. Finally and more importantly, all else equal women with a relatively low taste for children were unambiguously more likely to abort.

The unexpected implementation of the pro-natalist decree in Romania forced many women who would have had an abortion otherwise to give birth. Thus, the normal selection process to abort given by expressions (3) and (4) was suddenly interrupted. As a result and given the derivatives of the propensity to abort $g(.)$ just described, first-generation women who gave birth after the policy cutoff are expected to be older, have more children, with a relatively low taste for children, and probably of 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 determined exclusively by their taste for children in relation to a threshold $\bar{\xi}_x$ as indicated in (5).

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

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) is obtained after inserting equality (3) and using the selection mechanism (4). Since $g(X^f, \xi)$ in (3) is strictly negatively sloping in ξ^f (i.e. high fertility preferences always imply a lower propensity to abort), then it can be inverted. As a result, the threshold is given by $\bar{\xi}_x = g^{-1}(X^f)$.[12]

When the pro-natalist policy was implemented, the probability of having an abortion ceased to be driven by (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, which explanation follows below.

Fertility decisions of women born around the anti-abortion decree (second-generation women) The inter-temporal fertility problem (1)-(2) does not only characterize the decision process of pregnant women in 1966 around the anti-abortion decree, but also the fertility problem of their daughters in adulthood (i.e., the second generation).

$$n(y_0^s, A_0^s, D^s, \xi^s, \mu) \qquad (6)$$

[12]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 (5).

Equation (6) is the demand for children obtained from such model. 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. Notice that 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 that is not present in the model with the intention of capturing 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.

Defining the vector of observable variables $X^s = (y_0^s, A_0^s, D^s)$ - where the subscript s denotes the second generation - and assuming that the taste for children, and the 'unwantedness' and crowding effects are additively separable, equation (6) becomes:

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

After taking the conditional expectation of (7) separately for cohorts born before and after the policy cutoff marked in Figure 1, the resulting average demand for children of second-generation women whose own mothers could have aborted them legally ($pre = 1$), and that of women whose own mother where already pregnant when the anti-abortion decree was implemented but were banned from aborting them ($pre = 0$) are as follows.

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

$$E(n|X^f, X^s, pre = 0) = h(X^s) + E(\xi^s|X^f, X^s) + \mu \tag{9}$$

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, which, as shown in (8) and (9), is only relevant for individuals born after the policy implementation.[13] Equa-

[13]There were certainly unwanted children even before the abortion ban was implemented. However, what matter for the empirical analysis is the increase in such proportion. Assuming $\mu = 0$ in pre-policy periods is innocuous.

tions (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 who were not aborted are realized in the sample. As equation (5) indicate, women who were not aborted had mothers with relatively high taste for children ($\xi^f > \bar{\xi}_x$). In contrast, equation (9) indicates that the anti-abortion decree eliminated the possibility of abortion for all first-generation women irrespectively of whether they had high or low fertility preferences. Thus, if there was full compliance of the anti-abortion decree, then $\xi^f > \bar{\xi}_x$ should be eliminated from the conditioning set in this second case.[14]

The difference between equations (8) and (9) is driven entirely by $E(\xi^s|X^f, X^s, \xi^f > \bar{\xi}_x) - E(\xi^s|X^f, X^s) - \mu$, 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 mayor role in the demand for children. The precise empirical approach is explained in subsection 9.1 after presenting the data structure. An empirical strategy to estimate $corr(\xi^f, \xi^s) \neq 0$ is in section 9.3.

6 Data and descriptive statistics

The data used in this paper come from Romanian Censuses carried out in years 1977, 1992, 2002 and 2011. For each of these Censuses, IPUMS-International (2015) made publicly available microdata for a 10% random sample. In 1977, residents of Alba and Arad counties where not included in the census.[15] To be consistent across years, people in other Censuses born in these two counties are dropped from the analysis. 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 when the first cohort affected by anti-abortion policy was born.

Tables A1 and A2 in appendix 1 show descriptive statistics from the 1977 Census, when second-generation individuals born at the policy cutoff - June 1967 - were nine years olds and living with their parents. This Census provides key information about the socio-economic

[14]The case of incomplete compliance is discussed in the empirical sections.
[15]Only 3.8% of the Romanian population resides in these counties.

status of families that is 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). The standard of living cannot be measured with income or consumption because these variables are absent in Census data. However, there is a rich set of information about housing characteristics, which indicates asset usage, as well as variables associated with the earnings 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 1960s and 1970s. Little dispersion, if any, in standards of living is expected after conditioning on these and other variables such as age, county of residence and whether the household resides in a urban area.

Table A3 in appendix 1 shows descriptive statistics for relevant variables in Censuses 1992, 2002 and 2011. These Census data are used to analyze fertility choices of second-generation individuals at different points in their lives - ages 24, 34 and 44. Men and women are analyzed separately. The information available for women is more accurate. Only women report the number of children ever born to them and the age when they first got married. The data limitation for second-generation men implies that the study of their reproductive behavior is somehow restrictive. The number of children that men ever had is proxied by the number of own children living in their house and the children ever born to their female partners in each Census. The statistical consequences of this data limitation are discussed in section 10.

7 The anti-abortion decree and the fertility of second-generation women

This section graphically analyzes to what extent the fertility of second-generation women discontinuously changed across birth cohorts as a results the anti-abortion decree issued in 1966. This pro-natalist policy directly affected the parents of these women (i.e., the first generation) but imposed little constraint on the availability of family planning methods for second-generation individuals since it ended in 1989 after the revolution.[16]

[16]As mentioned in section 3, second-generation individuals were 22 years old when the pro-natalist policies ended. Despite that the initial years of their reproductive lives were directly affected by the policy, the research

The upper portion of Figure 2 shows the average children ever born to women in each month-year birth cohort. The graphs are built such that a cross-section of women ranging from 15 to 55 years old is depicted for each Census year, covering all reproductive ages and beyond. The red vertical line cuts the horizontal axis at June-1967, the month when first-generation women firstly affected by the pro-natalist policy gave birth. The shaded areas cover an 11-year period 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 2(a) shows no apparent disruption in the life-cycle fertility profile among cohorts born around this policy cutoff. This fact may lead to the conclusion that the anti-abortion policy that affected first-generation women and changed the composition of mothers in relation to the socio-economic status and presumably the 'taste' for children had no consequences on fertility decisions of second-generation cohorts. However, women in their mid-twenties are far from completing their reproductive cycle. Fertility differences between women born before and after June-1967 may not be evident at that age.

Figure 2(b) shows the cross-sectional life-cycle fertility in 2002 when women born in June-1967 were 34 years old and had lived two-thirds of their reproductive life. Contrary to Census 1992, the disruption in the life-cycle fertility trend becomes evident in this year. This break in the trend is substantially more apparent in Census 2011 when women born in 1967 approached the end of their reproductive life.

The lower part of Figure 2 shows the relative cohort size in each census year. These graphs are similar to Figure 1 but covering a longer time span. The spike observed in the third trimester of 1967 is similar in magnitude in all years indicating that there was no significant change in the relative size of cohorts over time.

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 Figure 2 but only for women born between January-1962 and December-1972 (the shaded areas in Figure 2).

The fertility patterns observed in Figures 2 and 3 reveal that the anti-abortion policy that af-

design that exploits the discontinuity across birth cohorts is robust to this fact.

fected first-generation parents changed the average reproductive behavior of second-generation cohorts. What these figures cannot tell are the mechanisms underlying such association. On the one hand, 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 couples may have inherited or learned through socialization such preferences. Additionally, 'unwanted' children may have been nurtured differently affecting their preferences in relation to childbearing. On the other hand, couples who would have aborted had the policy not been implemented may be of different socio-economic condition. 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 Figures 2 and 3.

The inheritance of wealth competes with the intergenerational transmission of preferences and the direct psychological effect being 'unwanted' to explain the decline in the average children born to second-generation women born after June-1967. The following section describes the selection of first-generation women into motherhood as a function of observable socio-economic variables. Subsequently, section 9 uses this information to disentangle preferences from the inheritance of income and wealth.

8 The 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 their pregnancy in the last trimester of 1966 had no choice but to keep the baby and give birth by mid-1967.

Equation (3) indicates how each relevant variable is expected to affect 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 more likely to abort. Because the pro-natalist policy 'forced' these women to give birth, then the mothers of children born by mid-1967 are expected to be on average different in these dimensions.

Figure 4(a) shows the average age of the mother as a function of her child's birth date. Consistent with the model in section 5, mothers who gave birth after June-1967 were one year older. Figure 4(b) shows that the child's average number of older siblings is higher for children born after the implementation of the decree. Although this figure is consistent with the theory, the magnitude of the change after the policy implementation seems small. A plausible explanation for this fact is that the predictions of the model in equation (3) are conditional on other determinants to be fixed (i.e., a *ceteris paribus* analysis obtained from derivatives), but figure 4(b) shows the unconditional relationship. If this is the case, a better indicator of completed fertility is directly measuring whether women decided to have more children after the policy shock. Figure 4(c) shows the probability that the child was the youngest in the household in 1977. In this case, the discontinuity is clearer; 30% of children born before the implementation of the decree were the youngest in the household, this proportion raised to 46% after the implementation of the decree indicating that at least 35% of these children were born to mothers who had already achieved the optimal number of children or were considering to have only one extra child.

Variables associated with the socio-economic status of women are very relevant, not only because they are expected to have a significant role in the abortion decision of expecting mothers, but also because the SEC is likely to pass from one generation to another affecting the future fertility decisions of their daughters. In census data, the socio-economic condition of a household is measured by housing characteristics (i.e., assets) and the education and labor outcomes of adults (i.e., income).

Figures 4(e) and 4(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 more rooms and better access to utilities, such a piped water. Part but not all of these differences can be attributed to the fact that the anti-abortion decree increased the proportion of mothers living in urban areas - Figure 4(d).

Figures 4(g) to 4(l) show 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. Nonetheless, there seems to be no discontinuity around June-1967 in relation to female employment rates. Father's education and labor outcomes show similar patters. Importantly, the proportion of children living with their fathers was not significantly affected by the anti-abortion policy - Figure 4(j) - suggesting that men did not leave the household as a result of an unexpected child. Appendix 1 shows the profile of all other variables used in the analysis.

The plots in Figure 4 are all consistent with predictions of the selection process (3)-(4). However, they show unconditional relationships and should be viewed only as descriptive. The identification and estimation of the conditional relationship (3) is discussed in subsection 9.3.

9 Empirical analysis of the intergenerational transmission of fertility

The first part of this section analyzes to what extent the fertility profiles observed in figures 2 and 3 are explained by socio-economic characteristics. The second part of the section deals with the point estimation of the intergenerational transmission of fertility preference. This second part is methodologically more involving and requires extra assumptions.

9.1 The role of inherited socio-economic status on the demand for children of second-generation individuals

After imposing a linear functional form approximation, equations (8) and (9) combined lead naturally to the following regressions.

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

The dependent variable is the number of children ever born to 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 for 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 and zero otherwise. Figure 1 suggests that first-generation women who gave birth in the second trimester of 1967 were only partially affected by the policy. Figure A8 in appendix 2 reinforces this statement. For this reason, second-generation individuals born in this period are dropped from the sample in some regressions. Results are robust to this adjustment.

The variable W_c is the distance in months between the second-generation woman's date of birth and the cutoff date June 1967. Thus, the third and fourth terms in in regression (10) fit linear trends on both sided 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 in the regression is γ_{1t}. It measures fertility differences of women born on each side of the cutoff, and arbitrarily close to it.

Graphically, the results of estimating regression (10) ignoring covariates $\bar{X}_c^s$ and $\bar{X}_c^f$ are in Figure 3. Each dot in the scatter plot is the average children ever born computed for each birth cohort. The estimated coefficient $\widehat{\gamma}_{1t}$ is the discontinuity in the regression predictions (fitted values) at the cutoff June-1967.

Specification (10) resembles the regression discontinuity design (RD). However, it is conceptually different. In the standard RD, individuals who are arbitrarily close to the cutoff on each side are assumed to be similar in both observables and unobservables (i.e., no selection in the limit). Thus, the discontinuity can be entirely attributed the treatment analyzed. On the contrary, as shown in section 8, women born on each side of the cutoff in regression (10) are significantly different in relation to observables (socio-economic status) and plausibly unobservables (preferences). For this reason, the role of covariates is different in the standard RD that in regression (10). While in the first case covariates are included only to improve the statistical efficiency of the estimator (Calonico et al. (2016)), here $\bar{X}_c^s$ and $\bar{X}_c^f$ are included to purge the influence of the socio-economic status from the demand for children.[17] Although the interpretation of results is different, the standard RD estimation techniques can be used since

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

the objective is to compare averages on both sides of a cutoff. Therefore, the procedures to estimate equation (10) include OLS and local polynomial regression with triangular kernels as is common practice in the RDD literature.

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

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

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) - notice that 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 - and/or their preferences changed significantly as a result of being 'unwanted' children (μ). For simplicity, I will refer expression (11) as the aggregate role of preferences on estimates.[18] It contains the residual effect of the anti-abortion decree after eliminating the impact of family socio-economic variables.[19]

Regression (10) includes not only fertility determinants of second-generation women X^s, but also fertility determinants of their mothers X^f since 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-to-daughter correlation in the socio-economic status. A complication in the analysis is the lack of longitudinal data and the fact that the 1992, 2002 and 2011 Census samples used to compute equation (10) do not contain retrospective information about the parents of women born around the implementation of the anti-abortion decree, i.e., the vector X^f, neither socio-economic variables at the beginning of their reproductive life, X^s. However, much of this information can be obtained from the 1977 Census data at the cohort level, when women born around June-1967 where nine years old. For example, the father's education of a woman born in February 1967 is not available in Census data 2002 when she was 34-year-old. But, it is possible to obtain the average level of education of all

[18] γ_{1t} aggregates the two role played by preferences: the parent-to-child transmission of 'tastes' for children and the direct psychological impact of the policy.

[19] As previously 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. The role of this component is discussed below.

men who fathered a child born in February 1967 in the same county than this woman was born.

Generally speaking, the variables in the vectors $\bar{X}_c^f$ and $\bar{X}_c^s$ are computed as average characteristics observed in 1977 by month-year birth cohort and county of birth, and then imputed to observation in Censuses 1992, 2002 and 2011 belonging to the same birth cohort and county of birth. Because the anti-abortion decree varies only across but not within cohorts, including average values in regression (10) is as good in relation to identification (i.e. no bias is added) as including variables at the individual level if these were observed.

9.2 Results

Table 1 shows the results of estimating regression (10) for a variety of specifications. It consists of three panels that show regression outcomes for women belonging to the cohorts of interest at different points of their reproductive life. In Census 1992, women born in June 1967 - at the policy cutoff - were 24 years old. In Census 2002, member 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 life.

The cells in the table report the point estimates and associated standard errors of coefficient γ_{1t} in equation (10). Each cell is obtained from a separate regression combining 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 equation (10) is estimated excluding $\bar{X}_c^f$ and $\bar{X}_c^s$ from the specification.

The 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 is depicted in Figure 3. The magnitude of the discontinuities in the fitted values at June-1967 correspond to the intersection of column (1) and rows (a) in Table 1. These results suggest that the daughters of first-generation women who were banned from aborting in 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 1967 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. Figure A8 in appendix 2 reinforces this conclusion. It indicates that April was the fist month when the number of births were above trend. 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 is not a standard RDD (see previous sections for a discussion), the computational method is appropriate to compare cohorts born around June-1967. The OLS results and the RDD results 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 more weight on observations located away from the sample mean, being this approach more sensitive to non-linearities in the life-cycle fertility trend, see Figure 2. The smaller coefficients from the RDD methods appear more consistent with Figure 3, where the discontinuity in 1992 seems smaller than in the other two Censuses. Although the discontinuities is Censuses 2002 and 2011 are certainly larger, the plots in Figure 3 should be interpreted with caution because these graphs are plotted at different scales due to the steeper slope of the fertility profile at younger ages.

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 equation (10) including dwelling characteristics and father's earnings capacity variables - age, education, industry, sector of work and occupation (see appendix 1 for the comprehensive list of regressors). In addition, the specification includes a set of county of birth fixed effects and a urban residence indicator. These variables, measured in 1977 when second-generation individuals were on average nine years old, are expected to accurately control for the initial socio-economic condition 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 to remain after conditioning on the covariates. The coefficient of interest declines from rows (a) to rows (b) - bottom lines in each panel- but less than a third in most of the cases. If covariates accurately control for differences in the socio-economic status, then the heterogeneity in preferences is the only fertility determinant not included among regressors.

Rows (c) in Table 1 add the characteristics of the mother associated with her earing capacity - 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 results of the anti-abortion decree may have decided to stop working to take care of the child. Although it is not possible to know the initial response to an unexpected baby, Figure 4(h) suggests that there was none in Census 1977, ten years after the decree was issued. It shows no differential employment rate among fist-generation women affected by the anti-abortion policy. Although results in rows (c) should be interpreted with some caution, they reveal that the inclusion of mother characteristics adds no explanatory power. The percentage decline in the discontinuity remained virtually unchanged from rows (b) to rows (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. There are two reasons why this variable is relevant. Firstly, the fertility quantity-quality literature (Becker and Lewis (1973), Willis (1973)) stresses that larger families invest relatively little on each child. Since 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. Secondly, being born in larger families may affect the view in relation to the optimal number of children. For example, second-generation women who had many siblings may be more likely to have many children by replicating the family structure they were raised in.

Results in Table 2 show that including birth order and sibship size have a mild contribution to explain the discontinuity. Comparing results from rows (c) in Table 1 to those in the odd columns of Table 2 indicates that the estimated values for the coefficient of interest are similar.

Education and marital status of second-generation women Section 8 clearly shows that the anti-abortion decree changed the composition of mothers across birth cohorts. It has been argued that after conditioning on the socio-economic status of the family, differences in omitted preferences remain, emphasizing the 'taste' for children. Nonetheless, not only preferences in relation to the demand for children are likely to pass from parents to children. All other preferences may pass across generations potentially affecting the fertility of second-generation women. For example, Figure 4 shows that women who gave birth after the implementation of the anti-abortion policy were on average more educated. If the taste for education is passed from mothers to daughters, then second-generation women were more likely to stay in school for more years postponing marriage and childbearing. Separating the intergenerational transmission of fertility preferences from educational preferences is not possible - i.e, obtaining point identification. However, whether parents pass the taste for children to daughters can be empirically analyzed - i.e., testing the null hypothesis of no intergenerational correlation in the taste for children.

The even columns in Table 2 show results when the level of education, the marital status and the age of first marriage of second-generation women are included as regressors. These variables are certainly endogenously determined and its inclusion in the regression likely dwarfs the coefficient of interest. Notice that although a relative high taste for education may lower the demand for children, a relatively low 'taste' for children may also give more time for women to study more.

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 explained by their desire to study more and postpone marriage, then conditioning on these variable should reduce the coefficient of interest to zero. Table 2 shows that the indicator variable for being born before June 1967 is cut in half 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 explained either by a positive correlation in the taste for children across generations or by the direct change in preferences as a result of 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 the regression result in Table 2 account for it. However, if the *quality* of education was also affected, then the covariates in 2 are insufficient to eliminate such effect. A priori, it is unclear whether the quality of education is able to affect women's fertility in a way that is not related to the 'taste' for children. However, if such effect existed it is not possible to be disentangled from the aggregate role of preferences in (11).

9.3 Measuring the intergenerational correlation of preferences in the absence of direct 'unwantedness' effect

This section estimates the parent-to-child correlation of preferences under the somewhat strong assumption that the direct effect of being 'unwanted' on the demand for children is negligible (i.e., the variable $\mu = 0$ in expression (6)). The reason for making such assumption is based on the lack of evidence of an 'unwantedness' effect and on the importance of providing an estimate of preference correlation as benchmark for related studies.[20] Additionally, the aggregate fertility correlations found in this section are similar in magnitude to those found in other studies where the direct 'unwantedness' effect is not a major concern (Danziger and Neuman (1989), Murphy (1999), Kolk (2014), Murphy (2012), Murphy and Knudsen (2002)).

The idea that children born as a result of unwanted pregnancies are less loved and more likely to be neglected is sounded. But, it is difficult to find causal evidence supporting this presumption. Although the medical and public health literature documents that unwanted chil-

[20] For example, macro-economic studies performing calibration exercises.

dren 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 affect the allocation of resources between wanted and unwanted children, from the different socio-economic environment and parental values that unwanted children are exposed to. For example, Dytrych et al. (1975) study 220 individuals born in Czechoslovakia during the period 1961-1963 whose mothers requested to abort then when pregnant, but were denied to do so by the government. These individuals form the 'treatment' group of unwanted children. Acknowledging the importance of finding 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 this approach is that the self-selection of mothers to get an abortion is determined by their taste for children making the treatment and control group not comparable in relation to preferences. Despite the limitation of this study, the authors state that:

> *"The expectation that unwanted conceptions would lead inevitably to the children being unwanted proved* ***not*** *to be the case"* p165 (emphasis in original paper)

In line with the medical and public health literature, economics studies on abortion and the adult (mis)behavior of 'unwanted' children (e.g., Donohue III and Levitt (2001), Pop-Eleches (2006)) emphasize that women who are more likely to abort face a less favorable socio-economic environment. Precisely the exposure to such environment is argued to explain why 'unwanted' children perform worse in adulthood. In the current paper, if the information available in Romanian Censuses is sufficient to account for 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, while nurturing unexpected children differently is a possibility, there is no causal evidence showing that mothers engage in such behavior. Rather, the literature has documented that unwanted children are exposed to less favorable socio-economic environments, which is accounted for in the empirical analysis of section 9. Nonetheless, the estimated correlations shown below should be interpreted with caution.

The procedure The expression for γ_{1t} obtained in the previous section can be further developed to extract the intergenerational transmission of preferences affecting the demand for children. Assuming that $\mu = 0$ in (11) and the distribution of preferences (ξ^f) among first-generation women is well-approximated by a normal distribution. Then, γ_{1t} in regression (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) follows closely Heckman (1979). As long, as the fertility preferences of second-generation women can be written as a function of their mother's preferences in a linear fashion as

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

,where σ^s is the standard deviation of ξ^s and $\rho = corr(\xi^f, \xi^s)$, then expression (14) follows immediately from (13).[21] Furthermore, assuming that ξ^f is normally distributed with mean zero and standard deviation 1, then the conditional expectation in (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 equation (10) with the right hand side of (15) gives the following estimating equation.

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

$$chborn_{ict} = \beta_{0t} + \beta_{\lambda}\ pre_c\ \bar{\lambda}(\tilde{\xi}_x) + \eta_{1t} W_c\ pre_c + \eta_{2t} W_c\ (1 - pre_c) + \bar{X}_c^s\ \beta_{st} + \varepsilon_{ict} \tag{17}$$

where

$$\beta_{\lambda} = \sigma^s \rho$$

$$\lambda(.) = \phi(.)/(1 - \Phi(.))$$

The coefficient β_{λ} is the correlation of preferences across generation scaled by a 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 into motherhood. However, when $pre_c = 0$, the second term on the right-hand side disappears since the anti-abortion decree eliminated the possibility of having an abortion.[22]

In comparison to regression (10), X^f enters regression (17) only through the inverse mills ratio since 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 X^s because the socio-economic status variables that affect first-generation women 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 is an important advantage in relation to the standard Heckman model, which relies on the functional form of $\lambda(.)$ when no exclusion restriction is available.

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 not observed makes the Heckman (1979) approach infeasible. The second complication is the estimation σ^s, such that the correlation coefficient of preferences can be computed as $\widehat{\rho} = \widehat{\beta}_{\lambda}/\widehat{\sigma}^s$. Since some of the regressors are aggregated at the cohort level because the information about first-generation women is obtained from a

[22] Notice that λ is aggregated at the cohort-geographic level due to data limitations. For this reason, the direct 'unwantedness' effect μ in (11) cannot be separately identified.

different dataset, the error term in (17) contains extra variability that should be eliminated. The solution to these two complications requires a reformulation of the likelihood function and a cross-dataset variance estimation. For reasons of space, the these methodological issues are addressed in appendix 1.

Results Figure 5 summarize the results obtained for the intergenerational fertility correlation and the portion that is explained by inherited preferences - for a full set of results see Tables A5 and A6 in appendix 4. Figure 5(a) shows correlations under assumptions A1 (no anticipation) and A2 (full compliance), see section appendix 1. In this case, the intergenerational fertility correlations when women were 34 years old (Census 2002) and 44 years old (Census 2011) is between 0.12 and 0.14. After conditioning on variables associated with the socio-economic status of first-generation parents and second-generation women, the correlations are two-thirds of the unconditional ones, suggesting that the transmission of preferences across generations play a major role in parent-to-daughter fertility correlations. Figure 5(b) shows the same correlations but assuming that there were 5 aborting for each live birth in pre-policy years as stated in Dethier et al. (1994) rather than assuming full compliance of the anti-abortion law, see assumption B1 in appendix 1. The magnitudes of the correlations increases in all cases. However, the part of the correlation attributed to inherited preferences remained robust at approximately two-thirds of the total fertility correlation.

10 Second-generation men

The previous sections analyzed the fertility behavior of second-generation women. However, the demand for children of men born around the policy cutoff is also important. Men can also inherit the 'taste' for children of their families and affect the choices that they and their female partners make in relation to childbearing.

Census data do not provide information about the number of children ever fathered by men. But, there are 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. The two variables have some limitations. In the first case, a lower demand for children among second-generation men born after the policy cutoff can be confounded with their children leaving the household at younger ages. In the second case, the analysis can be performed only among men who decided to get marry. This second case is less problematic. Although 20 per cent of second-generation men report not living with a spouse in adulthood, there is no discontinuity in the marital status of men around the policy cutoff (see Figure A9 in appendix 2), which rules out the possibility that changes in the demand for children are confounded with changes in the selection into marriage.

Figure 6 shows the scatter plot for own children in the household and wife's children ever born. The plots are analogous to those in Figure 3 but for men born between January 1962 and December 1972. The discontinuity in the demand for children becomes apparent in Census 2011, when second-generation men born at the policy cutoff were 44 years old. In this Census year and contrary to previous years, the number of own children living in the household increases with second-generation father's date of birth (i.e. decreases with father's age) since children are more likely to live in a different household as they grow old. The life-cycle relationship between own children in the household and wife's children ever born is depicted clearly in Figure A9 in appendix 2.

Importantly, the empirical approach that compares individuals around the policy cutoff, as Figure 6 suggests, measures exclusively the influence that the preferences of second-generation men had on their wives' fertility, unconfounded from the direct influence of their wives' preferences. This is true as long as the proportion of second-generation men's wives does not discontinuously change around the husband's date of birth. Figure A10 in appendix 2 confirms that such discontinuity did not exist.

Table 3 is similar to Table 1 but for the sample of second-generation men. Rows (a) shows the unconditional relation, that is, the magnitude of the discontinuities in Figures6. After conditioning on variables associated with the socio-economic status, the discontinuity declines in magnitude but do not disappear, particularly in Census 2011. This result suggests that men do not only inherit wealth, but also preferences from their families, which affects the demand for children.

11 Summary and conclusions

In 1966, President Ceausescu issued a decree banning abortion and other fertility control methods. This pro-natalist policy discontinuously affected women who become pregnant in the last trimester of 1966 (the first generation) but had little direct effect on their daughters and sons (the second generation) since the abortion ban was eliminated in 1989. This paper studied the intergenerational fertility effect of this abortion ban and the mechanisms underneath the relationship of interest.

The understanding of the intergenerational fertility association has important implications. For example, it gives a better and more comprehensive view of the impact that population policies in the medium and the long run, it provides a valuable input to understand the pattern and speed of the demographic transition, and it can be used to model economic mobility and income inequality persistence.

Results indicate that the intergenerational fertility correlation ranges from 0.15 to 0.25. One-third of these correlations are explained by inherited socio-economic status. The rest can be presumably explained by preferences. After conditioning on mediators such as the education of second-generation women and their age at first marriage, results remain strong suggesting an important role of the inherited 'taste' for children. Conditioning on sibship size affects little the magnitude of results suggesting the it is not that the second generation imitated the family structure of their parents. Rather, they likely inherited the values from the norms from the previous generation.

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Figures

Figure 1: Cohorts size in 1977 Romanian census
(by month of birth)

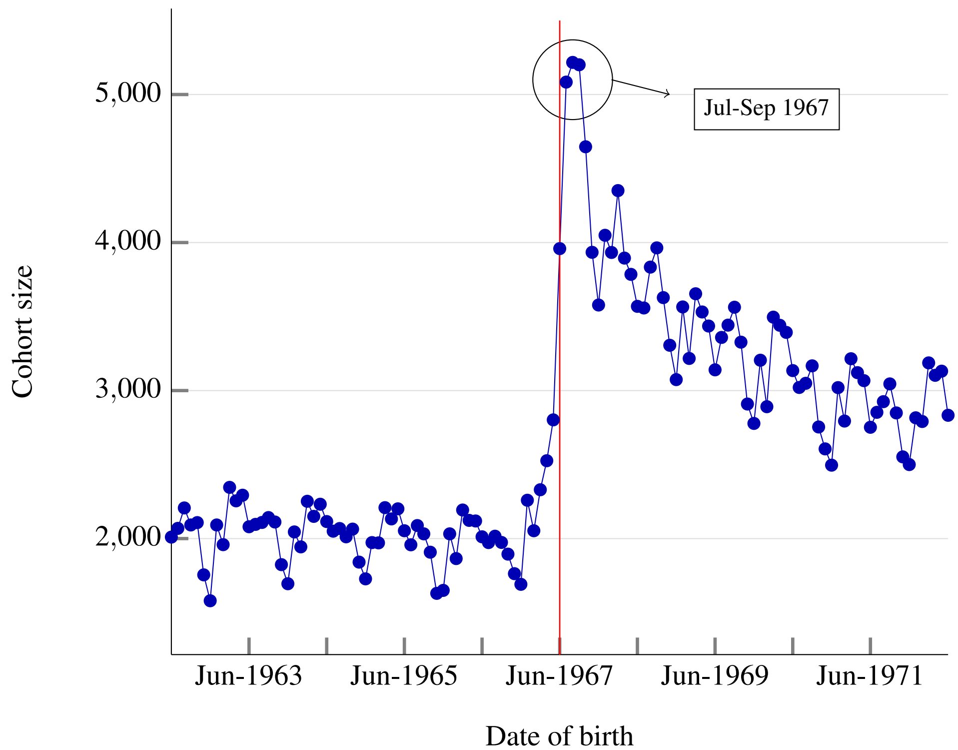


Figure 2: Life-cycle fertility, children ever born to women 15 to 55 years old by Census year

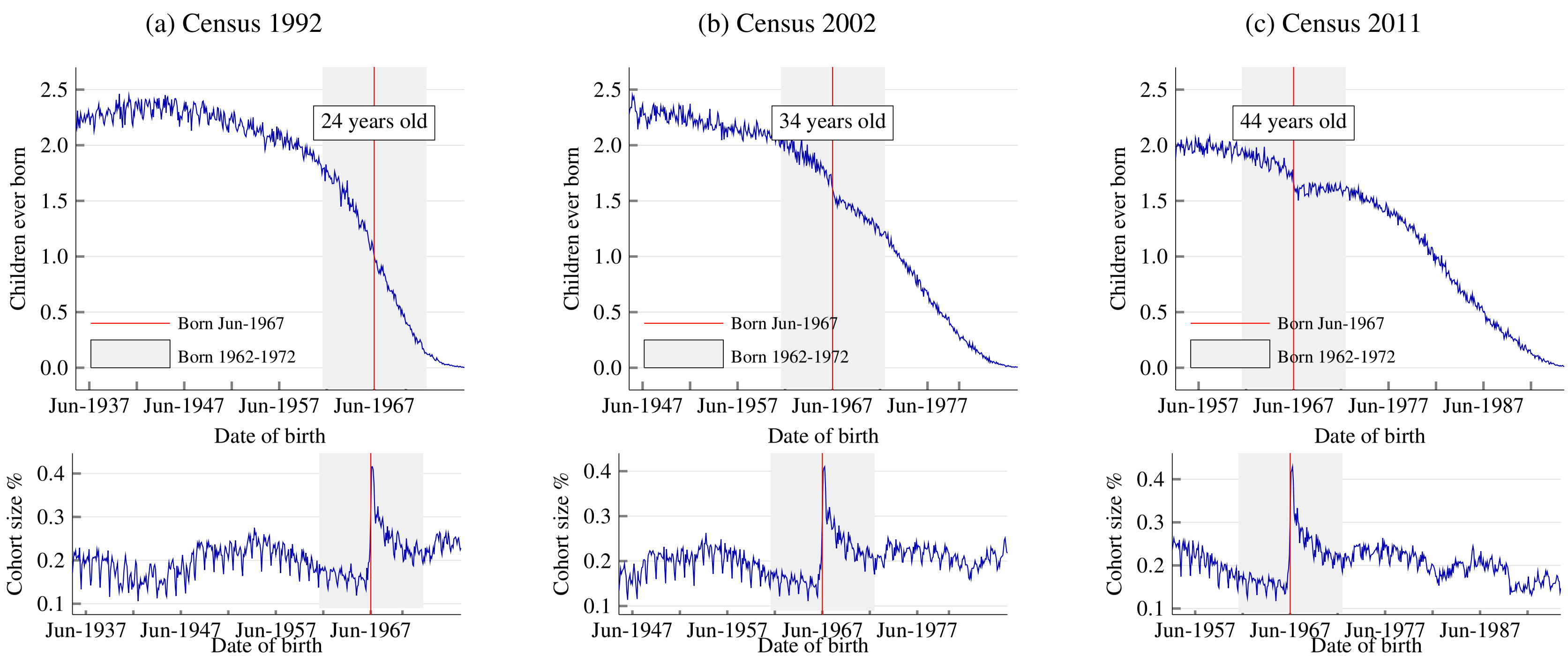


Figure 3: Children ever born to women in cohorts Jan-1962 to Dec-1972 in by Census year

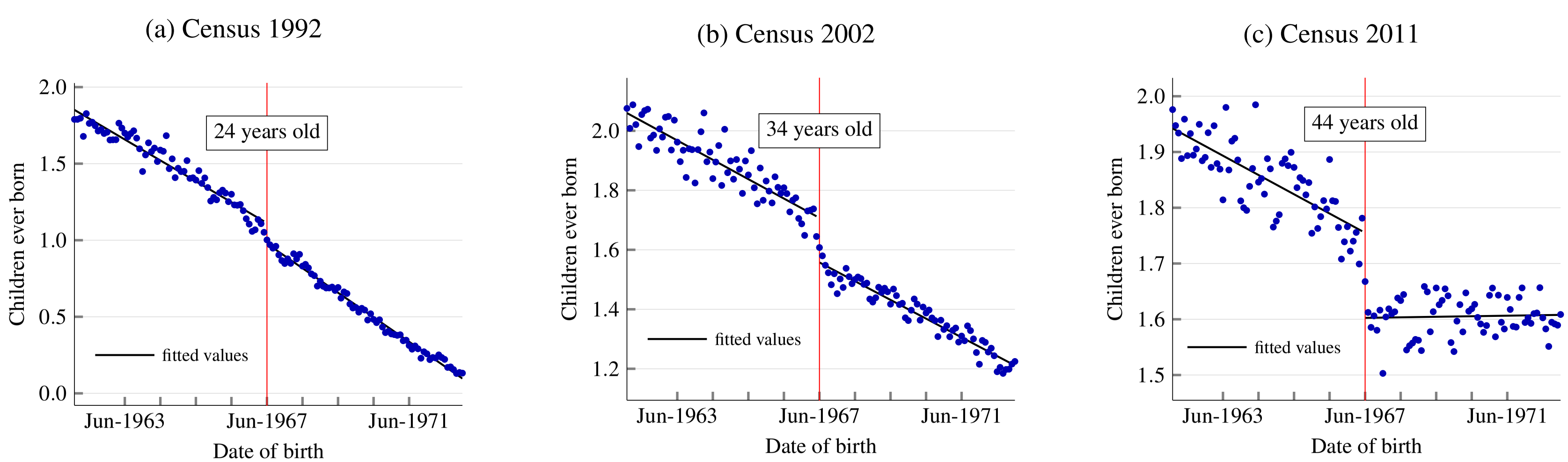


Figure 4: Selected covariates (Census 1977)

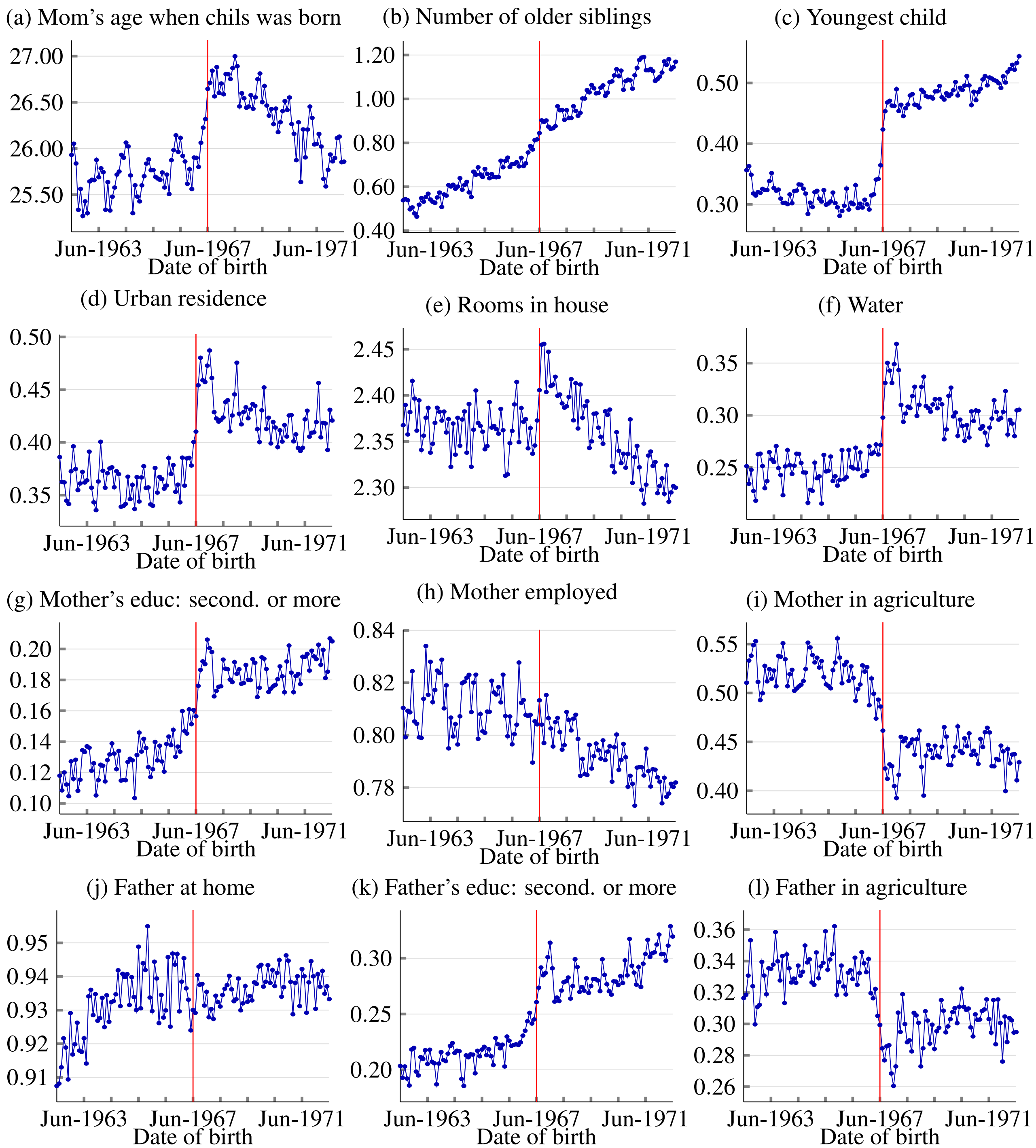
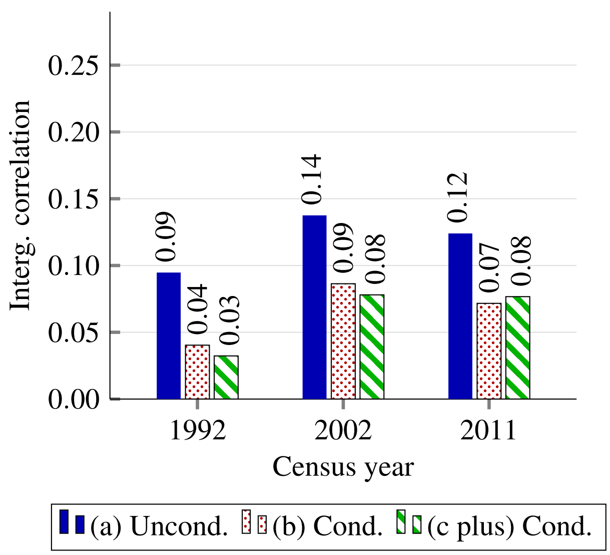
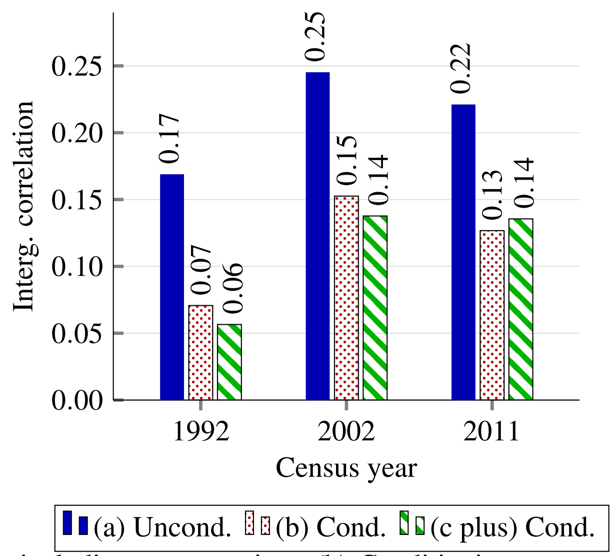


Figure 5: Intergenerational correlations

(a) Full compliance and no anticipation



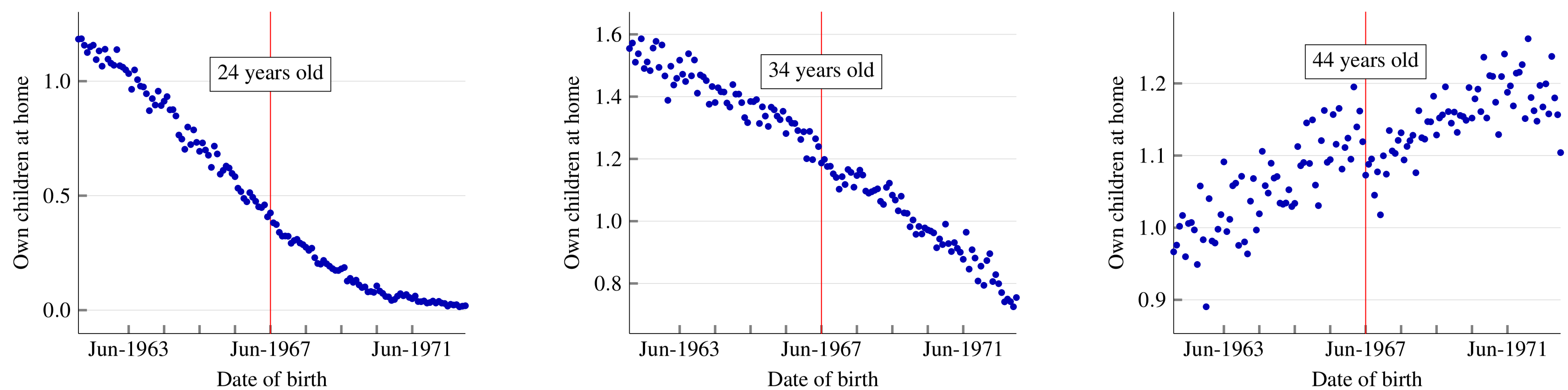
(b) Five abortions per living birth



(a) The intergenerational correlation without including any covariate; (b) Conditioning on dwelling and father characteristics as defined in the appendix plus urban residence in 1977, county of birth fixed effect and quarter of birth fixed effects, (c plus) adds mother characteristics (see appendix) to the covariates included in the previous case plus tot. number of siblings and older siblings. All results computed by RDD using triangular kernel of 36 month band width. See appendix for complete set of results.

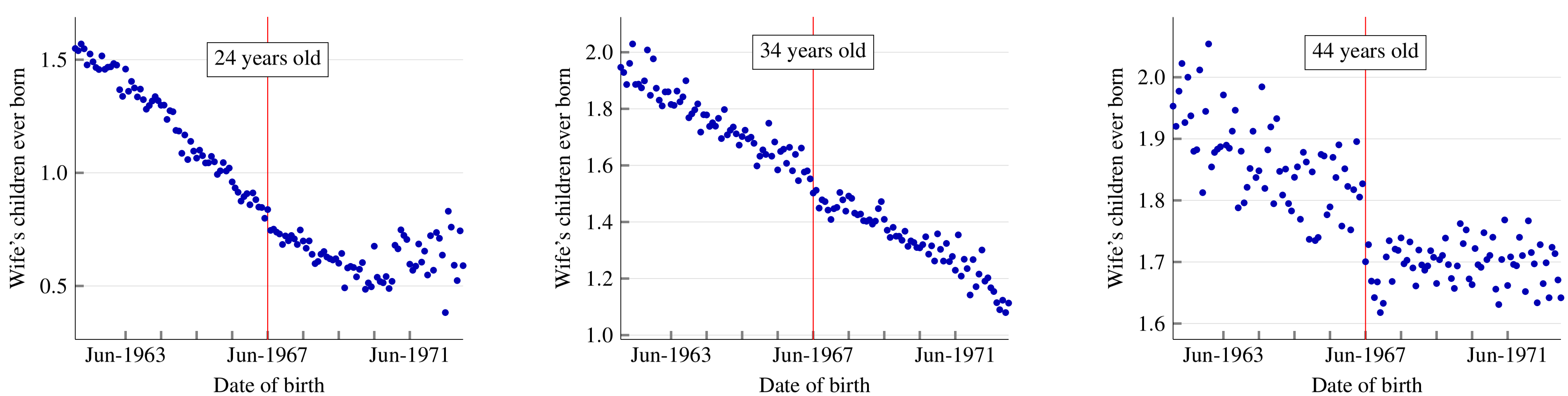
Figure 6: Own children at home (men 15 to 55 years old)

(a) Census 1992 | (b) Census 2002 | (c) Census 2011



Wife's children ever born (men 15 to 55 years old)

(a) Census 1992 | (b) Census 2002 | (c) Census 2011



Tables

Table 3: Own children in household and wife's children ever born

SAMPLE: MEN BORN BETWEEN JAN-1962 AND DEC-1972

	Census 2002 (34 years old$^{\pm}$)			
	Own children at home†		Wife's children ever born‡	
	OLS	RDD	OLS	RDD
Born pre Jun-1967 (a)	0.046***	0.045*	0.055***	0.087***
	(0.013)	(0.019)	(0.015)	(0.021)
Born pre Jun-1967 (c)	0.020	0.024	0.026	0.062**
	(0.012)	(0.017)	(0.015)	(0.021)
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$^{\pm}$)			
	Own children at home†		Wife's children ever born‡	
	OLS	RDD	OLS	RDD
Born pre Jun-1967 (a)	0.056***	0.090***	0.097***	0.148***
	(0.014)	(0.020)	(0.019)	(0.028)
Born pre Jun-1967 (c)	0.043**	0.076***	0.060**	0.119***
	(0.014)	(0.018)	(0.018)	(0.021)
Expl. by covariates [(a)-(c)]/(a)	0.23	0.15	0.39	0.20
Obs.$^{\wr}$	142,267	142,267	106,986	106,986

Clustered standard errors in parentheses

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

Note: Each cell is obtained from a separate regression. RDD uses a triangular kernel with 36 months band width. Rows (a): no covariates. Rows (c), analogously to that in Table 1, include dwelling, father and mother and characteristics as defined in the appendix plus urban residence in 1977, county of birth fixed effect and quarter of birth fixed effects. All regressions include independent linear trends on each side of the cutoff as indicated in equation (10).

± Age of individuals born in June 1967.

† Sample includes both married and unmarried men.

‡ Sample includes only men whose wife is at home.

≀ For RDD methods is the total number of observations, not the effective observations after the bandwidth selection.

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

DEPENDENT VARIABLE: CHILDREN EVER BORN
SAMPLE: WOMEN BORN BETWEEN JAN-1962 AND DEC-1972

	OLS all month-years 1962-1972	OLS excluding 2nd quar. 1967	RDD band width 24 months	RDD band width 36 months
	(1)	(2)	(3)	(4)
	Census 1992 (24 years old$^{\pm}$)			
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
Obs.†	164,209	159,894	159,894	159,894
	Census 2002 (34 years old$^{\pm}$)			
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
Obs.†	157,412	153,316	153,316	153,316
	Census 2011 (44 years old$^{\pm}$)			
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
Obs.†	142,712	138,953	138,953	138,953

Clustered standard errors in parentheses
* $p<0.05$, ** $p<0.01$, *** $p<0.001$

Note: Each cell is obtained from a separate regression. The columns indicate the estimation method and the rows the covariates included in the specification. RDD uses a triangular kernel. Rows (a): no covariates. Rows (b) include dwelling and father characteristics as defined in the appendix plus urban residence in 1977, county of birth fixed effect and quarter of birth fixed effects. Rows (c) adds mother characteristics (see appendix) to the covariates included in the previous case. All regressions include independent linear trends on each side of the cutoff as indicated in equation (10).

± Age of individuals born in June 1967.

† For RDD methods is the total number of observations, not the effective observations after the bandwidth selection.

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

DEPENDENT VARIABLE: CHILDREN EVER BORN
SAMPLE: WOMEN BORN BETWEEN JAN-1962 AND DEC-1972

	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
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
Obs.†	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
* $p<0.05$, ** $p<0.01$, *** $p<0.001$

Note: Women born in the second quarter of 1967 are excluded from the sample. Triangular kernel with 36 months band width for RDD.
† For RDD methods is the total number of observations, not the effective observations after the bandwidth selection.

Appendix 1

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 choses to give birth instead of aborting. The probit model can be written in latent variable form, which is derived from equations (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.

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

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

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 given the structure of the data available. The standard probit requires women to be classified into those who had an abortion ($a = 1$) and those who gave birth ($a = 0$). In Census data, neither abortion, nor the pregnancy status of women is observed, which makes such classification infeasible. However, the way the anti-abortion policy was implemented can be used to overcome this problem.

In any pre-policy period, it is possible to know who did not abort because the child's date of birth is observed. But, it is not possible to know who aborted because the pregnancy status of women is unknown. One the other hand, when the policy was implemented none of the women who were in their first trimester of pregnancy had the option to abort. 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 Figure 1). Thus, the unconditional probability of ***not*** having an abortion is given by the size of the cohort born in July-Sept 1966 (i.e., the number of women who decided not to abort when abortion was legal) in relation to the size of the cohort born in July-Sept 1967 (i.e., the total number of pregnant women). Formally, define the variables b as follows

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

Additionally let s be 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 prior to the implementation of the pro-natalist policy is

$$\begin{aligned} P(a_i = 0) &= \frac{\sum_{i=1}^{n} b_i}{s} & (21) \\ &= \frac{\sum_{i}^{n} b_i}{\sum_{i=1}^{n}(1-b_i)} & (22) \\ &= \frac{\sum_{i}^{n} b_i}{n} \times \frac{n}{\sum_{i=1}^{n}(1-b_i)} & (23) \\ &= \frac{P(b_i = 1)}{P(b_i = 0)} & (24) \end{aligned}$$

Equalities (21)-(24) make evident the assumptions needed to compute the probit model. The right hand side of (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 it can replaced by a similar quantity under two assumptions.

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 1 is plausible. The cohort size prior to 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 gave women with pregnancy due date in Jul-Sep 1967 no time adjust their behavior in anticipation of the policy. That is, these

[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 in relation to assumption A1 - and ii) the average trimester pregnancies in the past five years. Results are almost identical to these changes in assumption A1.

women should have not known that an anti-abortion policy was going to be implemented before they got pregnant. Otherwise, some of then could have adopted contraceptive measures in response (e.g., lower the intercourse frequency). 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 with due date in Jul-Sep 1967 effectively resulted in births - up to a normal 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 (21) can be replace by the number of births in Jul-Sep 1967 resulting in expression (22). Then, multiplying the numerator and the denominator by n in (23) gives the final result in (24): the probability of not having an abortion is equivalent to the ratio of the probability of being born in Jan-Mar 1967 to the probability of being born in Jul-Sep 1967.

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).

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

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

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

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

The specific functional form in (27) comes directly from the normality assumption of probit model (18)-(19). Replacing (27) into (25) gives the computationally 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 (17) is straightforward.

The selection probabilities obtained from the maximization of (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 having been banned (violation of assumption A2). The World Bank country report that describes the social and economic environment in socialist Romania mentions that there were five abortions for every live birth in 1965, the year before the anti-abortion decree was issued (Dethier et al. (1994) p142). 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, Figure 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

[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 (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.

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 children ever born, is observed in Censuses 1992, 2002 and 2011 when they were adults. On the other hand, the inverse mills ration $\lambda(.)$ 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.

$$chborn_{ict}=\beta_0+\beta_\lambda\ pre_c\ \lambda(\tilde{\xi}_{xi})+\eta_1 W_c\ pre_c+\eta_2 W_c\ (1-pre_c)+X^s_{ic}\ \beta_s+\xi^s_{ic} \tag{31}$$

$$\begin{aligned}&=\beta_0+\beta_\lambda\ pre_c\ \bar{\lambda}(\tilde{\xi}_x)+\eta_1 W_c\ pre_c+\eta_2 W_c\ (1-pre_c)+\bar{X}^s_c\ \beta_s+\ldots\\ &\ldots\underbrace{\beta_\lambda\ pre_c\ (\lambda(\tilde{\xi}_{xi})-\bar{\lambda}(\tilde{\xi}_x))+(X^s_{ic}-\bar{X}^s_c)\ \beta_s+\xi^s_{ic}}_{\varepsilon_{ict}}\end{aligned} \tag{32}$$

The ideal fertility equation (31) contains regressors $\lambda(\tilde{\xi}_{xi})$ and X^s_{ic} at the individual level. Adding and subtracting cohort aggregates for these covariates gives the computationally feasible equation (32). The error term ε_{ict} contains not just preferences ξ^s_{ic} but deviations of individual regressors from cohort averages. Then, since $(X^s_{ic}-\bar{X}^s_c)$ is orthogonal to ξ^s_{ic}, 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:

$$Var(\xi^s_{ic})=Var(\varepsilon_{ic})-Var((X^s_{ic}-\bar{X}^s_c)\ \beta_s)$$

[25]Under assumptions A1 and A2, the unconditional probability of not getting an abortion is 0.428 (aprox. 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.

The estimator for $Var(\varepsilon_{ic})$ is the sample variance of the residuals in (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 (17) is

$$\widehat{\rho}=\frac{\widehat{\beta_{\lambda}}}{\widehat{\sigma^{s}}} \tag{33}$$

$$s.t.\ \ \widehat{\sigma^{s}}=\sqrt{\widehat{Var}(\widehat{\varepsilon}_{ic})-\widehat{Var}((X_{ic}^{s}-\bar{X}_{c}^{s})\ \widehat{\beta_{s}})} \tag{34}$$

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 (32), but also the usual heteroscedasticity that arises in standard selection models (Greene (1981)).

Appendix 2

Tables A1 and A1 show descriptive statistics from the 1977 Census, when second-generation individuals born at the policy cutoff - June 1967 - were nine years olds and living with their parents. This Census provides key information about the socio-economic status of families that is 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). The standard of living cannot be measured with income or consumption because these variables are absent in Census data. However, there is a rich set of information about housing characteristics, which indicates asset usage, as well as variables associated with the earnings capacity of each household member such as education, employment status, industry, sector and occupation.

Table A3 shows descriptive statistics for relevant variables in Censuses 1992, 2002 and 2011. These Census data are used to analyze fertility choices of second-generation individuals at different points in their lives - ages 24, 34 and 44. Men and women are analyzed separately. The information available for women is more accurate. Only women report the number of children ever born to them and the age when they first got married. The data limitation for second-generation men implies that the study of their reproductive behavior is somehow restrictive. The number of children that men ever had is proxied by the number of own children living in their house and the children ever born to their female partners in each Census.

Table A1: Demographics and house characteristics
(Census 1977 - 9 years old)$^{(*)}$

	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.	
	mean	s.d.
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	

$(*)$ It corresponds to the age of individuals born in June-1967 - at the anti-abortion policy cutoff - in 1977. 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 A2: Characteristics of the father and the mother
(Census 1977 - 9 years old)$^{(*)}$

	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$^{\mp}$		327,897	
Birth counties	38		38	

$(*)$ It corresponds to the age of individuals born in June-1967 - at the anti-abortion policy cutoff - in 1977. 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.
$\mp$ 5% of fathers not living in the household.

Table A3: Own characteristics at different ages in adulthood

	Census 1992 (24 years old)$^{(*)}$		Census 2002 (34 years old)$^{(*)}$		Census 2011 (44 years old)$^{(*)}$	
			Female sample			
children ever born	0.93	1.11	1.59	1.28	1.70	1.31
nevermarried	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	

$(*)$ They correspond to the ages of individuals born in June-1967 - at the anti-abortion policy cutoff - 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.

Appendix 3:

Descriptive statistics (cont.)

Online material

Figure A1: Basic demographics

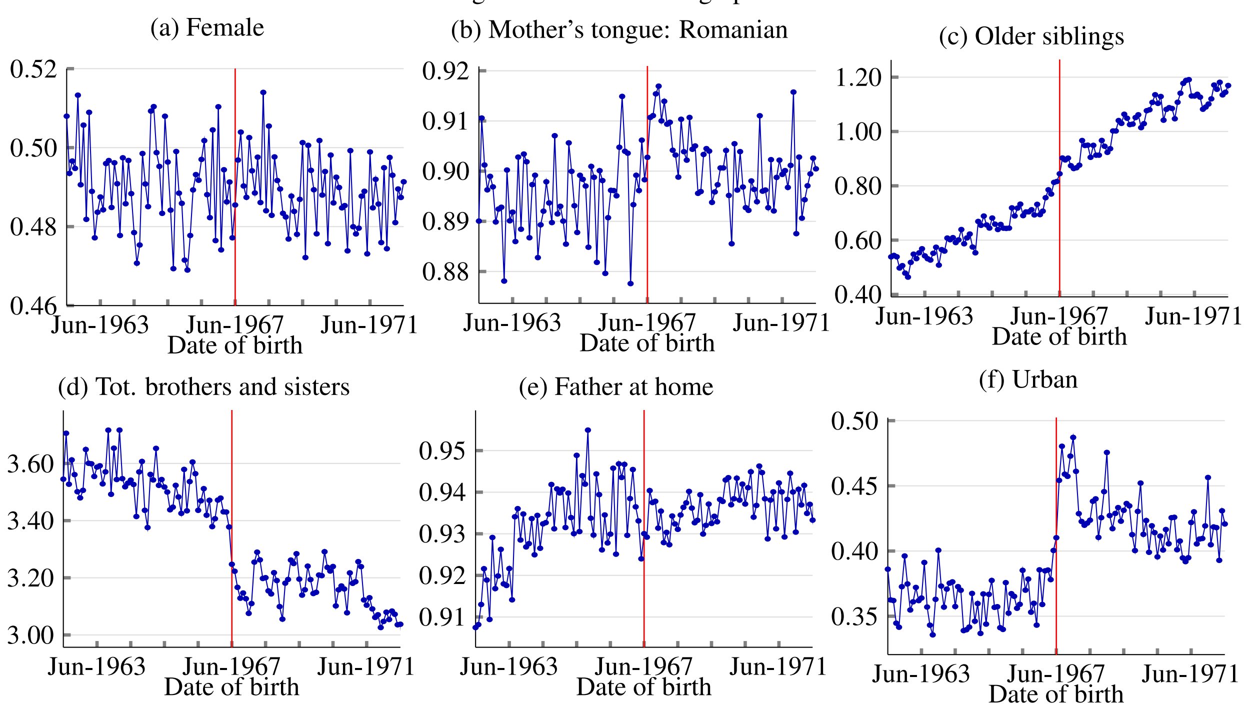


Figure A2: Housing

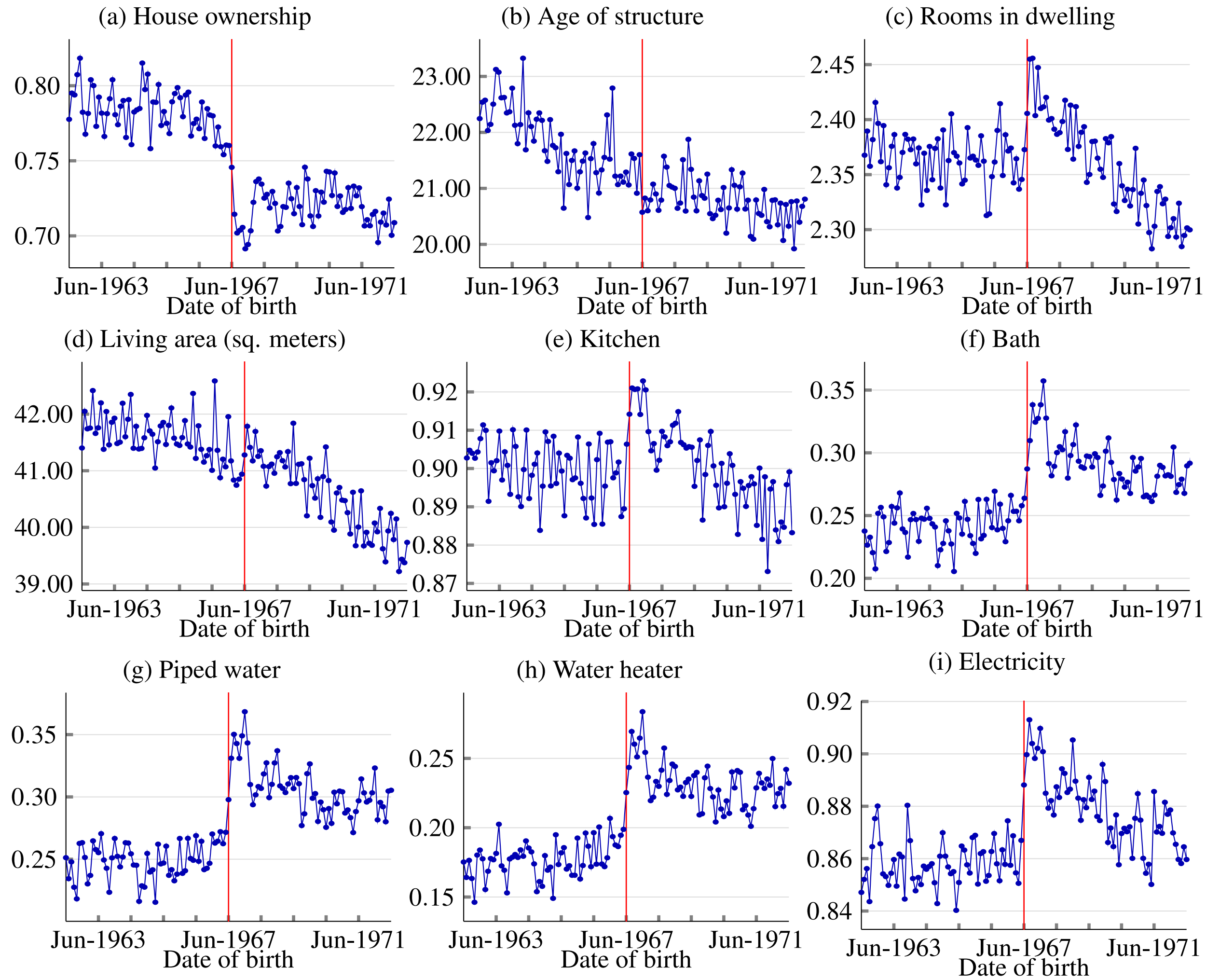


Figure A3: Housing (cont.)

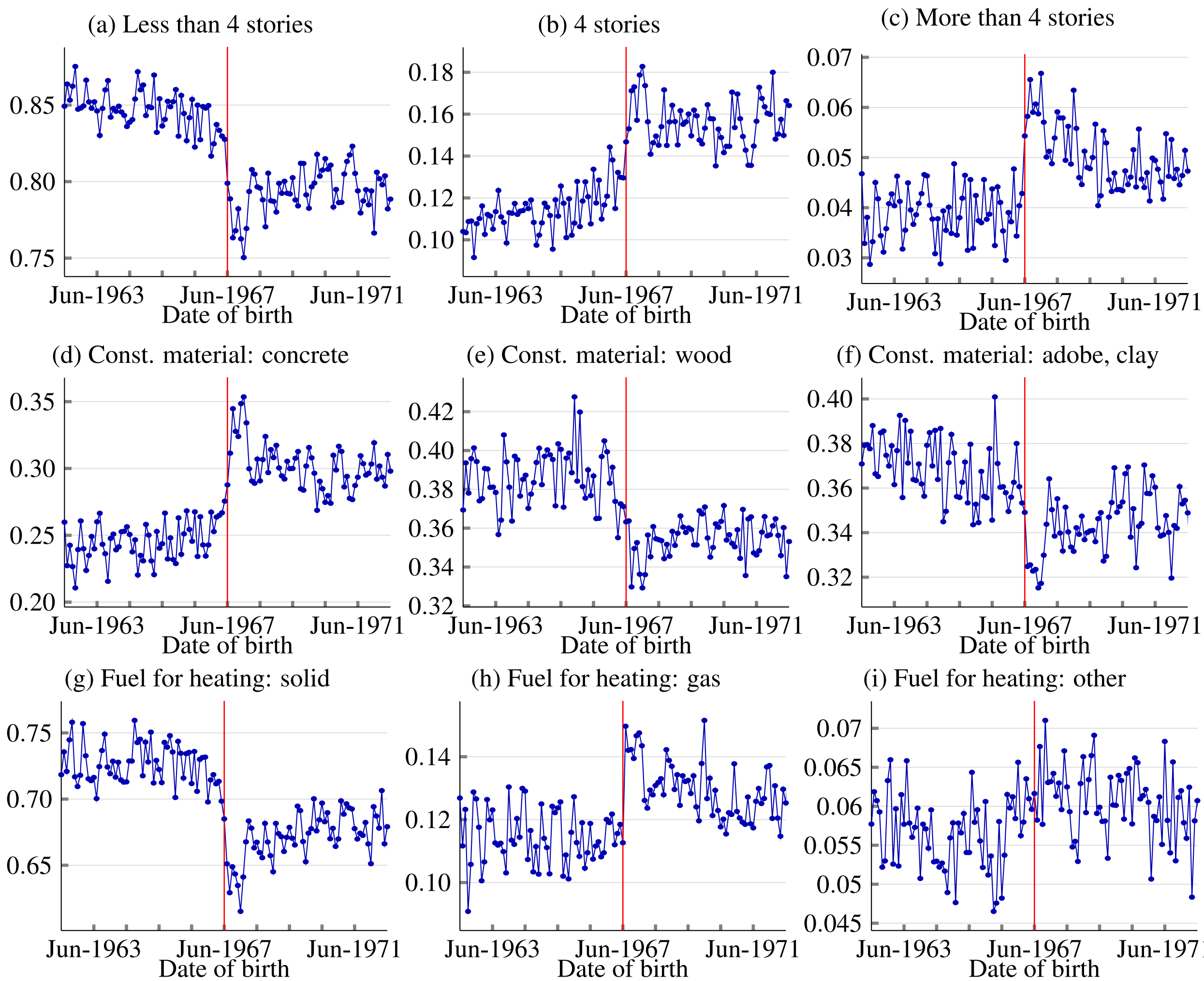


Figure A4: Father characteristics

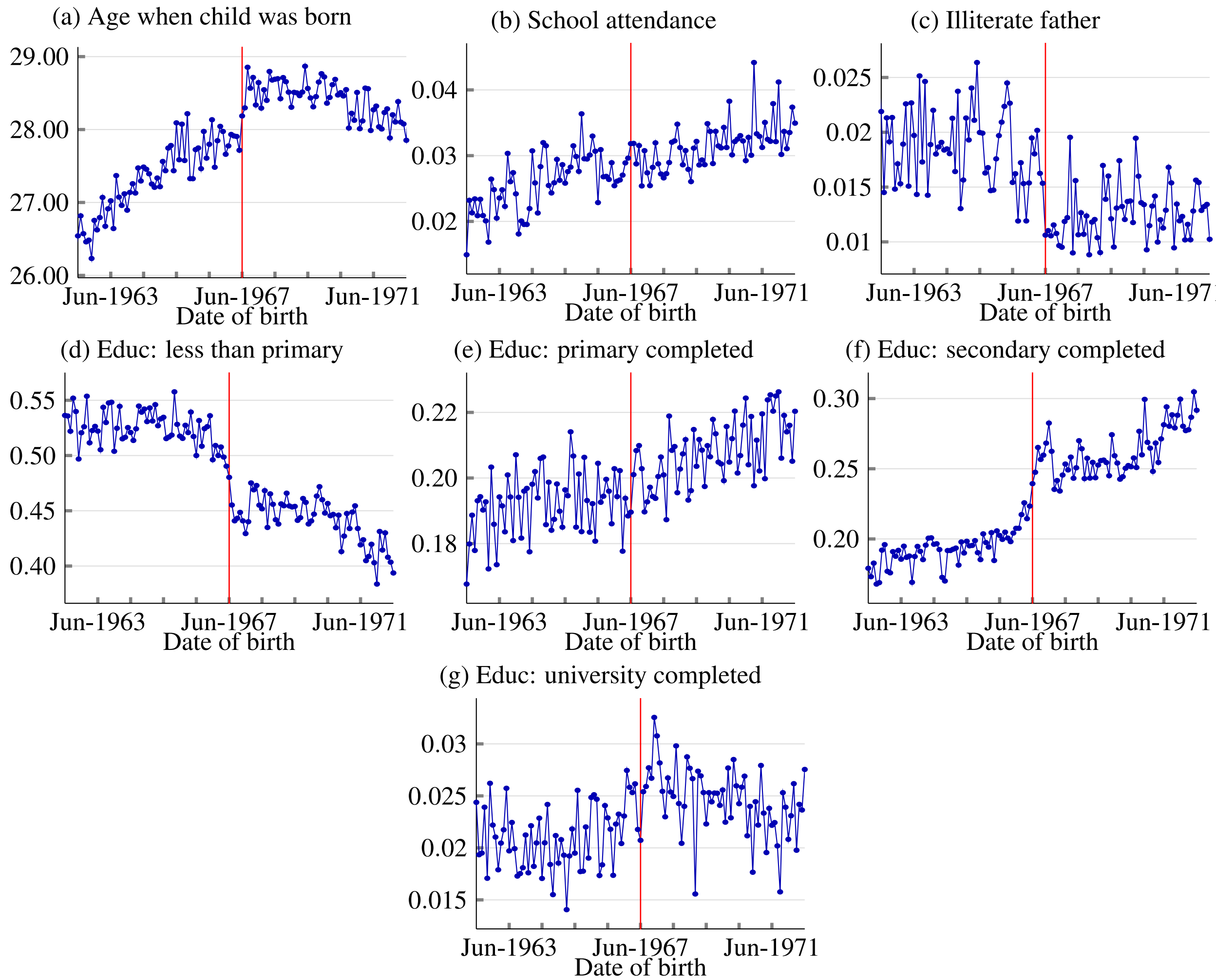


Figure A5: Father characteristics (cont.)

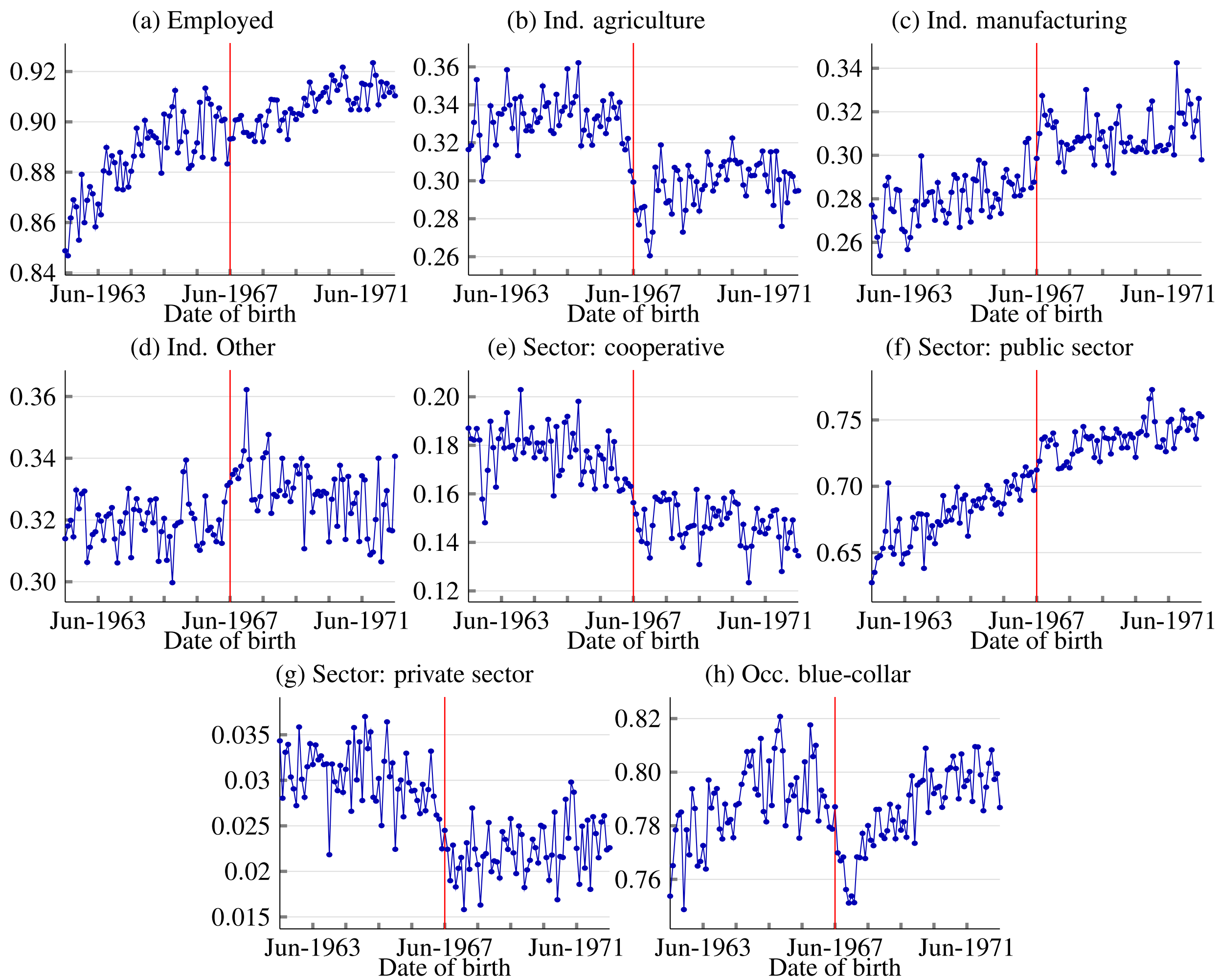


Figure A6: Mother characteristics

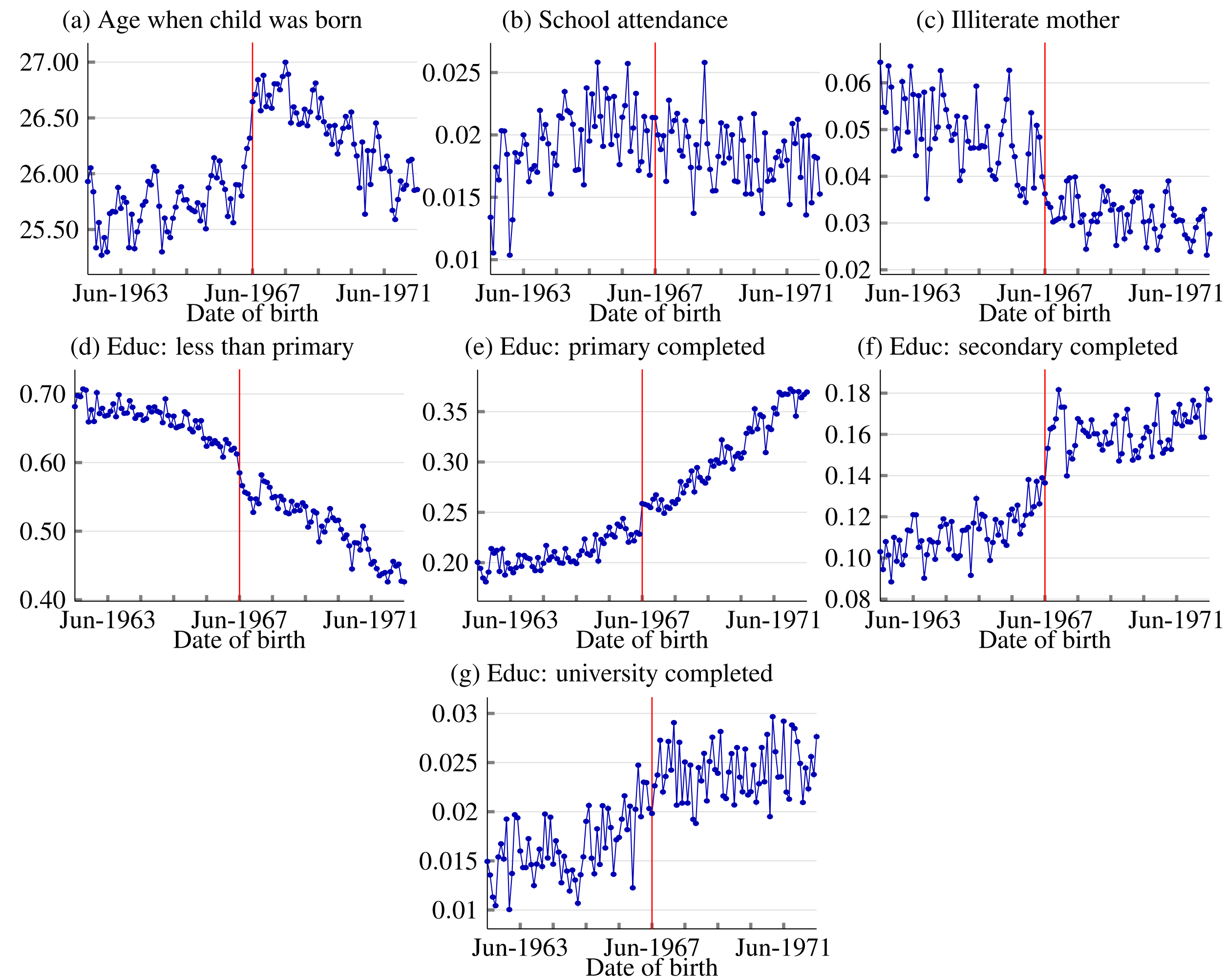
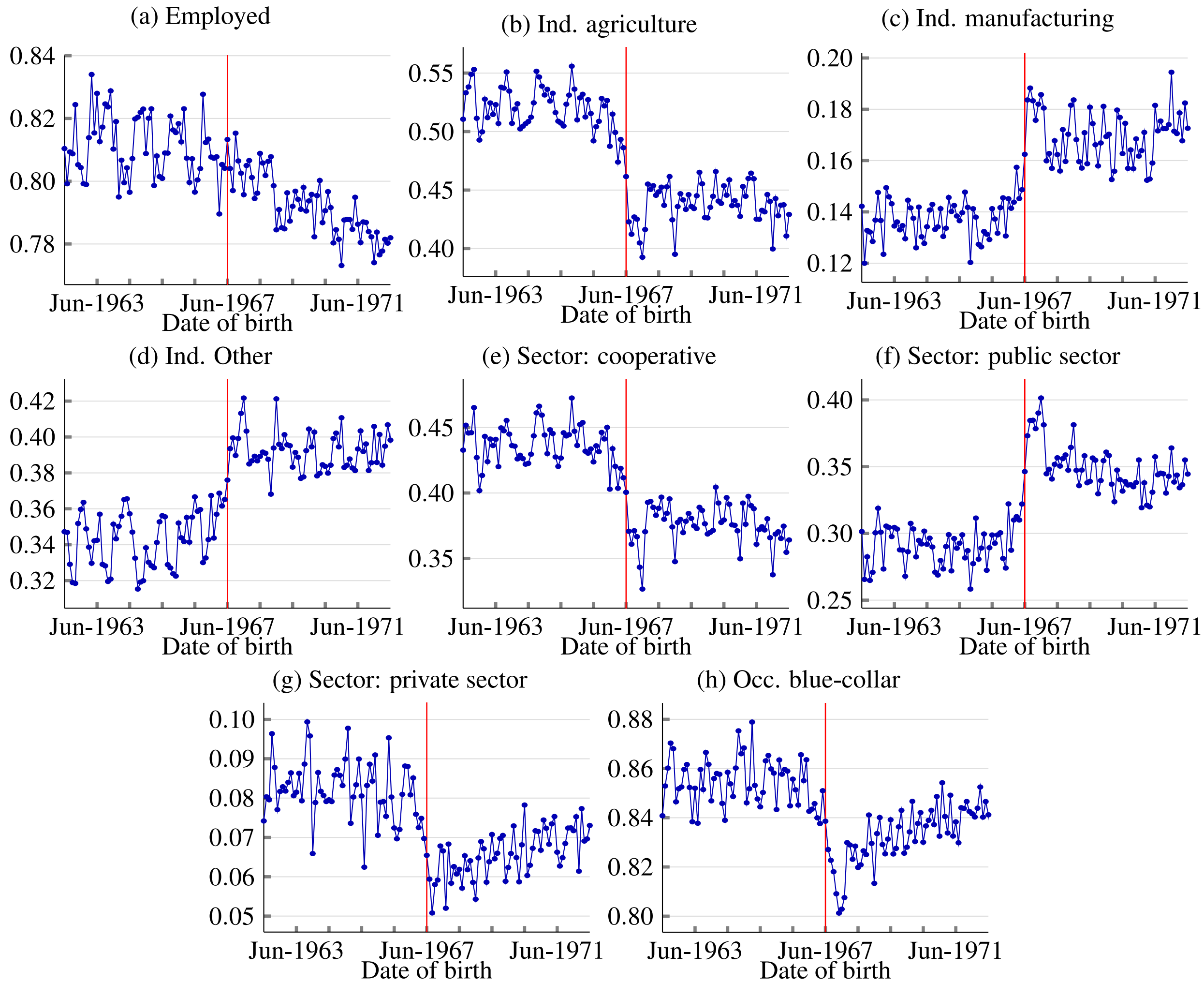


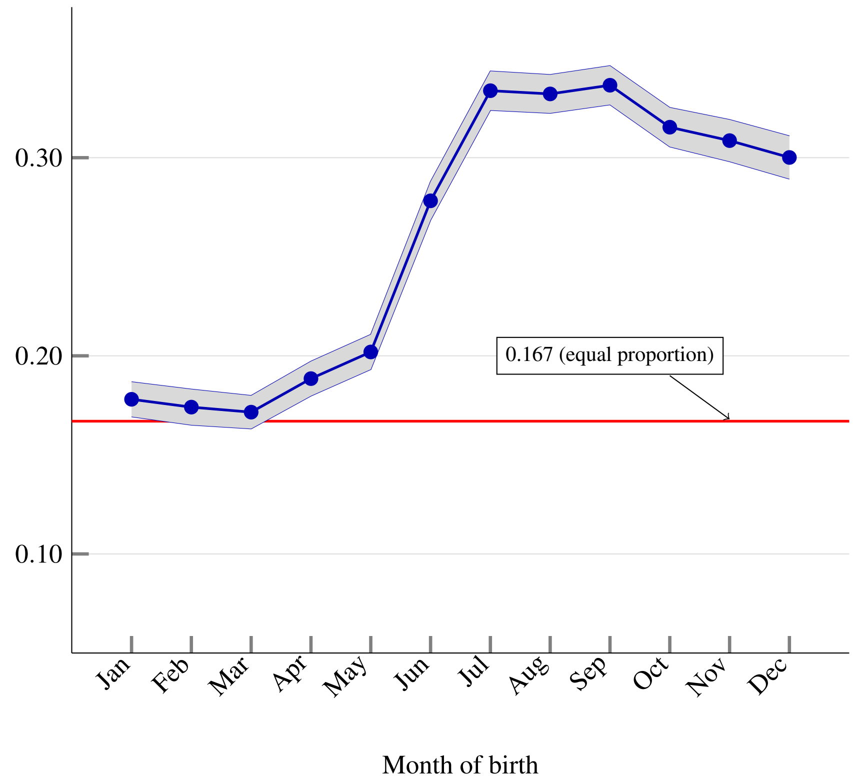
Figure A7: Mother characteristics (cont.)



Appendix 4:
Additional results

Online material

Figure A8: Proportion of birth in 1967 among all 1962-1967 births
(by month of birth)



Note: Each point is computed as the ratio of births in 1967 for a given month to total births in the period 1962-1967 for the same month. In the absence of the anti-abortion policy and under stationary conditions, births should be uniformly distributed across years at 16.7% birth in each of them.

Panel (a) in Figure A9 shows the histogram of births by mother's age in two periods of time: Jul-Sept 1966 (pre policy cutoff) and Jul-Sept 1967 (post policy cutoff). If women who gave birth in the second part of 1967 were caught by surprise by the decree, the percentage difference between the two histograms gives an accurate measure of abortion rates by women's age. Panel (b) shows that the abortion rate was increasing in women's age reaching more than 7 out of 10 pregnancies aborted among 36-year-old women.[1]

[1] The abortion rates are computed as the percentage difference in the two histograms, where each point in Figure A9 Panel (b) is computed as $\frac{births_{1a}-births_{0a}}{births_{1a}}$, where $births_{ta}$ is the number of births obtained in period $t \in \{0,1\}$ (Jul-Sep 1967 and Jul-Sep 1966) at age a.

Figure A9: Births and abortions by mother's age

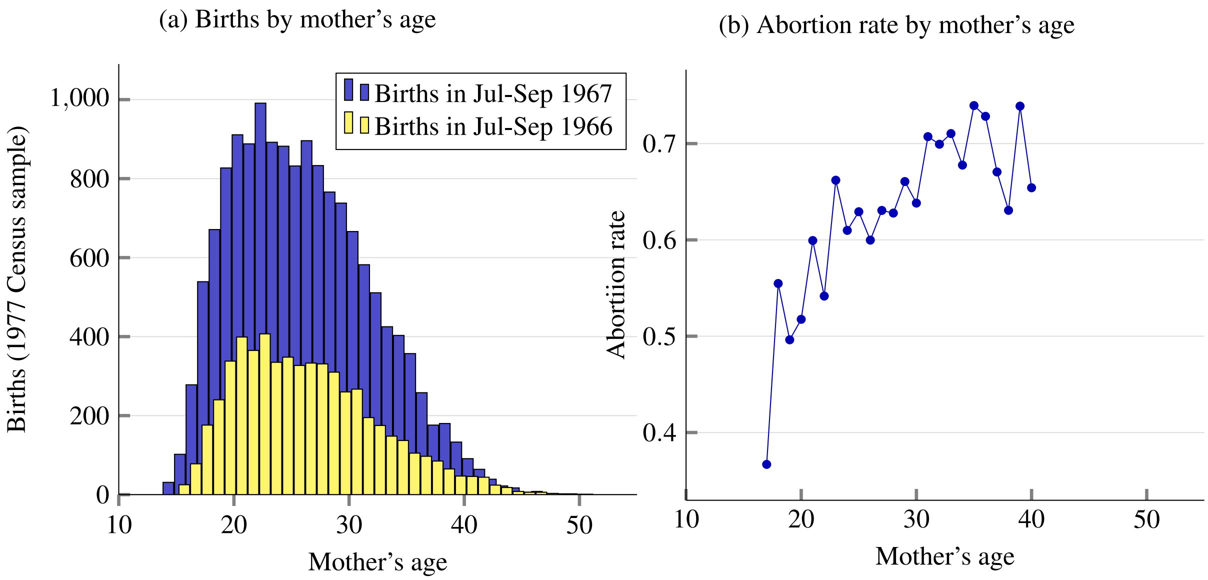


Table A4: Tests of joint significance in first stage regression:
(used to compute the inv. Mills ratio)

SAMPLE: CHILDREN BORN IN JUL-SEP 1966 OR JUL-SEP 1967	(b)	(c)
joint test: house charact.		
$\chi^2(15)$	59.73	47.61
p-val	0.000	0.000
joint test: father charact.		
$\chi^2(13)$	90.37	46.90
p-val	0.000	0.000
joint test: mother charact.		
$\chi^2(13)$		69.82
p-val		0.000
joint test: geograph. location		
$\chi^2(38)$	151.11	127.91
p-val	0.000	0.000
Obs.	20,469	20,469

Note: House characteristics are listed in Table A1. Father characteristics include variables in Table A2 plus an indicator for whether the father belongs to the household. Mother characteristics include variables in Table A2 plus and indicator for whether the mother's tongue was Romanian. Geographical location includes 37 dummies for the 38 counties of birth (Bucharest is the omitted county) and an indicator for urban residence.

Figure A10: Own children at home and wife's children ever born
(men 15 to 55 years old in each Census)

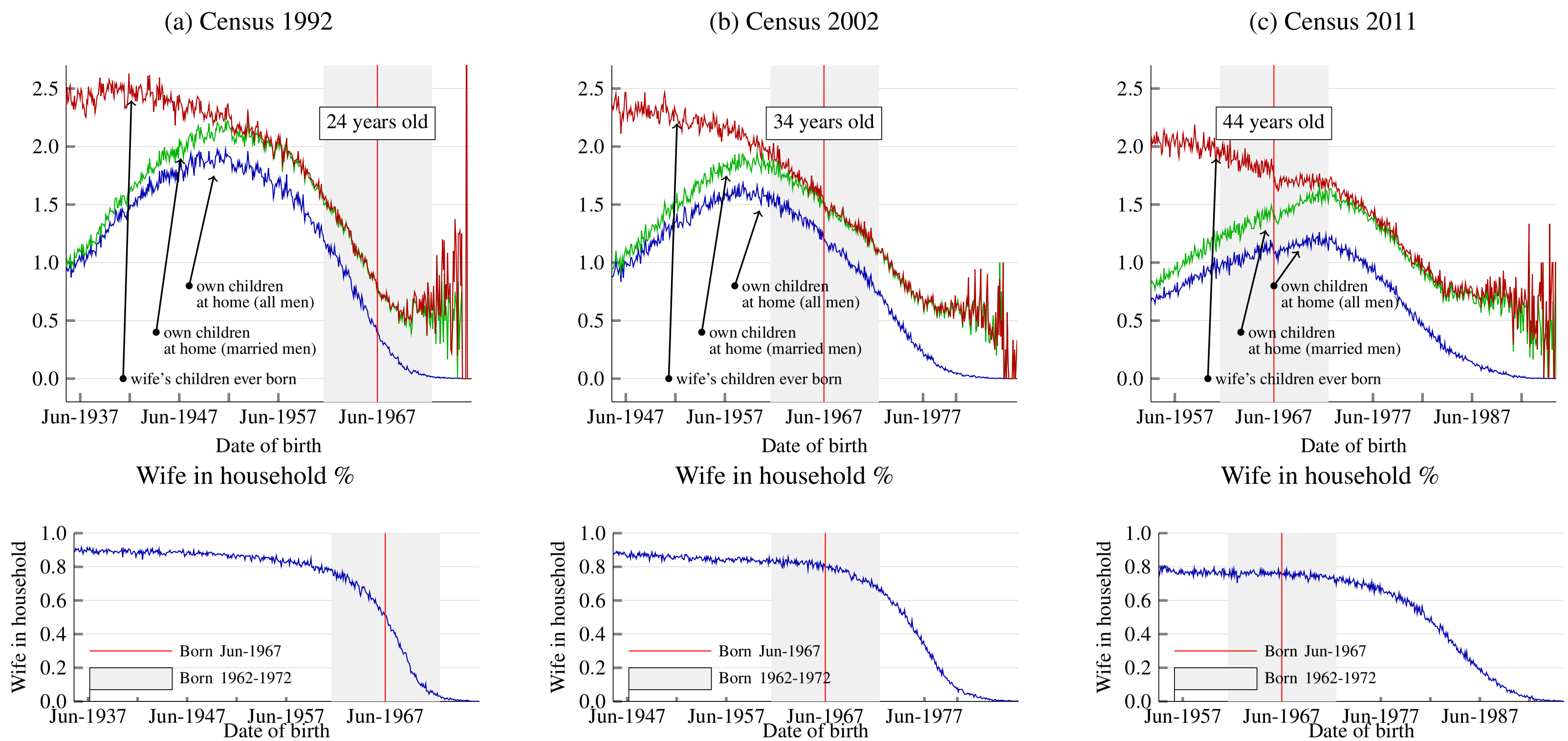


Figure A11: Men's spouse born before June 1967 - Census 2011
(by men's month of birth)

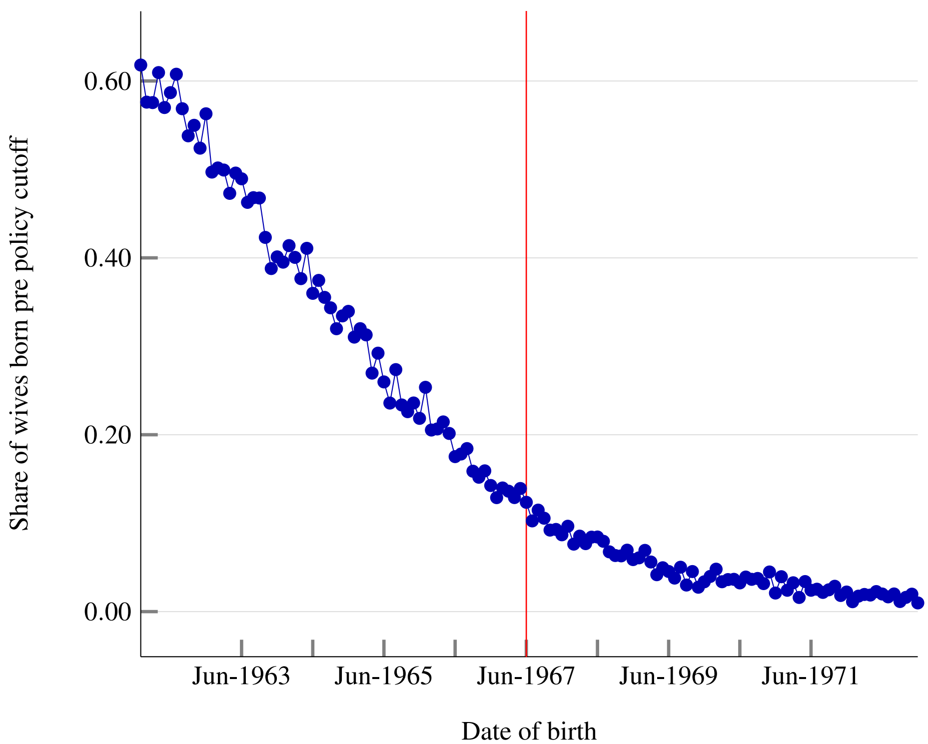


Table A5: Intergenerational correlation of fertility:
(assuming full compliance with the anti-abortion policy)

DEPENDENT VARIABLE: CHILDREN EVER BORN
SAMPLE: WOMEN BORN BETWEEN JAN-1962 AND DEC-1972

	Census 1992 (24 years old)		Census 2002 (34 years old)		Census 2011 (44 years old)	
	OLS	RDD	OLS	RDD	OLS	RDD
	(1)	(2)	(3)	(4)	(5)	(6)
inv. Mills * Born pre Jun-1967 (a)	0.164***	0.089***	0.175***	0.161***	0.175***	0.152***
	(0.019)	(0.023)	(0.016)	(0.025)	(0.015)	(0.019)
Interg. corr.	0.21	0.09	0.15	0.14	0.14	0.12
inv. Mills * Born pre Jun-1967 (b)	0.063***	0.037*	0.099***	0.098***	0.105***	0.084***
	(0.014)	(0.018)	(0.013)	(0.019)	(0.013)	(0.019)
Interg. corr.	0.08	0.04	0.09	0.09	0.09	0.07
inv. Mills * Born pre Jun-1967 (b plus)	0.044**	0.028	0.089***	0.087***	0.098***	0.082***
	(0.014)	(0.019)	(0.013)	(0.020)	(0.014)	(0.022)
Interg. corr.	0.06	0.03	0.08	0.08	0.08	0.07
inv. Mills * Born pre Jun-1967 (c)	0.064***	0.037	0.098***	0.098***	0.110***	0.092***
	(0.014)	(0.019)	(0.014)	(0.019)	(0.014)	(0.021)
Interg. corr.	0.08	0.04	0.09	0.09	0.09	0.08
inv. Mills * Born pre Jun-1967 (c plus)	0.048***	0.029	0.089***	0.088***	0.104***	0.089***
	(0.014)	(0.019)	(0.013)	(0.020)	(0.014)	(0.023)
Interg. corr.	0.06	0.03	0.08	0.08	0.09	0.08
Obs.	159,894	159,894	153,316	153,316	138,953	138,953

Clustered standard errors in parentheses
* $p<0.05$, ** $p<0.01$, *** $p<0.001$

Note: Women born in the second quarter of 1967 are excluded from the sample. Triangular kernel with 36 months band width for RDD.

Table A6: Intergenerational correlation of fertility:
(assuming five abortions per living birth)

DEPENDENT VARIABLE: CHILDREN EVER BORN
SAMPLE: WOMEN BORN BETWEEN JAN-1962 AND DEC-1972

	Census 1992 (24 years old)		Census 2002 (34 years old)		Census 2011 (44 years old)	
	OLS	RDD	OLS	RDD	OLS	RDD
	(1)	(2)	(3)	(4)	(5)	(6)
inv. Mills * Born pre Jun-1967 (a)	0.292***	0.159***	0.312***	0.287***	0.311***	0.270***
	(0.033)	(0.041)	(0.029)	(0.045)	(0.027)	(0.034)
Interg. corr.	0.37	0.17	0.28	0.25	0.25	0.22
inv. Mills * Born pre Jun-1967 (b)	0.110***	0.065*	0.174***	0.173***	0.185***	0.149***
	(0.025)	(0.032)	(0.024)	(0.034)	(0.024)	(0.033)
Interg. corr.	0.14	0.07	0.16	0.15	0.15	0.13
inv. Mills * Born pre Jun-1967 (b plus)	0.077**	0.049	0.158***	0.153***	0.172***	0.145***
	(0.025)	(0.034)	(0.023)	(0.036)	(0.024)	(0.038)
Interg. corr.	0.10	0.05	0.14	0.14	0.14	0.12
inv. Mills * Born pre Jun-1967 (c)	0.112***	0.064	0.172***	0.174***	0.195***	0.162***
	(0.025)	(0.034)	(0.024)	(0.034)	(0.025)	(0.036)
Interg. corr.	0.15	0.07	0.16	0.15	0.16	0.14
inv. Mills * Born pre Jun-1967 (c plus)	0.084**	0.051	0.158***	0.155***	0.183***	0.157***
	(0.025)	(0.034)	(0.024)	(0.035)	(0.025)	(0.040)
Interg. corr.	0.11	0.06	0.14	0.14	0.15	0.14
Obs.	159,894	159,894	153,316	153,316	138,953	138,953

Clustered standard errors in parentheses
* $p<0.05$, ** $p<0.01$, *** $p<0.001$
Note: Women born in the second quarter of 1967 are excluded from the sample. Triangular kernel with 36 months band width for RDD.