The Impact of a Restrictive Abortion Policy on Newborns' Health: The Case of Hungary

Gábor Hajdu[a,*] – Tamás Hajdu[b]

[a] Institute for Sociology, Centre for Social Sciences, Hungarian Academy of Sciences, Hungary
[b] Institute of Economics, Centre for Economic and Regional Studies, Hungarian Academy of Sciences, Hungary
[*] corresponding author; e-mail: hajdu.gabor@tk.mta.hu

March 20, 2019

First draft; not for citation or distribution

Abstract

This paper examines the consequences of the restrictive Hungarian abortion policy that was introduced in 1974. We analyze the impacts on the newborns' health using individual-level live birth and infant mortality registry data. Using difference-in-difference estimations, we show that the law change had a negative impact on the health of the newborns. We found a statistically significant decrease in birth weight and an increase in the probability of being born with low birth weight and in the probability of infant mortality. Performing robustness and placebo tests, we provide evidence that the unwantedness of the children might be a major mechanism in these results.

Keywords: abortion, policy change, health at birth, Hungary

Acknowledgment

This work was supported by the Hungarian National Research, Development and Innovation Office – NKFIH (grant no. FK 124384). The work uses individual-level registers of live births and infant mortalities of the Hungarian Central Statistical Office (HCSO). The source of funding and the HSCO had no role in study design; in analysis and interpretation of data; in the writing of the article.

1. Introduction

The direct and indirect impacts of change in abortion law are an important topic in social science. However, we know much more about these effects for the U.S, than for other countries. Many studies analyzed the impact of the abortion legalization in the United States on fertility (Ananat et al. 2007; Guldi 2008; Levine et al. 1999), on the socio-economic outcomes of children (Ananat et al. 2009; Charles and Stephens 2006; Donohue et al. 2009; Donohue and Levitt 2001, 2004; Gruber et al. 1999; Lin and Pantano 2015), or on health of the newborns (Grossman and Jacobowitz 1981; Gruber et al. 1999; Joyce 1987). There is much less evidence from other countries. Previous studies showed that the elimination of severe abortion restrictions reduces the number of birth in Eastern European countries (Levine and Staiger 2004), in Romania (Pop-Eleches 2010), and in Uruguay (Antón et al. 2018). Pop-Eleches (2006) analyzed the 1966 Romanian abortion ban and found that children born after the ban had worse educational and labor market outcomes when compositional changes were controlled for. These results can be interpreted as the consequences of the unplanned, mistimed, or unwanted pregnancies. Mølland (2016) showed that the liberalization of abortion access in Norway in the 1960s decreased the probability of teen fertility, led to higher educational attainment for women, and it also had similar impacts for their first-born offspring.

Regarding the health of the newborns, Mitrut and Wolff (2011) analyzed the impact of abortion legalization in Romania in 1989. They found that it had limited impact on the children's health: only the probability of being born with low birth weight (under 3000 g) decreased after the legalization. Antón et al (2018) found that the abortion legalization in Uruguay had a positive impact on the APGAR score of the newborns but had no impact on the birth weight and on the probability of premature birth. Similarly, Valente (2014) found no evidence that improved access to a legal abortion center in Nepal decreased the likelihood of neonatal mortality or increased the average level of investments in neonatal health.

In this paper, we examine the consequences of the restrictive Hungarian abortion policy that was introduced in 1974. Due to the law change, the number of induced abortions decreased by almost 70 000 from a baseline of 170 000, and the number of live births increased by 30 000 from a baseline of 156 000 between 1973 and 1974. We analyze the impacts on the health at birth of the affected children using individual-level live birth and infant mortality registry data. We estimate the effects by comparing children born just before and after the law change.

We show that the law change had a negative impact on the health of the newborns controlling for the socio-economic characteristics of the parents. These results indicate the effect of the unwantedness of these children. Using difference-in-difference estimates, we found a statistically significant decrease in birth weight and an increase in the probability of being born with low birth weight and in the probability of infant mortality.

In a series of robustness tests, we show that the results remain the same even if we use additional control variables, different estimation method, and alternative treatment and control periods. We also show that when the outcome variables are changed to the health indicators of the younger sibling of the newborns, there are no differences between the treatment and the control group, which suggests that the results cannot be explained by the compositional change of the mothers. In a placebo reform test, we assume that the law change was not introduced in 1974, but on other random dates in a 13-year period, then we re-estimate the model using these dates as the start of the restriction. In this placebo test, we could not replicate the baseline results.

This paper contributes to the literature in several ways. There are only a very small number of papers that analyze the impact of abortion restrictions (rather than the impact of legalization) on the health of the newborns. Since significant changes in abortion law are rarely introduced, every evidence about the impacts of these law changes gives important information. Second, unlike the previous papers, we use complete registry data. It makes us able to explore the consequences of the law change for smaller subgroups. Lastly, we do not know any other paper that studies the impact of the Hungarian abortion policy. Since in Hungary in the 1970s, abortion had become a standard method of family planning and contraception (Sándor 1999) and in the late 1960s the number of reported, legal abortions per thousand women of childbearing age was one of the highest in Europe (Gal 1994), the case of Hungary can reveal interesting evidence about the effect of restricting access to abortion.

The paper is structured as follows. First, we present the law change and discuss the mechanisms through which a restrictive abortion policy might affect newborns (Section 2). In Section 3, we present the data and the empirical strategy. In Section 4, we show the results and the robustness tests. Section 5 concludes.

2. Background

2.1. The abortion restriction in Hungary

From January 1, 1974, new abortion rules were introduced in Hungary.[1] They were justified as they were intended to protect the health of the women, but the implicit goal was to reduce the number of abortions and to increase the fertility (Bognár and Czeizel 1976; Gal 1994). Although the new rules were not as extreme as abortion regulation in the 1960s in Romania (Haney 2002), or in the early 1950s in Hungary (Szalai 1988), the new regime allowed the access to legal abortion to targeted groups: only for women who were unmarried, or had at least three children, or were at least 35 years old, or had serious housing problems or lived in poverty, or if the birth would cause serious health hazards for the woman (Gal 1994; Haney 2002; Sándor 1999). In addition, women seeking abortion for non-medical reason were charged a substantial fee (Sándor 1999; Szalai 1988).[2] Medical lay committees were set up to decide whether abortion was permissible. The application procedure was humiliating; even for those who had a good chance of a positive decision (Gal 1994; Haney 2002; Szalai 1988). In a media campaign, abortion (and birth control) was attacked as it is rooted in the "»unhealthy« spirit of individualism" that is "unacceptable in a socialist society" (Gal 1994, p. 264).

[1] Governmental Decision No 1040/1973 (X.8.) MT, Ministry Order No 4/1973 (XII.1.) EüM and Ministry Order No 5/1973 (XII.5.) EüM.
[2] 600-1000 HUF that was 21-35% of the average gross monthly earnings of employees. (Source: http://www.ksh.hu/docs/eng/xstadat/xstadat_long/h_qli001.html)

Figure 1: Number of induced abortions and live births between 1967 and 1980

Source: Hungarian Central Statistical Office
(http://www.ksh.hu/docs/eng/xstadat/xstadat_long/h_wdsd001a.html and
http://www.ksh.hu/docs/eng/xstadat/xstadat_long/h_wdsd001b.html)

The law change had substantial impacts immediately. The number of induced abortions decreased from 169 650 to 102 022 between 1973 and 1974 (Figure 1). At the same time, the number of live births increased by 30 000 births: from 156 224 in 1973 to 186 288 in 1974. These changes mean that the number of induced abortions per 100 live births decreased by 50 percent: from 108.6 to 54.8.

Since the decrease in the number of induced abortions is twice as large as the increase of the number of births, we can assume that some women adapted relatively quickly to the new rules. The adaptation process might include using available legal birth control technologies, remaining abstinent or having an illegal abortion. In the early 1970s, access to contraceptives increased (Makay 2016). On the other hand, the number and proportion of spontaneous abortion in total pregnancies, and the number and proportion of abortions induced for medical reasons in total induced abortion also increased between 1973 and 1974. Moreover, the proportion of spontaneous abortion in week 13-27 also increased substantially (A. Czeizel et al. 1984). Many of the spontaneous abortions were illegally induced (Bognár and Czeizel 1976) and a substantial proportion of the increased number of abortions induced for medical reasons can be explained by illegal help of doctors in terminating unwanted pregnancies (E. Czeizel 1983). These suggest that illegal abortion or terminating an unwanted pregnancy for medical reasons might have been substituted legal abortion. In 1973, other policies were also introduced that aimed to increase the fertility of women. Some of them increased the value of the existing allowances (e.g. childcare allowance, housing support), others increased the availability of contraceptives. Contrary to the impact of abortion restriction, the impact of the increased allowances might be positive on the health of the children (Amarante et al. 2016; Bradley et al. 2011), whereas the better availability of contraceptives might decrease the number of the unwanted or unplanned children. Consequently, these policy changes and behavioral changes make harder to reveal the full effect of the abortion restriction and makes its effects look smaller than in reality.

2.2. Mechanisms

There are three main mechanisms through which a restrictive abortion policy might affect outcomes of children (Mitrut and Wolff 2011; Pop-Eleches 2006): unwantedness effect, composition effect, and crowding effect. Since restrictive abortion laws make abortion less available, women have three options: using other birth control technologies, remaining abstinent or having more births. Therefore, reduced access to abortion increase the number of unplanned, mistimed, or unwanted children (*unwantedness effect)*. This might affect the children in three ways. (i) According to the standard model of child quality-quantity trade-off, the higher number of children leads to a decrease in child quality if parents desire equal levels of quality for all of their children (Becker 1993; Becker and Lewis 1973). (ii) Women are less able to delay the childbearing to an optimal time when it does not conflict with their marriage, labor market and educational plans (Angrist and Evans 2000; Goldin and Katz 2002). This might lead to unfavorable emotional, material, and physical conditions for raising a child, which might have direct and indirect negative impacts on the child. (iii) Restricted access to abortion may not allow the parents to end the pregnancy if the fetus is in bad health condition. It is also possible that it leads to delayed initiation of prenatal care and/or unhealthy prenatal care due to unwantedness of the fetus (Eggleston 2000; Gipson et al. 2008; Grossman and Jacobowitz 1981; Kost et al. 1998; Kost and Lindberg 2015).

The composition of women carrying pregnancies to term might also change after changes in abortion policy *(composition effect)*. The direction of this effect is ambiguous both theoretically and empirically. Evidence from the United States suggests that low-status women are more likely to have an abortion (Ananat et al. 2009; Gruber et al. 1999; Levine et al. 1996). On the other hand, Pop-Eleches (2006) found that in Romania highly educated women were more likely to use abortion in the 1960s, and a similar result is documented in Norway as well (Mølland 2016).

A negative *crowding effect* might emerge due to changes in cohort size. A larger cohort shares the same resources which might affect negatively the outcomes of the children. Such an effect was found in Romania after the abortion ban in 1967 (Pop-Eleches 2006). However, regarding health outcomes, the crowding effect is less relevant since newborns' health depends primarily on the characteristics of the parents (Mitrut and Wolff 2011).

3. Data and empirical strategy

3.1. Data

We use individual-level registers of live births and infant mortalities of the Hungarian Central Statistical Office (HCSO). These datasets contain every live birth and every infant death in the first year of life in Hungary. We accessed the deidentified datasets in the Safe Centre of the HCSO where the protection of individual statistical data is ensured. Infant mortality and live birth datasets are linked. The matching is almost perfect: more than 99% of the infant mortality records are linked to the live birth dataset.

These datasets contain information on the date of the birth and the infant death, the sex of the newborn, and characteristics of both parents (age, education, marital status, employment, occupation code from the standard classification of occupations in Hungary, place of residence, and pregnancy history of the mother).

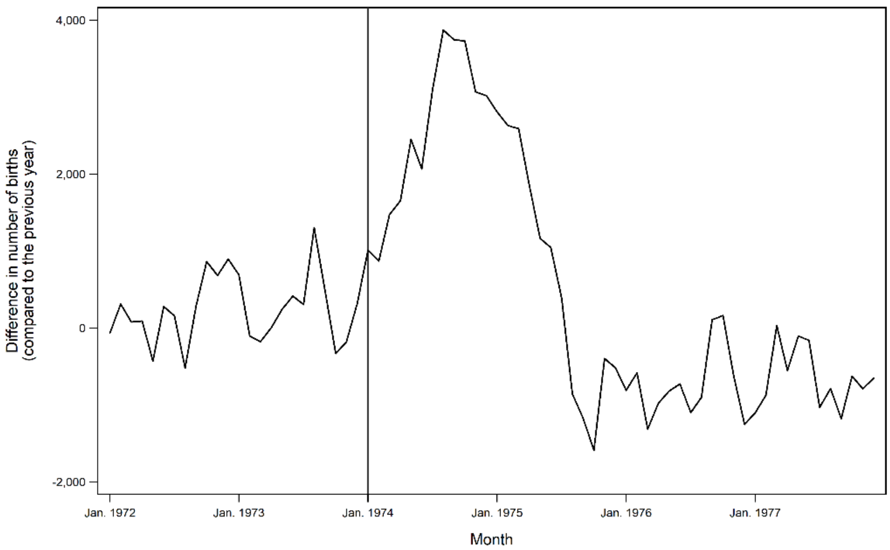
We use four outcome variables. (1) Birth weight (in grams). We choose two main definitions of low birth weight. (2) The standard definition as a birth weight of less than 2500 grams. (3) We also use a low birth weight variable with a higher threshold (<3000 g) that makes the results comparable to the results of Mitrut and Wolff (2011). The other reason behind the decision is that only 9.8% of the newborns are below the 2500 grams limit, but 33.1% is under the 3000 grams limit. (4) Infant mortality is defined as death during the first year of life.

We exclude twin births because of their special characteristics (e.g their birth weight is not comparable to singletons).

3.2. Empirical strategy

Figure 2 shows the number of monthly births compared to the same month of the previous year. The number of births increased after the first months of 1974 and the change reached a peak level in the summer months. A similar six months lag and a change with a similar magnitude in the number of births were found after the legalization of abortion in Romania (Mitrut and Wolff 2011). This six-month lag in the fertility-increase is not unexpected since abortion was permitted before the twelfth gestational week before 1974. Our identification strategy is based on the assumption that a large proportion of these "new" children would not have been born if abortion policy had not changed.

Figure 2: Difference in number of births compared to the same month in the previous year between 1972 and 1977



Note: The vertical line refers to January 1974, the first month when the new law came into effect.
Source: own calculation from Registry of live birth of the Hungarian Central Statistical Office.

Our identification strategy is similar to Pop-Eleches (2006) and Mitrut and Wolff (2011): we estimate the effect of the law change by comparing children born just before and after the law change, i.e. we rely on data of children born within a reasonably short timespan around the law change. However, similarly to the empirical strategy of Adamecz-Völgyi et al. (2018), we utilize that the new abortion rules made abortion permissible to selected groups of women: e.g.

for women who were at least 35 years old, for women who were unmarried or for women who had at least three children (Gal 1994; Haney 2002; Sándor 1999). We use the newborns of these women as control group and create treatment groups that are as similar as possible to these groups. Using these treatment and control groups, we can apply a difference-in-differences framework to estimate the impact of the abortion restriction.[3]

The control and treatment groups are the following. (1) Women who were 35.5-36.5 years old at the conception (control group) vs. women who were 33.5-34.5 years old at the conception (treatment group).[4] (2) Women who were not married (control group) vs. women who were married (treatment group). In this case, to have more homogeneous groups, the sample is restricted to women aged 25-29. (3) As a similar but somewhat different strategy, we build on the fact that previous research shows that unintended and mistimed pregnancies are more likely among teen women (D'Angelo et al. 2004; Finer and Henshaw 2006; Kamarás 2006; Pulley et al. 2002). Younger women also want less likely a baby with their partners than older women (Wilson and Koo 2006). It is reasonable to assume that for young women giving birth to a child after the law restrictions was more likely to be conflicted with their educational or labor market plans, than for older women. In the presence of the option of abortion, they would have used it to delay their pregnancies to an optimal time. Therefore, we created a treatment group consisting of mothers who were 16-18 years old at the time of the birth and a control group consisting of mothers who were 19-20 years old at the time of the birth. We assume that women in the treatment group were more likely to give birth to an unwanted child, hence, the abortion restriction is more likely to have affected them.

In the baseline estimation, we compare children born between July and September 1974 to children born between April and June 1974. Most of the children born between July and September 1974 were conceived between October and December 1973, thus their mothers were pregnant when the law change happened.

We estimate the following equation:

$$Y_i = \beta_0 + \beta_1 Treat_i + \beta_2 After_i + \beta_3 Treat_i \times After_i + \beta_4 X_i + \varepsilon_i \quad (1)$$

where Y_i is an outcome of interest (birth weight, low birth weight or infant mortality) for child i, $Treat_i$ is a dummy that takes the value of 1 if the child is born to a mother in the treatment group and 0 if the child is born to a mother in the control group. $After_i$ is a dummy that takes the value of 1 if the child is born between July and September 1974 and 0 if the child is born between April and June 1974. X_i is a vector of control variables that includes sex of the newborn, characteristics of the mother (age, age squared, dummies for marital status, dummies for education level, dummies for first language, dummies for labor force status, dummies for occupation, dummies for type of employment, dummies for number of pregnancies and live births, the number of years between the current and previous live birth, dummies for county and for the type of settlement), characteristics of the father (age, age squared, dummies for education level, dummies for labor force status, dummies for occupation, dummies for type of

[3] A similar empirical strategy is used by Adamecz-Völgyi et al. (2018) when analyzing the impact of the abortion restriction on birth defects. They compare newborns of 34-year-old and of 36-year-old mothers.

[4] We excluded women around age 35 as there is no information about the exact decision process of the abortion committees and we do not know how abortion requests of women around this age limit were treated.

work), and interaction terms for some of the characteristics of the parents (education, occupation, labor force status).

Although the composition of women carrying pregnancies to term might change after changes in abortion policy (Ananat et al. 2009; Pop-Eleches 2006), with the rich set of control variables we can control for a substantial part of this composition effect. Since regarding health outcomes, the crowding effect is less concern (Mitrut and Wolff 2011), with this empirical strategy we assume to estimate the unwantedness effect.[5] The key coefficient is β_3 that captures the unwantedness effect. We report only this coefficient throughout the paper.

Since there is no information about the mothers' applications for abortion (or their intentions of having an abortion), and since many children have been born regardless of the law change (see Figure 1), we note that β_3 measures an intention-to-treat effect (ITT) and the treatment-on-the-treated effect (TOT) might be higher.

We estimate Equation (1) using an OLS regression. Standard errors are robust to heteroscedasticity. Dummies are included for missing control variables.

4. Results

4.1. Main results

Table 1 shows the results from estimating Equation (1).[6] Panel A presents the results for mothers aged 33.5-34.5 (treatment group) vs. aged 35.5-36.5. All the four coefficients have a negative sign and despite the small sample size, all are significant. The law change decreases the birth weight of newborns of mothers aged 33.5-34.5 by 125 g which is a 4.0% decrease compared to the average birth weight before the law change came into effect. Low birth weight increased by 7.9 percentage points and infant mortality by 4.2 percentage points. Low birth weight using a higher threshold (3000 g) also increased by 9.3 percentage point.

Panel B shows the results when comparing married mothers (treatment group) and non-married mothers (restricting the sample to women aged 25-29). Birth weight of the newborns of married mothers decreased by 94 g (or 3.0%) whereas both the probability of being born with low birth weight and the probability of infant mortality increased (by 6.0 and 4.2 percentage point, respectively).

Panel C presents the results for mothers aged 16-18 (treatment group) vs. mothers aged 19-20. Again, three of the four health indicators are significant at the 5% or at the 10% level. In this case, since abortion was similarly restricted for both the treatment and the control groups we can identify from the differences in probability of having an unwanted child between mothers aged 16-18 and aged 19-20 (Finer and Henshaw 2006; Kamarás 2006). This might be a weaker difference than having different access to abortion thus the size of the coefficients is smaller than in Panel A and Panel B. The law change decreased the birth weight of the newborns of young mothers by 32.6 g, increased probability of infant mortality by 1.1 percentage points and

[5] Nevertheless, we also show a robustness test that suggests that the control variables indeed sufficiently control for the compositional changes of women after the law change.
[6] Table A1, Table A2, and Table A3 show summary statistics of the three samples.

probability of being born with low birth weight (using 3000 g threshold) by 2.7 percentage points.

Since many children have been born regardless of the law change (see Figure 1) these estimates are ITT effects. The impact of the law change on the really affected newborns (who would not have been born without the law change) might be much higher (TOT effect).

Table 1: The effect of the abortion restriction on birth outcomes

		Treat×After	Robust SE	p	N
A.	Mothers aged 33.5-34.5 (vs. aged 35.5-36.5)				
(1)	Birth weight	-124.940	64.065	0.051	2128
(2)	Low birth weight (<2500 g)	0.079	0.034	0.020	2128
(3)	Low birth weight (<3000 g)	0.093	0.045	0.037	2128
(4)	Infant mortality	0.042	0.021	0.049	2128
B.	Married mothers (vs. non-married mothers)				
(5)	Birth weight	-94.463	50.350	0.061	25076
(6)	Low birth weight (<2500 g)	0.060	0.030	0.046	25076
(7)	Low birth weight (<3000 g)	0.032	0.038	0.398	25076
(8)	Infant mortality	0.042	0.016	0.010	25076
C.	Mothers aged 16-18 (vs. aged 19-20)				
(9)	Birth weight	-32.550	15.513	0.036	22366
(10)	Low birth weight (<2500 g)	0.011	0.010	0.255	22366
(11)	Low birth weight (<3000 g)	0.027	0.014	0.057	22366
(12)	Infant mortality	0.011	0.005	0.032	22366

Control variables: Panel A and Panel C: Sex of the newborn, characteristics of the mother (age, squared age, education, labor force status, occupation, type of employment, birth month, marital status, first language, number of pregnancies, number of live births, number years since the previous live birth, county, type of settlement), characteristics of the father (age, squared age, education, labor force status, occupation, type of employment), interactions for characteristics of the parents (education, occupation, labor force status). Panel B: as Panel A excluding marital status.

4.2. Robustness of the results

In Table 2, we change the treatment period. First, we excluded July since this month might be regarded as an intermediate period. Children born in the first days of July after a longer gestation than the average might be conceived before October 1973, thus, their mothers were able to use abortion before the twelfth week of pregnancy. In these estimations, we compare child born between August and September to children born between May and June. The coefficients have similar or only slightly smaller size compared to the main results, but standard errors are lager hence they lose some significance.

We also use a longer treatment period and compare July-October to March-June. These coefficients are usually smaller and lose some significance that suggests some adaptation. Nevertheless, these results suggest that using alternative treatment and control period does not change substantially the inference on the impact of the law change.

Table 2: The effect of the abortion restriction on birth outcomes for alternative treatment periods

	Aug-Sept. vs. May-June				July-Oct. vs. March-June			
	Treat × After	Robust SE	p	N	Treat × After	Robust SE	p	N
A. Mothers aged 33.5-34.5 (vs. aged 35.5-36.5)								
(1) Birth weight	-91.450	81.986	0.265	1433	-65.813	55.509	0.236	2792
(2) Low birth weight (<2500 g)	0.059	0.041	0.153	1433	0.048	0.029	0.093	2792
(3) Low birth weight (<3000 g)	0.084	0.057	0.138	1433	0.051	0.039	0.184	2792
(4) Infant mortality	0.028	0.027	0.288	1433	0.030	0.018	0.098	2792
B. Married mothers (vs. non-married mothers)								
(1) Birth weight	-91.287	61.709	0.139	16989	-51.679	43.884	0.239	32993
(2) Low birth weight (<2500 g)	0.064	0.037	0.082	16989	0.046	0.027	0.081	32993
(3) Low birth weight (<3000 g)	0.037	0.046	0.422	16989	0.021	0.033	0.531	32993
(4) Infant mortality	0.050	0.020	0.011	16989	0.026	0.015	0.080	32993
C. Mothers aged 16-18 (vs. aged 19-20)								
(1) Birth weight	-31.323	18.908	0.098	15030	-35.261	13.533	0.009	29537
(2) Low birth weight (<2500 g)	0.010	0.012	0.413	15030	0.011	0.009	0.209	29537
(3) Low birth weight (<3000 g)	0.019	0.017	0.278	15030	0.022	0.012	0.078	29537
(4) Infant mortality	0.011	0.006	0.080	15030	0.009	0.005	0.052	29537

Control variables: see Table 1.

We perform two placebo tests using placebo treatment groups, and placebo treatment periods, respectively. First, we change treatment and controls groups to mothers who were identically affected by the restricted abortion rules. If the main results indeed show the effect of the law change and not the effect of other unmeasured variables, we should observe no differences between placebo treatment and control groups. We compare mothers aged 30.5-31.5 and mothers aged 31.5-32.5, and mothers aged 20-21 and mothers aged 22-23.[7] The results are reported in Table A4 and confirm that there is no difference between the placebo treatment and the placebo control groups. This suggests that the main models capture the effect of the law change.

Next, to check that the estimated impacts are not just the result of coincidence or a general trend in these years, we perform a placebo reform test. We use data from other years between 1972 and 1985 and we assume that the law change was introduced one or two years before or one or more years later. We estimate the effect of placebo law changes in these years using identical empirical procedure to the main model. Table 3 reports the summarized results. For every year, we count the number of significant coefficients with the expected sign. Since we have three treatment groups and four outcome variables, the maximum number of the significant coefficients is twelve. For 1974, seven coefficients are significant at the 5% level and ten

[7] For married and unmarried women, we cannot create any meaningful placebo groups.

coefficients at the 10% level. In other years, the coefficients are hardly significant even at the 10 percent level. These results confirm that we cannot observe similar systematical changes in the health indicators of the newborns in other years that are similar to the observed negative changes in 1974. This suggests that the estimations show the causal effect of the law change and it is unlikely that they are the results of pure chance or the results of a long-run trend in this period.

Table 3: The results of the treatment effects in other years

Year	Number of significant estimations with the expected sign (at the 5 percent level)	Number of significant estimations with the expected sign (at the 10 percent level)
1972	0	0
1973	3	3
1974 (main results)	7	10
1975	3	3
1976	0	0
1977	1	2
1978	0	3
1979	0	0
1980[a]	1	2
1981[a]	2	3
1982[a]	0	0
1983	1	2
1984	0	0
1985	0	0

[a] Infant mortality data is missing in 1981, therefore no estimations for 1980, 1981 and 1982.

To rule out that compositional change of the young mothers explains the results, we use another placebo test where we compare the younger siblings of the children in the treatment and the control groups. Since all younger siblings are conceived after new abortions laws, therefore, they are affected similarly by them. Theoretically, it is possible that women who gave birth after the law change have different unobserved, time-invariant characteristics compared to women who gave birth before the law change. In this case, a compositional change would explain the main results – at least partially. However, if the composition effect explains the results, we should also observe similar differences between the younger siblings as in the baseline sample. Non-zero coefficients would suggest that the unobserved and stable "quality" of the mothers explains the results instead of the unwantedness effect. On the contrary, zero or close to zero associations would suggest that there is no remaining composition effect and the results are driven by the unwantedness effect.[8]

[8] For newborns of mothers around age 35, there is only a small number of younger siblings, therefore we exclude this group from this exercise.

Table 4 reports these estimations. To have meaningful benchmark coefficients, we re-estimate the model restricting the sample to children with a younger sibling. The re-estimated coefficients are similar to the coefficients in Table 1. The coefficients for younger siblings are insignificant and mostly close to zero or they point in the theoretically "wrong" direction. It means that younger siblings of the treatment group are similar to younger siblings of the control group. This suggests that the law change explains the results of the baseline estimation rather than any time-invariant compositional differences between the treated and control mothers.

Table 4: The effect of the abortion restriction on birth outcomes for younger siblings born in the next 10 years

		Younger siblings of the original sample			Original sample excluding observations without younger siblings			
		Treat × After	Robust SE	p	Treat × After	Robust SE	p	N
A.	Married mothers (vs. non-married mothers)							
(1)	Birth weight	-15.265	79.339	0.847	-141.711	79.677	0.075	7328
(2)	Low birth weight (<2500 g)	0.063	0.046	0.172	0.049	0.048	0.313	7328
(3)	Low birth weight (<3000 g)	0.038	0.061	0.534	0.105	0.062	0.088	7328
(4)	Infant mortality	-0.005	0.029	0.875	0.014	0.022	0.535	6347
B.	Mothers aged 16-18 (vs. aged 19-20)							
(5)	Birth weight	-0.114	18.628	0.995	-37.240	17.553	0.034	15473
(6)	Low birth weight (<2500 g)	-0.013	0.010	0.184	0.017	0.011	0.129	15473
(7)	Low birth weight (<3000 g)	-0.003	0.016	0.872	0.031	0.017	0.060	15473
(8)	Infant mortality	-0.004	0.006	0.495	0.009	0.006	0.115	13460

Younger siblings of the original sample: the outcome variables were changed to the health at birth indicators of the younger sibling of the newborns.
Original sample with younger siblings: As Table 1, but children without a younger sibling in the next 10 years are excluded.
Infant mortality data is missing in 1981, therefore, no estimations for siblings born in 1980, 1981 and 1982 in Row 4, and Row 8.
Control variables: see Table 1.

Lastly, since three of the four outcome variables are binary, we estimate Equation (1) using probit regressions. Average marginal effects are reported in Table A5 in the Appendix. These estimations yield identical results to Table 1.

5. Conclusion

In this paper, we estimated the impact of the restrictive Hungarian abortion policy that was introduced in 1974 on the health of the newborns. Using a difference-in-differences approach, we found that the law change had a negative impact on the health outcomes of the newborns. The estimated effects for subgroups of mothers who were more likely to give birth to an unplanned or unwanted child are between 32.6 g and 125.0 g regarding birth weight, between 1.1 percentage points and 7.9 percentage points regarding low birth weight, between 2.7

percentage points and 9.3 percentage points regarding low birth weight using higher threshold, and between 1.1 percentage points and 4.2 percentage points regarding infant mortality.

We also showed that the unwantedness of the children might be the most important mechanism in these results. Performing robustness tests (e.g. a placebo reforms in other years, a placebo test using siblings' data), we provided evidence that supports this interpretation.

These results are in line with the findings of Gruber et al (1999) who report a decrease both in infant mortality rates and in the percentage of children born with low birth weight after abortion legalization in the US in the 1970s. Similar results were obtained by Joyce (1987) regarding the impact of the abortion rate on neonatal mortality and on low birth weight. On the other hand, our estimations somewhat differ from the results of Mitrut and Wolf (2011). They found only limited health effects of the Romanian abortion legalization in 1990. As they note, the insignificant effects they found might be explained by the positive emotional changes after the collapse of communism (in 1990) that affected the control group as well. The difference in the sample sizes might also a relevant explanation since their effect sizes are comparable to the effects found in this paper.

According to the results of this paper, an adaptation to the new rules started soon after the law change. The estimated impacts are smaller when longer treatment periods were used. This suggests that the restricted access to abortion had relatively short-term impacts on newborns' health due to different strategies of adaptation. As we noted above, the adaptation process included having an illegal abortion or terminating an unwanted pregnancy formally for medical reasons (Bognár and Czeizel 1976; A. Czeizel et al. 1984) or using contraceptives (Makay 2016).

This paper has some limitations. As we noted above, in these years, other policies were also introduced that aimed to increase the fertility of women. However, the potential effect of these policy changes is the opposite of the effects of the abortion restriction and might be less relevant in a difference-in-differences framework.

We estimated the impacts of the law change by comparing women for whom abortion was permissible with women for whom access to abortion was restricted. That is, we estimated ITT effects for special subgroups: for older mother, for married mothers and for teen mothers. The generalizability of these effects is unknown. Also, since a sizeable proportion of the children have been born regardless of the law change (even for these subgroups) the estimated effects are intention-to-treat effects and the treatment-on-the-treated effects might be considerably higher.

Significant changes in the abortion laws are very rare, and the effect of restrictions in abortion laws can be even more rarely analyzed, hence these results give important information on the consequences of access to abortion and, in a wider perspective, to family planning. Health at birth has a long-lasting impact on children's health and other life outcomes (Almond et al. 2018; Almond and Currie 2011; Black et al. 2007; Currie 2009; Oreopoulos et al. 2008), therefore the restrictions in the access to abortion might have long-term consequences for the affected newborns. Since abortion policy is still an emerging issue in public debates in many countries, these results could provide important information for evidence-based policy.

References

Adamecz-Völgyi, A., Bárdits, A., Kertesi, G., & Szabó-Morvai, Á. (2018). Sokkok hatása a születési kimenetekre. In K. Fazekas & Á. Szabó-Morvai (Eds.), *Munkaerőpiaci Tükör 2017*. Budapest: MTA Közgazdaság- és Regionális Tudományi Kutatóközpont Közgazdaság-tudományi Intézet.

Almond, D., & Currie, J. (2011). Killing Me Softly: The Fetal Origins Hypothesis. *Journal of Economic Perspectives*, *25*(3), 153–172. doi:10.1257/jep.25.3.153

Almond, D., Currie, J., & Duque, V. (2018). Childhood Circumstances and Adult Outcomes: Act II. *Journal of Economic Literature*, *56*(4), 1360–1446. doi:10.1257/jel.20171164

Amarante, V., Manacorda, M., Miguel, E., & Vigorito, A. (2016). Do Cash Transfers Improve Birth Outcomes? Evidence from Matched Vital Statistics, Program, and Social Security Data. *American Economic Journal: Economic Policy*, *8*(2), 1–43. doi:10.1257/pol.20140344

Ananat, E. O., Gruber, J., & Levine, P. B. (2007). Abortion Legalization and Life-Cycle Fertility. *Journal of Human Resources*, *XLII*(2), 375–397. doi:10.3368/jhr.XLII.2.375

Ananat, E. O., Gruber, J., Levine, P. B., & Staiger, D. (2009). Abortion and Selection. *Review of Economics and Statistics*, *91*(1), 124–136. doi:10.1162/rest.91.1.124

Angrist, J. D., & Evans, W. N. (2000). Schooling and labor market consequences of the 1970 state abortion reforms. In *Research in Labor Economics* (Vols 1-0, Vol. 18, pp. 75–113). Emerald Group Publishing Limited. doi:10.1016/S0147-9121(99)18020-8

Antón, J.-I., Ferre, Z., & Triunfo, P. (2018). The impact of the legalisation of abortion on birth outcomes in Uruguay. *Health Economics*, *27*(7), 1103–1119. doi:10.1002/hec.3659

Becker, G. S. (1993). *A Treatise on the Family: Enlarged Edition* (Enlarged edition.). Cambridge, Mass.: Harvard University Press.

Becker, G. S., & Lewis, H. G. (1973). On the Interaction between the Quantity and Quality of Children. *Journal of Political Economy*, *81*(2, Part 2), S279–S288. doi:10.1086/260166

Black, S. E., Devereux, P. J., & Salvanes, K. G. (2007). From the Cradle to the Labor Market? The Effect of Birth Weight on Adult Outcomes. *The Quarterly Journal of Economics*, *122*(1), 409–439. doi:10.1162/qjec.122.1.409

Bognár, Z., & Czeizel, A. (1976). Mortality and morbidity associated with legal abortions in Hungary, 1960-1973. *American Journal of Public Health*, *66*(6), 568–575. doi:10.2105/AJPH.66.6.568

Bradley, E. H., Elkins, B. R., Herrin, J., & Elbel, B. (2011). Health and social services expenditures: associations with health outcomes. *BMJ Qual Saf*, *20*(10), 826–831. doi:10.1136/bmjqs.2010.048363

Buckles, K. S., & Hungerman, D. M. (2012). Season of Birth and Later Outcomes: Old Questions, New Answers. *Review of Economics and Statistics*, *95*(3), 711–724. doi:10.1162/REST_a_00314

Charles, K. K., & Stephens, J., Melvin. (2006). Abortion Legalization and Adolescent Substance Use. *The Journal of Law & Economics*, *49*(2), 481–505. doi:10.1086/508249

Currie, J. (2009). Healthy, Wealthy, and Wise: Socioeconomic Status, Poor Health in Childhood, and Human Capital Development. *Journal of Economic Literature*, *47*(1), 87–122. doi:10.1257/jel.47.1.87

Currie, J., & Schwandt, H. (2013). Within-mother analysis of seasonal patterns in health at birth. *Proceedings of the National Academy of Sciences*, *110*(30), 12265–12270. doi:10.1073/pnas.1307582110

Czeizel, A., Bognár, Z., & Rockenbauer, M. (1984). Some epidemiological data on spontaneous abortion in Hungary, 1971-80. *Journal of Epidemiology and Community Health*, *38*(2), 143–148. doi:10.1136/jech.38.2.143

Czeizel, E. (1983). A terhességmegszakítások orvosi javallat alapján történő engedélyezéseinek az elemzése. *Orvosi Hetilap*, *124*(22), 1297–1302.

D'Angelo, D. V., Gilbert, B. C., Rochat, R. W., Santelli, J. S., & Herold, J. M. (2004). Differences Between Mistimed and Unwanted Pregnancies Among Women Who Have Live Births. *Perspectives on Sexual and Reproductive Health*, *36*(5), 192–197. doi:10.1363/3619204

Donohue, J. J., Grogger, J., & Levitt, S. D. (2009). The Impact of Legalized Abortion on Teen Childbearing. *American Law and Economics Review*, *11*(1), 24–46. doi:10.1093/aler/ahp006

Donohue, J. J., & Levitt, S. D. (2001). The Impact of Legalized Abortion on Crime. *The Quarterly Journal of Economics*, *116*(2), 379–420. doi:10.1162/00335530151144050

Donohue, J. J., & Levitt, S. D. (2004). Further Evidence that Legalized Abortion Lowered Crime A Reply to Joyce. *Journal of Human Resources*, *XXXIX*(1), 29–49. doi:10.3368/jhr.XXXIX.1.29

Eggleston, E. (2000). Unintended pregnancy and women's use of prenatal care in Ecuador. *Social Science & Medicine*, *51*(7), 1011–1018. doi:10.1016/S0277-9536(00)00010-1

Finer, L. B., & Henshaw, S. K. (2006). Disparities in Rates of Unintended Pregnancy In the United States, 1994 and 2001. *Perspectives on Sexual and Reproductive Health*, *38*(2), 90–96. doi:10.1363/3809006

Gal, S. (1994). Gender in the Post-socialist Transition: the Abortion Debate in Hungary. *East European Politics and Societies*, *8*(2), 256–286. doi:10.1177/0888325494008002003

Gipson, J. D., Koenig, M. A., & Hindin, M. J. (2008). The Effects of Unintended Pregnancy on Infant, Child, and Parental Health: A Review of the Literature. *Studies in Family Planning*, *39*(1), 18–38. doi:10.1111/j.1728-4465.2008.00148.x

Goldin, C., & Katz, L. F. (2002). The Power of the Pill: Oral Contraceptives and Women's Career and Marriage Decisions. *Journal of Political Economy*, *110*(4), 730–770. doi:10.1086/340778

Grossman, M., & Jacobowitz, S. (1981). Variations in infant mortality rates among counties of the United States: The roles of public policies and programs. *Demography*, *18*(4), 695–713. doi:10.2307/2060956

Gruber, J., Levine, P. B., & Staiger, D. (1999). Abortion Legalization and Child Living Circumstances: Who is the "Marginal Child"? *The Quarterly Journal of Economics*, *114*(1), 263–291. doi:10.1162/003355399556007

Guldi, M. (2008). Fertility effects of abortion and birth control pill access for minors. *Demography*, *45*(4), 817–827. doi:10.1353/dem.0.0026

Haney, L. (2002). *Inventing the Needy: Gender and the Politics of Welfare in Hungary*. Berkeley: University of California Press.

Joyce, T. J. (1987). The impact of induced abortion on black and white birth outcomes in the United States. *Demography*, *24*(2), 229–244. doi:10.2307/2061631

Kamarás, F. (2006). Kívánt és nem kívánt terhességek, gyermekek. *Demográfia*, *49*(2–3), 150–172.

Kost, K., Landry, D. J., & Darroch, J. E. (1998). The Effects of Pregnancy Planning Status on Birth Outcomes and Infant Care. *Family Planning Perspectives*, *30*(5), 223–230. doi:10.2307/2991608

Kost, K., & Lindberg, L. (2015). Pregnancy Intentions, Maternal Behaviors, and Infant Health: Investigating Relationships with New Measures and Propensity Score Analysis. *Demography*, *52*(1), 83–111. doi:10.1007/s13524-014-0359-9

Levine, P. B., & Staiger, D. (2004). Abortion Policy and Fertility Outcomes: The Eastern European Experience. *The Journal of Law and Economics*, *47*(1), 223–243. doi:10.1086/380475

Levine, P. B., Staiger, D., Kane, T. J., & Zimmerman, D. J. (1999). Roe v Wade and American fertility. *American Journal of Public Health*, *89*(2), 199–203. doi:10.2105/AJPH.89.2.199

Levine, P. B., Trainor, A. B., & Zimmerman, D. J. (1996). The effect of Medicaid abortion funding restrictions on abortions, pregnancies and births. *Journal of Health Economics*, *15*(5), 555–578. doi:10.1016/S0167-6296(96)00495-X

Lin, W., & Pantano, J. (2015). The unintended: negative outcomes over the life cycle. *Journal of Population Economics*, *28*(2), 479–508. doi:10.1007/s00148-014-0530-z

Makay, Z. (2016). Contraceptive use in Hungary: Past trends and actual behaviour. *Demográfia English Edition*, *58*(5), 65–90. doi:10.21543/DEE.2015.3

Mitrut, A., & Wolff, F.-C. (2011). The impact of legalized abortion on child health outcomes and abandonment. Evidence from Romania. *Journal of Health Economics*, *30*(6), 1219–1231. doi:10.1016/j.jhealeco.2011.08.004

Mølland, E. (2016). Benefits from delay? The effect of abortion availability on young women and their children. *Labour Economics*, *43*, 6–28. doi:10.1016/j.labeco.2016.06.011

Oreopoulos, P., Stabile, M., Walld, R., & Roos, L. L. (2008). Short-, Medium-, and Long-Term Consequences of Poor Infant Health An Analysis Using Siblings and Twins. *Journal of Human Resources*, *43*(1), 88–138. doi:10.3368/jhr.43.1.88

Pop-Eleches, C. (2006). The Impact of an Abortion Ban on Socioeconomic Outcomes of Children: Evidence from Romania. *Journal of Political Economy*, *114*(4), 744–773. doi:10.1086/506336

Pop-Eleches, C. (2010). The Supply of Birth Control Methods, Education, and Fertility: Evidence from Romania. *The Journal of Human Resources*, *45*(4), 971–997.

Pulley, L., Klerman, L. V., Tang, H., & Baker, B. A. (2002). The Extent of Pregnancy Mistiming and Its Association with Maternal Characteristics and Behaviors and Pregnancy Outcomes. *Perspectives on Sexual and Reproductive Health*, *34*(4), 206–211. doi:10.2307/3097731

Sándor, J. (1999). From Ministry Orders towards Constitutional Debate: Lessons Drawn from the Past 50 Years of Abortion Laws in Hungary. *Medicine and Law*, *18*, 389–408.

Szalai, J. (1988). Abortion in Hungary. *Feminist Review*, (29), 98–100. doi:10.2307/1395151

Valente, C. (2014). Access to abortion, investments in neonatal health, and sex-selection: Evidence from Nepal. *Journal of Development Economics*, *107*, 225–243. doi:10.1016/j.jdeveco.2013.12.002

Wilson, E. K., & Koo, H. P. (2006). The Relationship Context: Its Effects on Low-Income Women's Desire for a Baby. *Journal of Marriage and Family*, *68*(5), 1326–1340. doi:10.1111/j.1741-3737.2006.00331.x

Appendix

Table A1: Descriptive statistics for older mothers (control group: aged 35.5-36.5, treatment group: aged 33.5-34.5)

	Period	Control	Treatment	Diff.	p	N_C	N_T
Birth weight	Before	3024.94	3113.93	88.99	0.051	399	600
	After	3133.98	3113.45	-20.53	0.626	396	733
Low birth weight (<2500 g)	Before	0.188	0.133	-0.055	0.019	399	600
	After	0.129	0.151	0.023	0.301	396	733
Low birth weight (<3000 g)	Before	0.414	0.328	-0.085	0.006	399	600
	After	0.348	0.351	0.002	0.943	396	733
Infant mortality	Before	0.083	0.050	-0.033	0.037	399	600
	After	0.038	0.044	0.006	0.643	396	733
Sex: female	Before	0.474	0.462	-0.012	0.710	399	600
	After	0.495	0.495	0.000	0.993	396	733
Mother's age	Before	36.261	34.218	-2.042	0.000	399	600
	After	36.210	34.202	-2.008	0.000	396	733
Mother's education: primary	Before	0.779	0.700	-0.079	0.005	399	600
	After	0.750	0.712	-0.038	0.174	396	733
Mother's education: vocational	Before	0.028	0.027	-0.001	0.931	399	600
	After	0.030	0.041	0.011	0.368	396	733
Mother's education: high-school	Before	0.153	0.188	0.035	0.148	399	600
	After	0.144	0.183	0.039	0.097	396	733
Mother's education: university	Before	0.040	0.085	0.045	0.005	399	600
	After	0.076	0.064	-0.012	0.460	396	733
Mother's residence: Capital	Before	0.164	0.185	0.021	0.406	396	590
	After	0.174	0.201	0.027	0.280	390	726
Mother's residence: Town with county rights	Before	0.129	0.154	0.025	0.265	396	590
	After	0.164	0.163	-0.002	0.946	390	726
Mother's residence: Town	Before	0.371	0.286	-0.085	0.005	396	590
	After	0.321	0.313	-0.008	0.788	390	726
Mother's residence: Village	Before	0.336	0.375	0.039	0.214	396	590
	After	0.341	0.324	-0.017	0.557	390	726
Mother's language: Hungarian	Before	0.972	0.980	0.008	0.435	399	600
	After	0.982	0.986	0.004	0.596	396	733
Mother's language: Roma	Before	0.023	0.013	-0.009	0.270	399	600
	After	0.013	0.011	-0.002	0.797	396	733
Mother's language: Other	Before	0.005	0.007	0.002	0.741	399	600
	After	0.005	0.003	-0.002	0.531	396	733
Mother's occupation: Non-manual	Before	0.375	0.433	0.058	0.127	264	450
	After	0.387	0.408	0.021	0.558	297	559
Mother's occupation: Manual	Before	0.625	0.567	-0.058	0.127	264	450
	After	0.613	0.592	-0.021	0.558	297	559
Father's age	Before	38.355	36.588	-1.766	0.000	389	588
	After	38.826	36.667	-2.160	0.000	386	720
Father's education: primary	Before	0.621	0.594	-0.027	0.399	390	588
	After	0.610	0.557	-0.053	0.087	385	720
Father's education: vocational	Before	0.177	0.129	-0.048	0.040	390	588
	After	0.169	0.192	0.023	0.351	385	720
Father's education: high-school	Before	0.128	0.138	0.010	0.668	390	588
	After	0.127	0.143	0.016	0.468	385	720
Father's education: university	Before	0.074	0.139	0.065	0.002	390	588
	After	0.094	0.108	0.015	0.441	385	720
Father's occupation: Non-manual	Before	0.206	0.258	0.053	0.055	399	600
	After	0.217	0.250	0.032	0.222	396	733
Father's occupation: Manual	Before	0.794	0.742	-0.053	0.055	399	600
	After	0.783	0.750	-0.032	0.222	396	733

Table A2: Descriptive statistics for married mothers (treatment group) and for non-married mother (control group)

	Period	Control	Treatment	Diff.	p	N_C	N_T
Birth weight	Before	2888.15	3173.48	285.34	0.000	346	11267
	After	2993.12	3198.84	205.72	0.000	327	13136
Low birth weight (<2500 g)	Before	0.225	0.096	-0.129	0.000	346	11267
	After	0.162	0.089	-0.073	0.000	327	13136
Low birth weight (<3000 g)	Before	0.494	0.299	-0.195	0.000	346	11267
	After	0.462	0.290	-0.172	0.000	327	13136
Infant mortality	Before	0.072	0.034	-0.039	0.000	346	11267
	After	0.028	0.027	0.000	0.996	327	13136
Sex: female	Before	0.477	0.489	0.012	0.663	346	11267
	After	0.511	0.487	-0.023	0.401	327	13136
Mother's age	Before	27.267	27.158	-0.109	0.159	346	11267
	After	27.269	27.118	-0.151	0.056	327	13136
Mother's education: primary	Before	0.777	0.481	-0.297	0.000	346	11267
	After	0.786	0.500	-0.286	0.000	327	13136
Mother's education: vocational	Before	0.046	0.078	0.031	0.031	346	11267
	After	0.080	0.080	0.000	0.990	327	13136
Mother's education: high-school	Before	0.133	0.332	0.199	0.000	346	11267
	After	0.116	0.324	0.208	0.000	327	13136
Mother's education: university	Before	0.043	0.110	0.067	0.000	346	11267
	After	0.018	0.095	0.077	0.000	327	13136
Mother's residence: Capital	Before	0.230	0.219	-0.011	0.621	343	11122
	After	0.196	0.198	0.002	0.939	326	12977
Mother's residence: Town with county rights	Before	0.120	0.198	0.078	0.000	343	11122
	After	0.172	0.194	0.023	0.309	326	12977
Mother's residence: Town	Before	0.303	0.309	0.006	0.821	343	11122
	After	0.316	0.323	0.007	0.780	326	12977
Mother's residence: Village	Before	0.347	0.274	-0.073	0.003	343	11122
	After	0.316	0.284	-0.032	0.212	326	12977
Mother's language: Hungarian	Before	0.945	0.988	0.043	0.000	346	11267
	After	0.939	0.991	0.052	0.000	327	13136
Mother's language: Roma	Before	0.052	0.003	-0.049	0.000	346	11267
	After	0.061	0.003	-0.059	0.000	327	13136
Mother's language: Other	Before	0.003	0.009	0.006	0.247	346	11267
	After	0.000	0.007	0.007	0.133	327	13136
Mother's occupation: Non-manual	Before	0.293	0.558	0.265	0.000	246	9853
	After	0.264	0.523	0.259	0.000	235	11637
Mother's occupation: Manual	Before	0.707	0.442	-0.265	0.000	246	9853
	After	0.736	0.477	-0.259	0.000	235	11637
Father's age	Before	31.773	30.241	-1.532	0.000	214	11257
	After	32.457	30.186	-2.271	0.000	200	13125
Father's education: primary	Before	0.698	0.339	-0.359	0.000	212	11265
	After	0.704	0.340	-0.364	0.000	203	13133
Father's education: vocational	Before	0.165	0.280	0.114	0.000	212	11265
	After	0.222	0.303	0.082	0.012	203	13133
Father's education: high-school	Before	0.080	0.221	0.141	0.000	212	11265
	After	0.049	0.213	0.164	0.000	203	13133
Father's education: university	Before	0.057	0.161	0.104	0.000	212	11265
	After	0.025	0.143	0.119	0.000	203	13133
Father's occupation: Non-manual	Before	0.206	0.344	0.139	0.000	345	11262
	After	0.171	0.323	0.151	0.000	327	13133
Father's occupation: Manual	Before	0.794	0.656	-0.139	0.000	345	11262
	After	0.829	0.677	-0.151	0.000	327	13133

Table A3: Descriptive statistics for young mothers (control group: aged 19-20, treatment group: aged 16-18)

	Period	Control	Treatment	Diff.	p	N_C	N_T
Birth weight	Before	3072.94	2987.82	-85.12	0.000	7429	3407
	After	3099.18	2993.85	-105.33	0.000	8126	3404
Low birth weight (<2500 g)	Before	0.115	0.149	0.033	0.000	7429	3407
	After	0.099	0.139	0.040	0.000	8126	3404
Low birth weight (<3000 g)	Before	0.375	0.444	0.069	0.000	7429	3407
	After	0.358	0.444	0.086	0.000	8126	3404
Infant mortality	Before	0.035	0.036	0.001	0.827	7429	3407
	After	0.024	0.034	0.010	0.003	8126	3404
Sex: female	Before	0.479	0.492	0.013	0.214	7429	3407
	After	0.482	0.488	0.006	0.527	8126	3404
Mother's age	Before	19.531	17.472	-2.059	0.000	7429	3407
	After	19.554	17.498	-2.056	0.000	8126	3404
Mother's education: primary	Before	0.670	0.873	0.203	0.000	7429	3407
	After	0.676	0.860	0.184	0.000	8125	3404
Mother's education: vocational	Before	0.160	0.101	-0.059	0.000	7429	3407
	After	0.162	0.114	-0.048	0.000	8125	3404
Mother's education: high-school	Before	0.168	0.026	-0.142	0.000	7429	3407
	After	0.161	0.026	-0.135	0.000	8125	3404
Mother's education: university	Before	0.002	0.000	-0.002	0.015	7429	3407
	After	0.001	0.000	-0.001	0.052	8125	3404
Mother's residence: Capital	Before	0.130	0.096	-0.033	0.000	7292	3349
	After	0.131	0.101	-0.030	0.000	7997	3351
Mother's residence: Town with county rights	Before	0.149	0.125	-0.025	0.001	7292	3349
	After	0.146	0.117	-0.029	0.000	7997	3351
Mother's residence: Town	Before	0.345	0.346	0.001	0.884	7292	3349
	After	0.350	0.360	0.010	0.305	7997	3351
Mother's residence: Village	Before	0.376	0.433	0.057	0.000	7292	3349
	After	0.373	0.422	0.049	0.000	7997	3351
Mother's language: Hungarian	Before	0.990	0.972	-0.017	0.000	7429	3407
	After	0.989	0.981	-0.008	0.001	8126	3404
Mother's language: Roma	Before	0.006	0.023	0.016	0.000	7429	3407
	After	0.006	0.016	0.011	0.000	8126	3404
Mother's language: Other	Before	0.004	0.005	0.001	0.484	7429	3407
	After	0.005	0.002	-0.003	0.028	8126	3404
Mother's occupation: Non-manual	Before	0.284	0.123	-0.160	0.000	6353	2381
	After	0.274	0.129	-0.145	0.000	7094	2449
Mother's occupation: Manual	Before	0.716	0.877	0.160	0.000	6353	2381
	After	0.726	0.871	0.145	0.000	7094	2449
Father's age	Before	23.672	22.314	-1.358	0.000	7245	3054
	After	23.684	22.285	-1.399	0.000	7910	3116
Father's education: primary	Before	0.374	0.511	0.138	0.000	7248	3064
	After	0.381	0.523	0.141	0.000	7914	3124
Father's education: vocational	Before	0.462	0.407	-0.055	0.000	7248	3064
	After	0.457	0.401	-0.056	0.000	7914	3124
Father's education: high-school	Before	0.141	0.074	-0.067	0.000	7248	3064
	After	0.136	0.069	-0.067	0.000	7914	3124
Father's education: university	Before	0.024	0.007	-0.016	0.000	7248	3064
	After	0.026	0.008	-0.018	0.000	7914	3124
Father's occupation: Non-manual	Before	0.119	0.051	-0.068	0.000	7428	3404
	After	0.118	0.056	-0.062	0.000	8124	3402
Father's occupation: Manual	Before	0.881	0.949	0.068	0.000	7428	3404
	After	0.882	0.944	0.062	0.000	8124	3402

Table A4: The effect of the abortion restriction on birth outcomes, placebo groups

		Treat×After	Robust SE	p	N
A.	Mothers aged 30.5-31.5 (vs. aged 31.5-32.5)				
(1)	Birth weight	-7.736	39.510	0.845	4525
(2)	Low birth weight (<2500 g)	0.008	0.020	0.673	4525
(3)	Low birth weight (<3000 g)	0.021	0.029	0.474	4525
(4)	Infant mortality	-0.012	0.013	0.355	4525
B.	Mothers aged 20-21 (vs. aged 22-23)				
(1)	Birth weight	11.826	11.733	0.314	33198
(2)	Low birth weight (<2500 g)	-0.007	0.006	0.298	33198
(3)	Low birth weight (<3000 g)	-0.016	0.010	0.125	33198
(4)	Infant mortality	-0.005	0.004	0.203	33198

Control variables: see Table 1.

Table A5: The effect of the abortion restriction on birth outcomes, average marginal effects of probit regressions

		Treat×After	Robust SE	p	N
A.	Mothers aged 33.5-34.5 (vs. aged 35.5-36.5)				
(1)	Low birth weight (<2500 g)	0.072	0.030	0.016	2106
(2)	Low birth weight (<3000 g)	0.083	0.041	0.042	2124
(3)	Infant mortality	0.043	0.020	0.031	1919
B.	Married mothers (vs. non-married mothers)				
(4)	Low birth weight (<2500 g)	0.038	0.019	0.046	24911
(5)	Low birth weight (<3000 g)	0.027	0.034	0.431	25057
(6)	Infant mortality	0.034	0.013	0.007	24184
C.	Mothers aged 16-18 (vs. aged 19-20)				
(7)	Low birth weight (<2500 g)	0.013	0.009	0.143	22170
(8)	Low birth weight (<3000 g)	0.027	0.014	0.053	22310
(9)	Infant mortality	0.012	0.005	0.012	21623

Control variables: see Table 1.