Online Appendix

Out of Labor and Into the Labor Force?
The Role of Abortion Access, Social Stigma, and Financial Constraints

Nina Brooks   Tom Zohar

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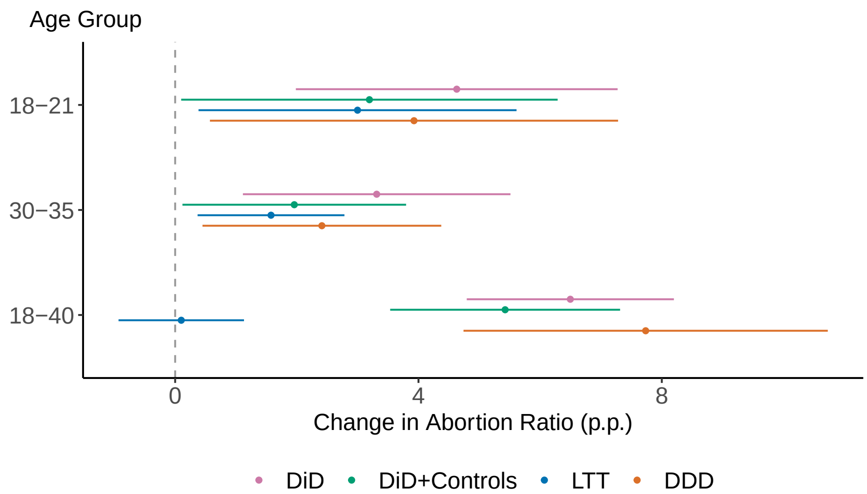
Contents

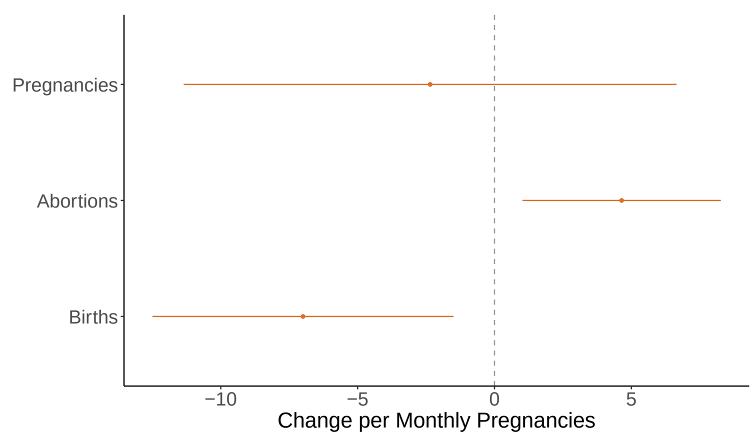
A Additional Figures and Tables

Figure A1: Effect of the 2014 Policy on Abortion Ratios



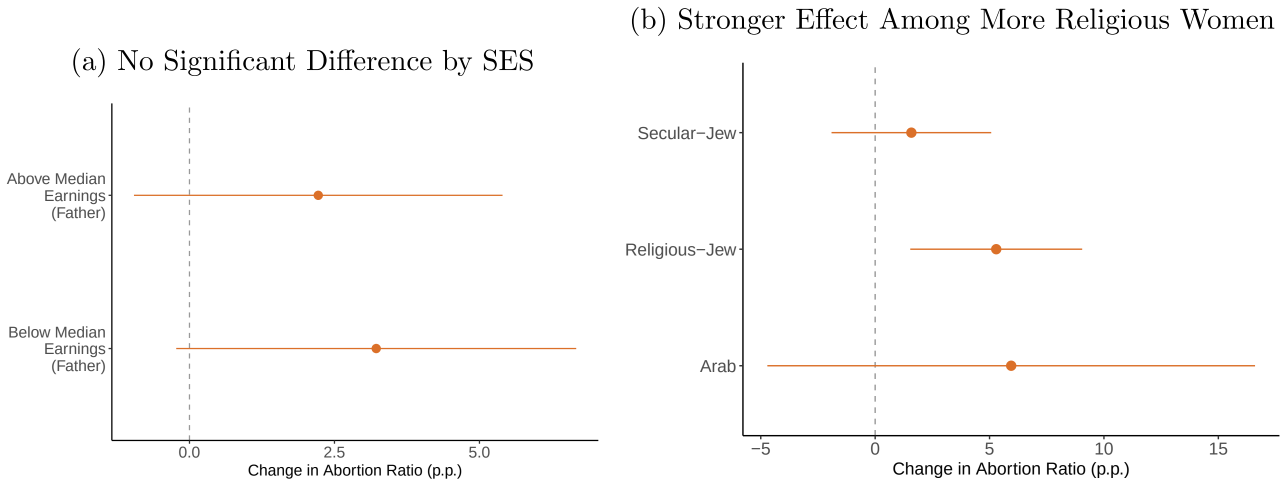
Notes: This figure presents difference-in-difference results for the effect of the 2014 policy on abortion ratios for three different populations (age groups) and specifications. Each color presents the results from a different specification (standard DiD in pink, DiD+Controls in green, a linear time trend specification in blue, and the triple difference in orange). The effect is estimated separately in three samples based on the women's age group: 18-21 (top row, main analytic results), 30-35 (middle row), and 18-40 (bottom row). In each row, the dot represents the treatment effect ($\delta \cdot Post_t \times T_{it}$) and the lines mark the 95% confidence interval around the point estimate. The dashed vertical line is at 0, indicating an insignificant result.

Figure A2: Test for Illegal Market Spillovers



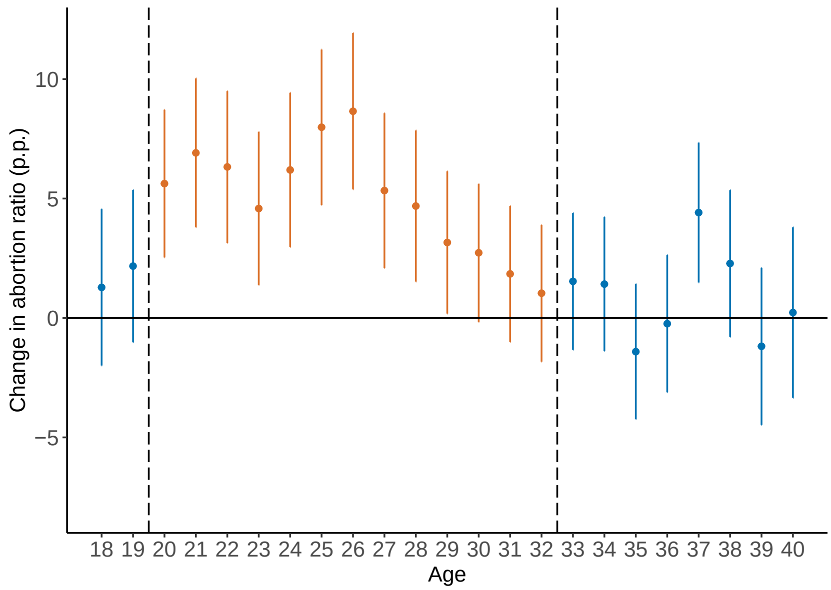
Notes: This figure presents difference-in-difference results for the effect of the 2014 policy on pregnancies, abortions, and births in levels, after collapsing the data to the year-month-age level. In each row, the dot represents the treatment effect ($\delta \cdot Post_t \times T_i$) and the lines mark the 95% confidence interval around the point estimate. The dashed vertical line is at 0, indicating an insignificant result. The sample includes all unmarried women in the country aged 18-21 from 2009-2016. Treated women are those aged 20-21. The estimates are percentage point changes that can be interpreted as the relative change per 100 pregnancies.

Figure A3: Effect of the Policy by SES and Ethnicity



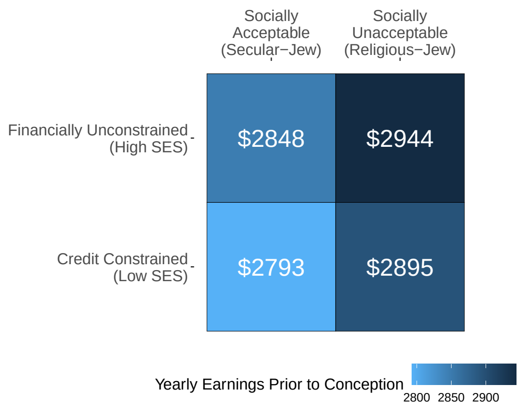
Notes: The figure presents the heterogeneous difference-in-difference results, where we split the sample by population groups. Panel A3a is split into two SES groups according to the yearly earnings of the father of the woman who conceived (our proxy for household SES level). Panel A3b is split by ethnic groups. In each row, the dot represents the percentage change of the treatment effect ($\delta \cdot Post_t \times T_i$), and the lines mark the 95% confidence interval around the point estimate. The dashed vertical line is at 0, indicating an insignificant result. The sample includes all unmarried women in the country aged 18-21 from 2009-2016 who conceived. Treated women are those aged 20-21.

Figure A4: Difference Between Pre and Post Policy Abortion Ratios by Age (Raw Data)

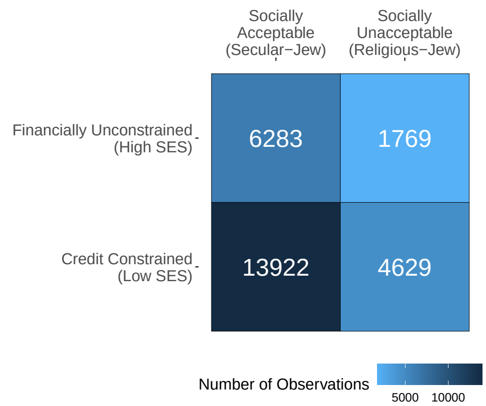


Notes: This figure presents the results of a before-and-after exercise in which we restrict the data to two years before and after the 2014 policy change (2012-2015) and estimate the post-policy difference in abortion ratios separately for each age (18-40). The point estimates can be interpreted as the difference in abortion ratio for each age group following the introduction of the 2014 policy. The lines are 95% confidence intervals and the horizontal line marks 0. The ages that were eligible for the 2014 subsidy expansion are indicated in orange (treated), while those ineligible are presented in blue. The dashed vertical lines mark the two age cutoffs for the subsidy change eligibility: 19-years-old and 33-years-old.

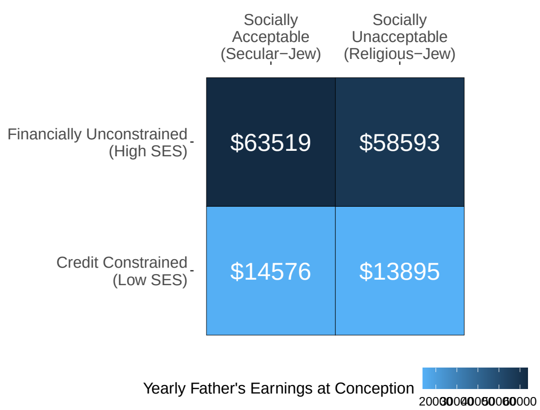
Figure A5: Baseline Characteristics by Religiosity and SES



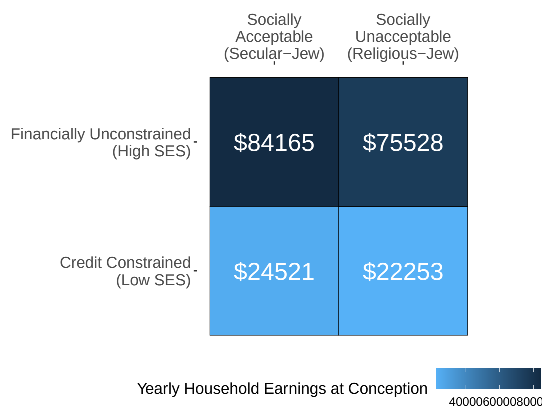
(a) Woman's Earnings



(b) Number of Observations



(c) Father's Earnings



(d) Household's Earnings

Notes: This figure presents the baseline characteristics of the analytic population split by religiosity level and SES background (based on parental earnings). Panel (a) presents the mean yearly earnings of the woman who conceived, one year prior to conception; Panel (b) presents the number of observations in each cell; Panel (c) presents the mean yearly earnings of the woman's father at the year of conception; Panel (d) presents the mean yearly earnings of the woman's household (sum of both parents earnings) at the year of conception. Darker blues correspond to higher levels in each panel.

Table A1: Selection Into Abortion

	Abortion			Birth			Difference	P-value	Full Sample		
	Mean	SD	N	Mean	SD	N			Mean	SD	N
Age	28.0	6.7	241,625	28.5	5.3	2,276,942	0.5	0.000	28.4	5.5	2,518,567
Married	0.41	0.49	241,625	0.90	0.30	2,276,942	0.49	0.000	0.86	0.35	2,518,567
Total Kids	1.1	1.2	241,625	1.3	1.1	2,276,942	0.2	0.000	1.3	1.1	2,518,567
Months Worked Pre	5.0	4.5	241,625	5.4	4.7	2,276,942	0.4	0.000	5.4	4.7	2,518,567
Yearly Earnings Pre	32,327	49,502	241,625	37,369	52,356	2,276,942	5,042	0.000	36,885	52,110	2,518,567
Monthly Earnings Pre	3,513	4,872	241,625	3,904	5,381	2,276,942	391	0.000	3,867	5,336	2,518,567
USSR Descendent	0.23	0.42	241,625	0.10	0.30	2,276,942	-0.13	0.000	0.11	0.32	2,518,567
Religious-Jew	0.07	0.26	241,625	0.28	0.45	2,276,942	0.21	0.000	0.26	0.44	2,518,567
Arab	0.10	0.30	241,625	0.22	0.41	2,276,942	0.11	0.000	0.21	0.40	2,518,567
Years of Education	12.0	2.9	241,625	12.4	2.9	2,276,942	0.5	0.000	12.4	2.9	2,518,567
less than HS	0.72	0.45	241,625	0.63	0.48	2,276,942	-0.08	0.000	0.64	0.48	2,518,567
HS + Vocational	0.10	0.31	241,625	0.10	0.30	2,276,942	-0.01	0.000	0.10	0.30	2,518,567
BA+	0.18	0.38	241,625	0.27	0.44	2,276,942	0.09	0.000	0.26	0.44	2,518,567
High-earnings	0.35	0.48	241,625	0.40	0.49	2,276,942	0.06	0.000	0.40	0.49	2,518,567
Low-earnings	0.04	0.20	241,625	0.04	0.18	2,276,942	-0.00	0.000	0.04	0.19	2,518,567
Nonemployed	0.36	0.48	241,625	0.35	0.48	2,276,942	-0.01	0.000	0.36	0.48	2,518,567

Notes: This table presents summary statistics of key variables separately for the samples of women who gave birth and aborted, as well as the overall population of conceptions (abortion + birth). All statistics are calculated pre 2014, when the reform took place. The first three columns present the mean, standard deviation, and sample size of characteristics for women who had an abortion. The next three columns present the same for women who gave birth. Next, we report the difference in means and the p-value for a t-test of the difference. Finally, in the last three columns we report the mean, standard deviation, and sample size for the full sample of women.

Table A2: Family earnings rank vs child earnings rank

	Family Earnings Measure		
	Household	Father	Mother
Coefficient	0.23	0.246	0.093
	(.003)	(.003)	(.003)
Obs	156555	156555	156555
R^2	0.049	0.055	0.008

Notes: This table presents the results from the the rank correlation between child's earnings rank and household, father's and mother's earning ranks. Both parent's and child's earnings are the residuals from a regression of age, age-squared and year fixed effects on log earnings.

Table A3: Effect on Downstream Social and Economic Outcomes for All Unmarried 18-21 Year Olds

	OLS	IV	RF	Mean	N
Is a parent	-0.847	-0.71	-0.033	37.7%	24,650
	(0.016)	(0.192)	(0.004)		
Married	-0.307	-0.785	-0.036	18.8%	24,650
	(0.005)	(0.242)	(0.02)		
Age at 1st Birth	3.825	5.263	0.41	21.27	15,369
	(0.04)	(0.78)	(0.062)		
Number of children	-1.438	-1.665	-0.077	0.6	24,650
	(0.047)	(0.141)	(0.016)		
BA Enrollment	0.021	0.022	0.001	5.7%	24,650
	(0.013)	(0.101)	(0.002)		
Working	0.26	-0.247	-0.011	75.8%	24,650
	(0.007)	(0.213)	(0.008)		
Employed by a firm	0.261	-0.129	-0.006	74.5%	24,650
	(0.008)	(0.152)	(0.006)		
Employed part-time	-0.175	0.036	0.002	72.8%	24,650
	(0.03)	(0.174)	(0.008)		
Self-employed	-0.001	-0.118	-0.005	1.2%	24,650
	(0.001)	(0.061)	(0.002)		
Earnings (NIS, Cond.)	13351.295	-4229.73	-176.258	26,865	22,340
	(2441.9)	(9475.7)	(355)		
Sector's Wage Premium	0.018	0.054	0.003	0.032	22,340
	(0.002)	(0.012)	(0.001)		

Notes: $^{*}p< 0.05$, $^{**}p< 0.01$, $^{***}p< 0.001$. Standard errors clustered by age at conception in parentheses. This table presents results for the effect of the 2014 policy on a range of human capital formation and labor market outcomes. The first column presents the naive OLS (Equation 2), the second column presents results from the IV (Equation 3), and the third column presents results for the reduced from (Equation 5). The sample includes all unmarried, 18-21 year old women.

B Israel Context

B.1 Abortion Committee

Israel's abortion committee process is fairly unique in the global context. The committee process for obtaining an abortion was motivated by medical concerns that abortion could affect a woman's future fertility (Amir, 2015). The concern over fertility was aligned with Israel's demographic project, which aims to reverse the decrease in the global Jewish population due to the Holocaust.[1] In pursuit of the demographic project, Israel has adopted aggressive, pro-natalist policies such as subsidized daycare, monthly child allowances, tax deduction, paid parental leave, and covered infertility treatments and oocyte cryopreservation (egg freezing) under the national health insurance system. Likewise, contraception is not covered by the national health insurance, and abortions are illegal without prior approval from the committee. Consequently, Israel's birth rate is the highest among the developed world.

Israel has a national health-care system which oversees the abortion committee. People may choose to seek care from private providers but if a woman decides to have an abortion through a private doctor, she is still required by law to go through the committee process for approval. A total of 42 abortion committees exist in Israel. They are contained in each of the national hospitals, and the rest operate in smaller clinics (either private or public).[2] Typically, once a woman becomes pregnant and is interested in having an abortion, her doctor directs her to make an appointment with a committee.

The committee is composed of two medical professionals and a social worker, one of whom must be a woman. The committee approves the abortion if *at least* one of the following conditions is satisfied: (1) the woman is under 18 or over 40 years of age; (2) the pregnancy is out of marriage; (3) the pregnancy is the result of an illegal act (rape or incest); (4) the pregnancy risks the life or the health of the woman; or (5) the fetus suffers from congenital disorders. These criteria are largely motivated by Jewish law (Amir, 2015). Judaism holds relatively liberal views with respect to abortion compared to Islam and Christianity (19% and 2% of the population, respectively). Jewish law emphasizes the mother's life and health, whereas the fetus is part of the mother and does not have its rights before it is born. Based on Jewish religion, a child outside of wedlock is considered to be illegitimate and is doomed to bad life outcomes (e.g., cannot get married by Jewish tradition), and should therefore be avoided.

Upon arrival at the committee, the woman fills out the necessary paperwork and pays the committee fees.[3] Then, she meets the social worker to discuss her decision and assess her eligibility per the criteria. The committee's social workers serve as the effective gatekeepers to the approval process, while the committee itself serves as a rubber stamp (Oberman, 2020). Given the criteria for a legal abortion, only one group of women is ineligible: married women between 18 and 40 with healthy pregnancies. In cases when a woman is ineligible based on the official criteria, but desires an abortion, the social worker often helps her navigate the system to meet the criteria. Most commonly, the social workers refers these women to a psychiatrist who can assert that the woman is not adjusting to the pregnancy, which allows her to obtain approval under the criteria for protecting women's health, which includes mental health (Oberman,

[1] "increasing the Jewish birth rate is in dire need..." David Ben-Gurion, Israel's first Prime Minister.

[2] See the full list here

[3] The committee fee is 400 NIS (or $155), which was also eliminated by the 2014 policy.

2020).[4]

Although the committee process may seem obstructive, given this process effectively, all applications are approved (unsurprisingly). Our data shows that 99% of applications were approved, and 97% are acted upon. The application and committee process is completely confidential for the woman, and neither parental nor partner consent is required. Women past the 24th week of pregnancy (1% of abortions) are referred to a select committee that reviews the request and has stricter standards for approvals (though approval rates are as high as the standard committee), which we exclude from this analysis. Oberman (2020) attributes the high approval rates to the social workers who direct women who would otherwise not be approved to a psychiatrist for sign-off under the women's health criteria.

B.2 Abortion and Contraceptive Use Norms and Prevalence

Abortion is not uncommon in Israel, despite the existence of the committee, as shown in Figure B1. The share of legal abortions out of pregnancies has remained relatively constant between 2002-2016, averaging approximately 8% of pregnancies each year and 10% overall. While this may sound high, Israel's legal abortion ratio is actually relatively low compared to global rates and other high income countries (see Figure B2). Twenty-five percent of pregnancies are aborted worldwide, while in Europe the rate is 26% and in North America the rate is 16% (Guttmacher, 2018).

Abortion is not a politically charged topic in the Israeli parliament (*Knesset*) relative to other settings, such as the United States, and is considered to be a "silenced phenomena" (Amir, 2015). While the liberal parties apposes the existence of the abortion committee, they know that opening the topic for discussion might result in a more restrictive abortion laws (Oberman, 2020; Rimaltt, 2017). On two rare occasions that a challenging bill was introduced it did not pass.[5] On the other side of the debate, the religious parties (both Orthodox and Arab parties) tried twice to challenge the status-quo of the current abortion law but could not gather the necessary political support.[6]

To illustrate the difference in abortion discussion in the public sphere between Israel and the U.S., and the salience of the policy we conduct a Google trends analysis that suggests the policy took time to arrive to the public discussion and even when it did, it was much less salient as compared to other abortion discussions in the U.S. Figure B3 presents the Google searches for the word "abortion" in the U.S. compared to it's Hebrew equivalent ("hapala") in Israel from 2009-2019 (normalizing base levels of both countries in January 2009). We can see the peak in Israel across time is indeed in 2014, but only a few months into the year (while the policy was

[4] Alternatively, some anecdotal evidence suggests women can report taking certain medications before getting pregnant, which puts the fetus at risk.

[5] At November 29th 2004, Reshef Hen ('Shinoi' party) submitted a legislative application for adding approval based on SES status. At 2006 Zehava Galon ('Meretz' party) propose to reconsider the committee practice altogether. Both bills did not pass.

[6] In 2008, Nissim Zeev ('Shas' party) proposed to make late-term abortions illegal. The bill failed not pass. In 2013 the two Chief Rabbis of Israel issued a letter in support of Efrat, an anti-abortion group that was established in the 1960's. In January 2017 Yehudah Glick (Likud party) and Abd al-Hakim Hajj Yahya (Joint List party) called a *Knesset* Committee on the Status of Women and Gender Equality meeting to propose incorporating religious representative in the abortion committee, the law failed to pass arguing that a religious entity in the committee will discourage women (especially in the Arab population) to apply to the committee, fearing information leaking to their communities.

in effect in January already). Having the intensity of the American abortion discussion, one can imagine that this extensive coverage will create an out of the ordinary discussion. However, we can see the surge in searches of the word "Abortion" responds much more aggressively in the U.S. even without a change of the law, but just when President Trump was elected, or when Brett Kavanaugh was nominated to the supreme court.

Israel is an interesting setting to study the effects of abortion access due to the vast heterogeneity in cultural views toward abortion, ranging from: immigrants from the former USSR and other secular Jews, religious and ultra-Orthodox Jews, and Christian and Muslim Israeli-Arab.[7] Figure 1c presents the wide heterogeneity in baseline abortion ratios which might suggest different latent costs of abortion (or differing abortion views) across groups. Understanding these heterogeneous views is key to understanding our results and our proposed mechanism - social-stigma cost, that will be discussed in Section 5.2.

Compared to Islam and Christianity, Judaism holds relatively liberal views with respect to abortions and accepts abortions in two broad cases: threat to the woman's life and if the fetus will be born into an "unstable life".[8] These two conditions are the motivation for the abortion committee criteria discussed above (Amir, 2015). Judaism place a supreme value on mother life and health, while the fetus is part of the mother and doesn't have it's own rights before it is born. Nevertheless, Israel's (Jewish) "demographic project" strives to limit abortions among the Jewish population (motivating the abortion committee), which are considered anti-Zionist. The Jewish population consists of a wide mixture of religiosity levels, ranging from secular Jews (45%), traditional Jews (25%), religious Jews (16%) and Orthodox Jews (14%) (Central Bureau of Statistics (Israel), 2018).

Broadly speaking, religiosity level is highly correlated with marriage age, fertility levels, contraceptive use, and opposition to abortion. The secular-Jewish population generally supports abortion, has relatively high contraceptive use rates (Figure B4), and has relatively low fertility rates, while the other end of the religiosity spectrum the Orthodox and Ultra-Orthodox populations are opposed to abortion, have low contraceptive use rates (Figure B4) and very high fertility rates. The Jewish religious populations also tends to marry and have children at a very young age (late teens to early 20s). For example, among 18-21 year-olds, 79% of the women who conceived are married. The abortion ratio (out of pregnancies) among 18-21 unmarried women is 67%, while it is essentially 0% for the married population in this age group. Thus, it is safe to assume that most pregnancies among the married population in this age group are planned, while the converse is true among unmarried populations. Immigrants from the former USSR make up 15% of the total population in Israel (Leshem, 2009), are both Jewish and predominantly secular (Leshem, 2009), and have extremely lenient attitudes toward abortion due to USSR's unique history with respect to abortion and contraception.

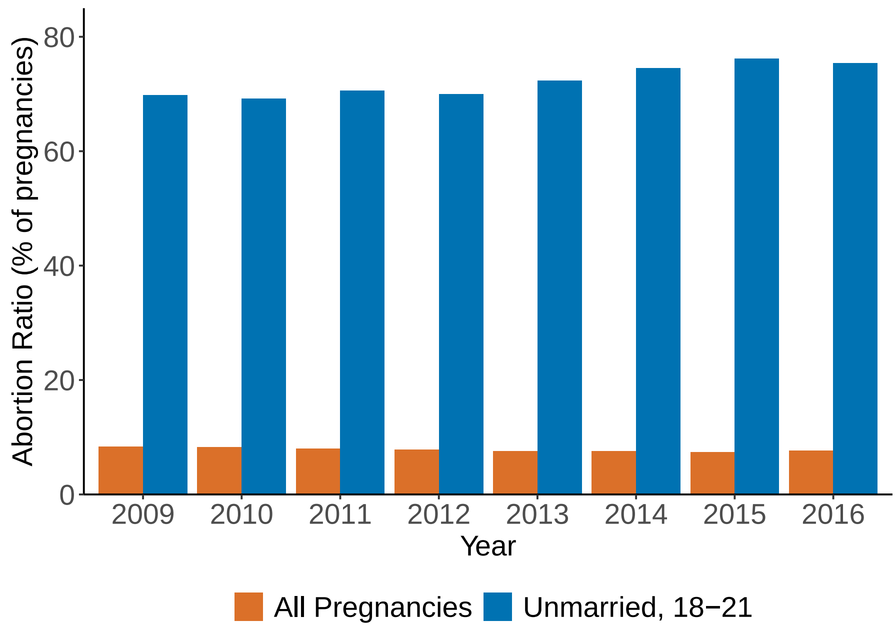
The Israeli-Arab population is mostly religious and considers abortion very taboo. The Muslim population consists of 11% secular, 57% traditional and 31% religious Muslims (Central Bureau of Statistics (Israel), 2018). In general, Islam opposes abortions other than in cases of a health risk for the child. Given the opposition to abortion, there may be a greater incentive for women in the Arab-Muslim community in Israel to turn to the illegal market or even self-induce abortion. The extent of self-induced and illegal abortions is extremely difficult to

[7] Israel is composed of 75% Jews, 18.6% Arab-Muslims, 2% Arab-Christians and 4.4% affiliated with other religious groups (or non-affiliated).

[8] There is no strict definition of an "unstable life", but characteristics of an unstable life may include cases such as unmarried parents, an extremely old or young mother, or being born with congenital disorder.

estimate, but a study of the Palestinian population may help contextualize the frequency with which this occurs among the Arab-Muslim population within Israel.[9] A 2006 survey conducted by Bethlehem University found that 10% of Palestinian-women self induced abortions and a quarter of the women stated it was necessary for unmarried women in order to prevent "honor killings" (Foster et al., 2007).[10]

Figure B1: Abortion in Israel Over time



Notes: This figure presents the abortion ratio (share of abortions per pregnancies) in Israel per year for each year between 2009-2016. The bars in orange presents the rates for the entire population of pregnant women, while the bars in blue restricts to our population of interest: unmarried, 18-21 year olds.

B.3 Child Rearing Cost

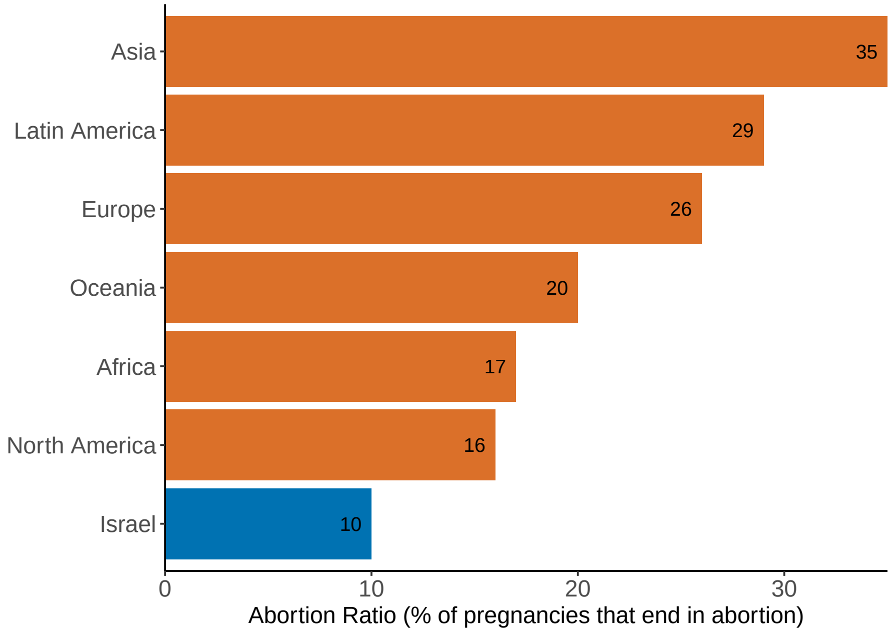
The cost of child-rearing in Israel is substantially lower than the cost in the United States. First, both education and healthcare are free in Israel.[11] Second, as part of the demographic agenda of the country there are several income transfers tools: birth transfers ($145 - $484 for birth), tax breaks ($118 a month per child per working parent) and social security transfers ($41-$52 a month per child). Finally, one should consider the opportunity cost of having a child - the woman's displacement out of the labor force. By the Israeli law, the government covers three months of paid maternity-leave with an option to extend three months unpaid, during which employers cannot fire the parent (only one of the parents can choose to take the leave,

[9]The Israeli-Arab community commonly identifies with the Palestinian population (Tamar-Sheperman, 2008).

[10]The practice of killing women by other family members when the women have brought dishonor to the family, for example by having an abortion or having premarital or extramarital sex.

[11]Healthcare coverage per child is merely $3 per month and Pre-K is free for children 3 years of age and older. Below the age of 3 free options exist, yet even the paid options are less than the minimum full-time monthly earnings.

Figure B2: Abortion Ratios Worldwide

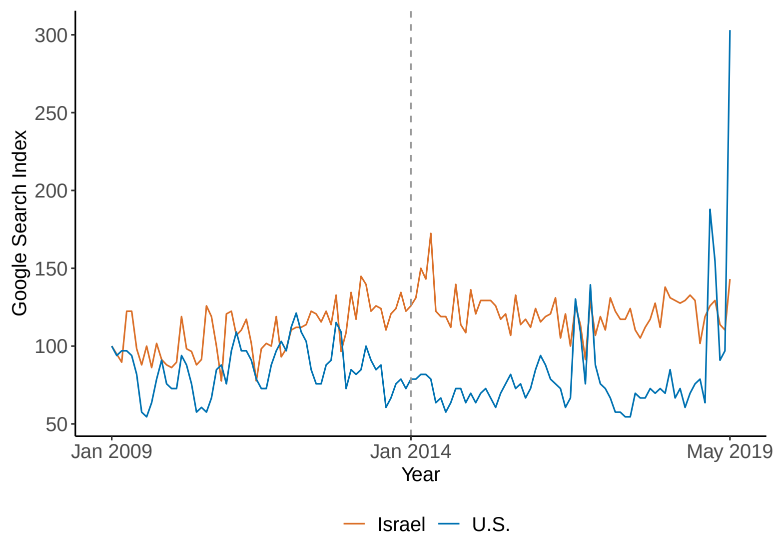


Notes: This figure presents the abortion ratio (share of abortions per pregnancies) across regions of the world and in Israel. Global data come from the Guttmacher Institute (https://data.guttmacher.org/) and the Israeli data are from the Central Bureau of Statistics used in the primary analysis.

though it is commonly the mother). In some cases the parent may ask for an unpaid extension of the maternity leave, but the employer is no longer obligated by law to do so.

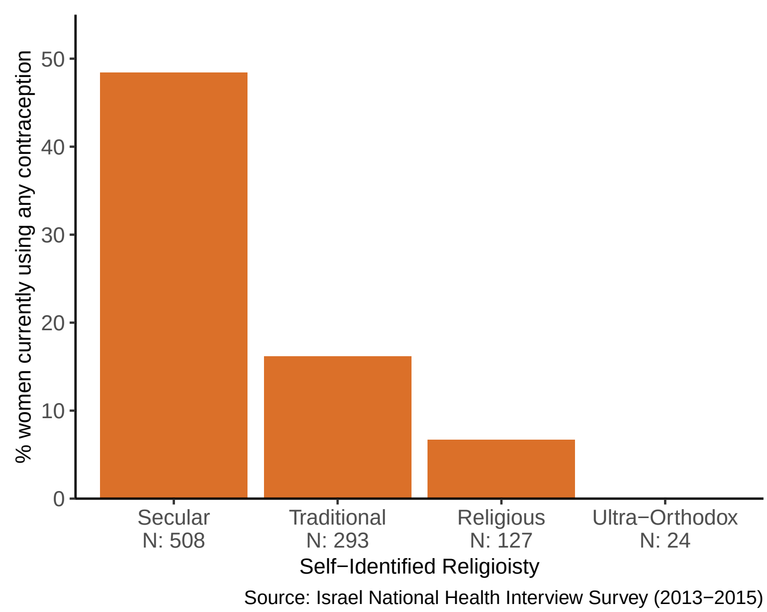
One important feature of Israel - it's size - allows a smoother return to the labor force due to support of the family members in those months where the child is still out of the education system (typically until age one years old). Still, child-rearing costs is a substantive consideration when taking a decision to abort.

Figure B3: Google Search for "Postinor (Plan B)" and "Abortion"



Notes: This figure presents Google searches for the word "abortion" in the United States compared to its Hebrew equivalent ("hapala") in Israel from 2009-2019 (normalizing base levels of both countries in January 2009).

Figure B4: Self-reported contraceptive use by religiosity



Notes: This figure presents self-reported contraceptive use by self-reported religiosity, collected by the Israel National Health Interview Survey, 2013-2015 (Einav et al., 2017). This data includes Jewish and Muslim women, who self-identified their religiosity level. The total number of women surveyed in each group is reported. Thus, these categories do not perfectly align with the religiosity level for Jewish women we constructed for our analysis. In this survey, no Ultra-Orthodox women reported currently using a method of contraception.

C Parallel Trends

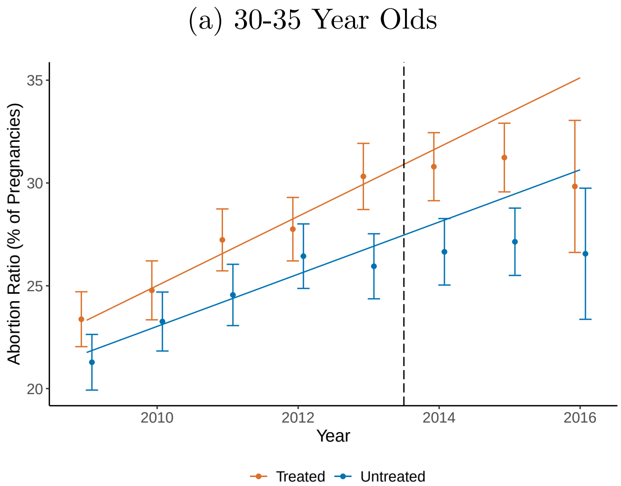
Our difference-in-differences approach requires that women eligible for the subsidy would have experienced similar changes in their abortion ratio over time as ineligible women in the absence of the 2014 subsidy. In this section we present several pieces of evidence to support the plausibility of parallel trends for the population of interest (unmarried, 18-21 year olds), as well as comparable plots for two other populations: unmarried 30-35 year olds, where the 33-year-old age cutoff was used to determine treatment, and unmarried women aged 16-40 with both age cutoffs (19-years-old and 33-years-old) used to define treatment.

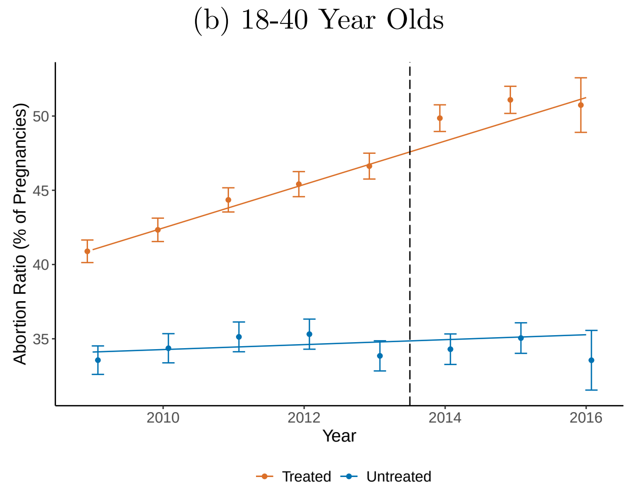
Figure C1 presents comparable versions of Figure 2 for two alternative age groups: unmarried 30-35 year olds and the entire sample of unmarried 18-40 year olds. Figure 2 shows parallel pre-trends for the treated and untreated groups and a clear increase in the abortion ratio relative to the counterfactual pre-trend for the newly eligible 20-21 year olds after the policy was implemented. In contrast, Figure C1a presents the comparable parallel trends for the 30-35 year old sample. For this age group, the trends are not parallel as the abortion ratio for the eligible 30-32 year olds increased more over time than that of the ineligible 33-35 year olds. There also does not appear to be much of a treatment effect among the eligible population after 2014.

Figure C1b presents the comparable parallel trends for the 18-40 year old sample, where women affected by the policy change (ages 20-32) are considered treated and unaffected women (ages 18-19 and 33-40) are considered untreated. The trends are clearly not parallel: the abortion ratio for the treated group increased over time, while the abortion ratio for the ineligible group is essentially flat. Given the vast number of ages we included (18-40) it is reasonable that trends in abortion ratios would be very different, nonetheless this is also a clear violation of the parallel trends assumption.

Based on this assessment, we choose to focus on the younger population of women for whom the difference-in-difference is valid, but we present a comparable set of results using the 33 year old cutoff among the sample of 30-35 year olds and the sample of 18-40 year olds in the appendix for completeness. This population of 18-21 year olds is also extremely relevant from a policy perspective. Due to the two-year mandatory military service in Israel, women usually start higher education at the age of 21-23. Thus, 18-21 year olds are deciding whether or not to obtain higher education and the ability to avoid an undesired birth at this time could have important impacts on this decision, whereas 30-35 year olds have largely completed their educational attainment. Additionally, 18-21 year olds are less financial independent and the subsidy is likely more valuable for them. Finally, given familial and cultural pressures, 18-21 year-olds may also be more likely to have a “shotgun marriage” in response to an undesired birth, whereas the majority of women in Israel are already married by the time they reach ages 30-35.

Figure C1: Parallel Trends Assessment By Age Group





Notes: This figure presents the difference in abortion ratios between treated and control over time (2009-2016) for three different age groups. In panel (a), we present the parallel trends for unmarried, 30-35 year olds, where the treated group is women aged 30-32 and the control group is women 33-35. In Panel (b), we present the parallel trends for all unmarried, 18-40 year olds, where the treated group is women aged 20-32 and the control group is women 18-19 and 33-40. In both panels, the dashed line indicates the timing of the 2014 policy change. Each dot represents the mean abortion ratio in a given month-year for the eligible and ineligible groups of women, respectively. The linear lines are fitted separately before and after the policy change for each group. The ineligible population (control) is presented in blue, and the eligible population (treated) is presented in orange.

To address the potential risk to identification that women aged 18-19 were on a different (time) trend then women aged 20-21, we run specifications in which we first residualize the abortion outcome on separate pre-trends for the control and treated groups (Equations 1 and 2), and then run the standard DiD on the residualized abortion (Equation 3):[12] We present the results of this linear-time trend specification in Table 3 for the main population of unmarried 18-21 year olds, and in Figure A1 for all age groups for comparison.

$$abort_{it}^{Pre} = \beta_{Pre}^{T} \times T_i \times t + \nu_{it} \tag{1}$$

$$\tilde{abort}_{it} \equiv abort_{it} - \widehat{\beta_{Pre}^{T}} \times T_i \times t \tag{2}$$

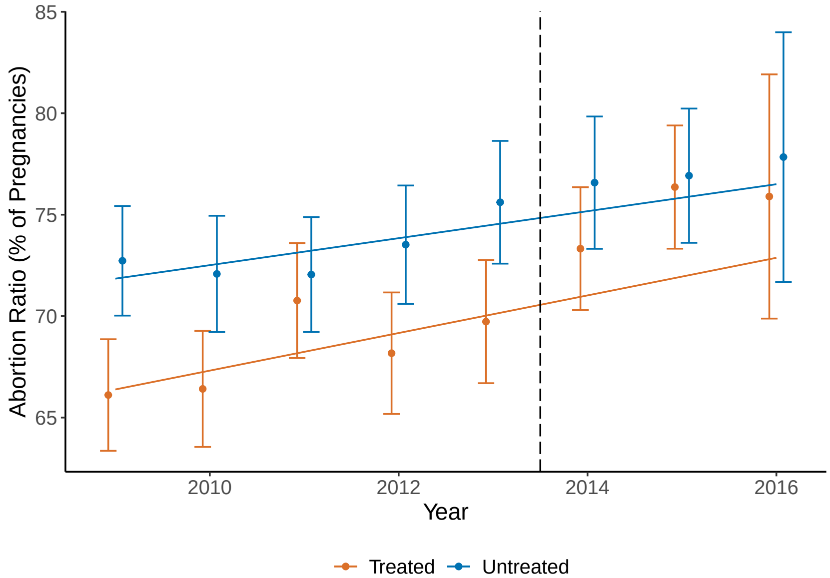
$$\text{Resid. Pre-trends: } \tilde{abort}_{it} = \delta \cdot Post_t \times T_i + \gamma_{a_i} + \gamma_{y_t} + \gamma_{m_t} + X'\gamma_i + \epsilon_{it} \tag{3}$$

To address two further potential identification concerns, we assess parallel trends in two additional groups. First, to restrict to an even narrower age bandwidth around the 19-year-old age cutoff, we plot the mean abortion ratio at the month-year level for 19-20 year olds, where 19 year olds represent the untreated (ineligible) group and 20 year olds represent the treated (eligible) group. This narrower age bandwidth ensures the treated and control populations are as similar as possible in their fertility trends (in the absence of the policy). As Figure C2 demonstrates, the parallel trends assumption appears to hold more strongly within this

[12] Since there is no straightforward way to calculate standard errors in this case, we calculate them using 1,000 bootstrapped replications.

narrower age bandwidth than in the 18-21 year olds, consistent with our prior. However, by doing so we lose power and the standard errors increase significantly. Therefore, to maintain statistical power, and because parallel trends also holds for 18-21 year olds (Figure 2), we use the entire population of unmarried 18-21 year olds.

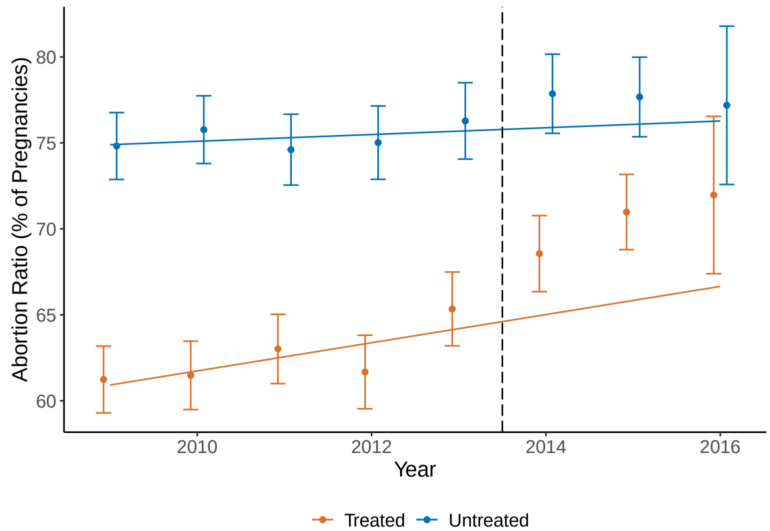
Figure C2: Parallel Trends: Only 19 and 20 Year Olds



Notes: This figure presents the abortion ratios for the treated and control groups over time (2009-2016) for 19-20 year olds. In this population, the treated group includes women aged 20 and the control group includes women aged 19. The dashed line indicates the timing of the 2014 policy change. Each dot represents the mean abortion ratio in a given month-year for the eligible and ineligible groups of women, respectively. The linear lines are fitted separately before and after the policy change for each group. The ineligible population (control) is presented in blue, and the eligible population (treated) is presented in orange.

Second, some 20 year olds serve in the military for a few months past their 20th birthday, which fully covers all medical procedures including abortion, and would thus contaminate the treatment. To address this concern, we assess the plausibility of using 21 and 22 year olds as the treated group (compared to 18-19 year olds as the control group) to determine whether it is reasonable to estimate specifications that exclude all 20 year olds. We present this version of the parallel trends in Figure C3.

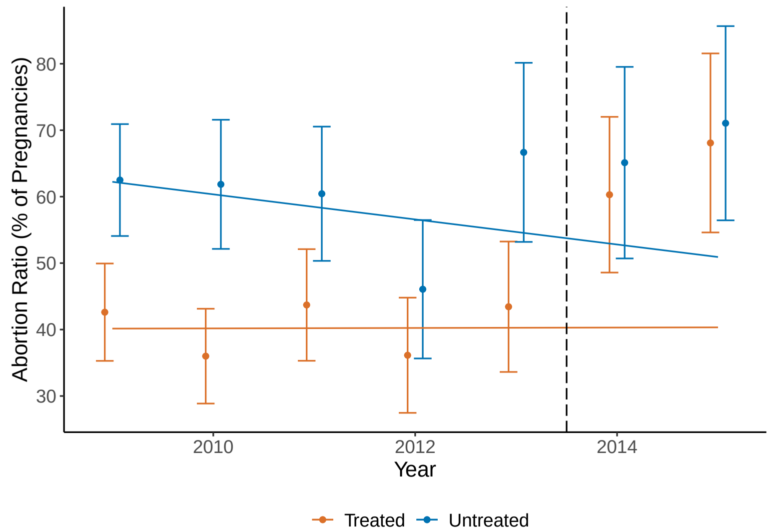
Figure C3: Parallel Trends: Excluding 20 Year Olds



Notes: This figure presents the abortion ratios for the treated and control groups over time (2009-2016) for the 18-19 year olds compared to 21-22 year olds (excluding 20 year olds to account for women finishing their military service at different points of time). In this population, the treated group includes women aged 21-22 and the control group includes women aged 18-19. The dashed line indicates the timing of the 2014 policy change. Each dot represents the mean abortion ratio in a given month-year for the eligible and ineligible groups of women, respectively. The linear lines are fitted separately before and after the policy change for each group. The ineligible population (control) is presented in blue, and the eligible population (treated) is presented in orange.

As we demonstrate in Section 5, the increase in the abortion ratio is driven by the population of socially and financially constrained women (ie., Jewish women from low-income, religious households). Because we restrict the analysis of downstream socioeconomic effects to this sub-population in Section 6, it is important to assess whether parallel trends holds for this specific subgroup as well. Figure C4 presents

Figure C4: Parallel Trends: Socially and Financially Constrained



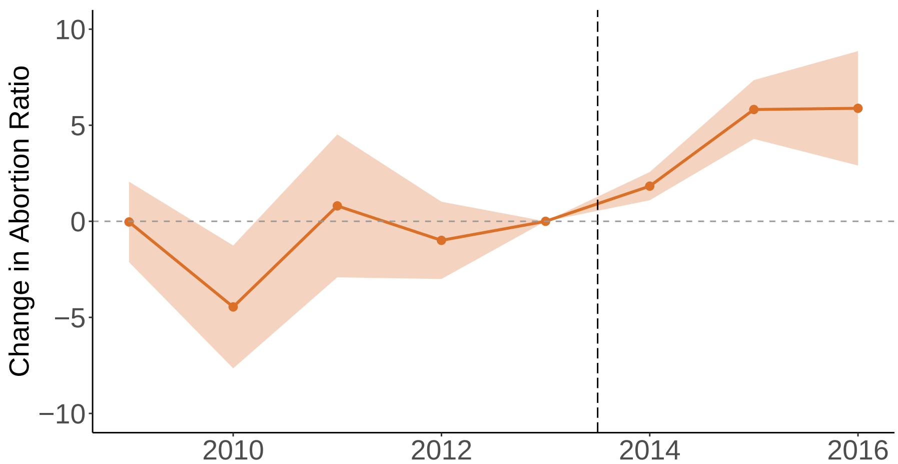
Notes: This figure presents the abortion ratios for the treated and control groups over time (2009-2016) for sub-population of socially and financially constrained women (below median father's income and religious Jew). In this sub-population, the treated group includes women aged 20-21 and the control group includes women aged 18-19. The dashed line indicates the timing of the 2014 policy change. Each dot represents the mean abortion ratio in a given month-year for the eligible and ineligible groups of women, respectively. The linear lines are fitted separately before and after the policy change for each group. The ineligible population (control) is presented in blue, and the eligible population (treated) is presented in orange.

Finally, as an alternative approach to testing for parallel trends, we interact the treatment status (T_i) with a dummy for each year in our sample (for years $k \in \{2009, 2016\}$) using the following equation:

$$abort_{it} = \sum_{k=2009}^{2016} \delta_k \times \mathbb{1}\{t = k\} \cdot T_i + \gamma_{a_i} + \gamma_{y_t} + \epsilon_{it} \tag{4}$$

Figure C5 plots the estimates of δ_k from Equation 4. The estimates represent the difference in abortion ratio between treated (aged 20-21) and untreated (aged 18-19) women over time (2009-2016), with the 2013 difference dropped as the reference year. The shaded regions mark the 90% and 95% confidence intervals around the point estimate, respectively. We see no statistical difference in the abortion ratio between treated and control women before the policy change (which supports the parallel trends assumption); after 2014, the abortion ratio among treated women increased.

Figure C5: Generalized Difference-in-Difference



Notes: This figure presents the difference in abortion rate between treated (aged 20-21) and control (aged 18-19) women over time (2009-2016). The dashed line indicates the timing of the 2014 policy change. Each dot represents the coefficient δ_k estimated from the generalized difference-in-differences (Equation 4). Note that 2013 is dropped as the reference year. The shaded regions mark the 95% confidence interval around the point estimates, respectively. The sample includes all unmarried women in the country aged 18-21 who conceived from 2009-2016.

D Toy Model

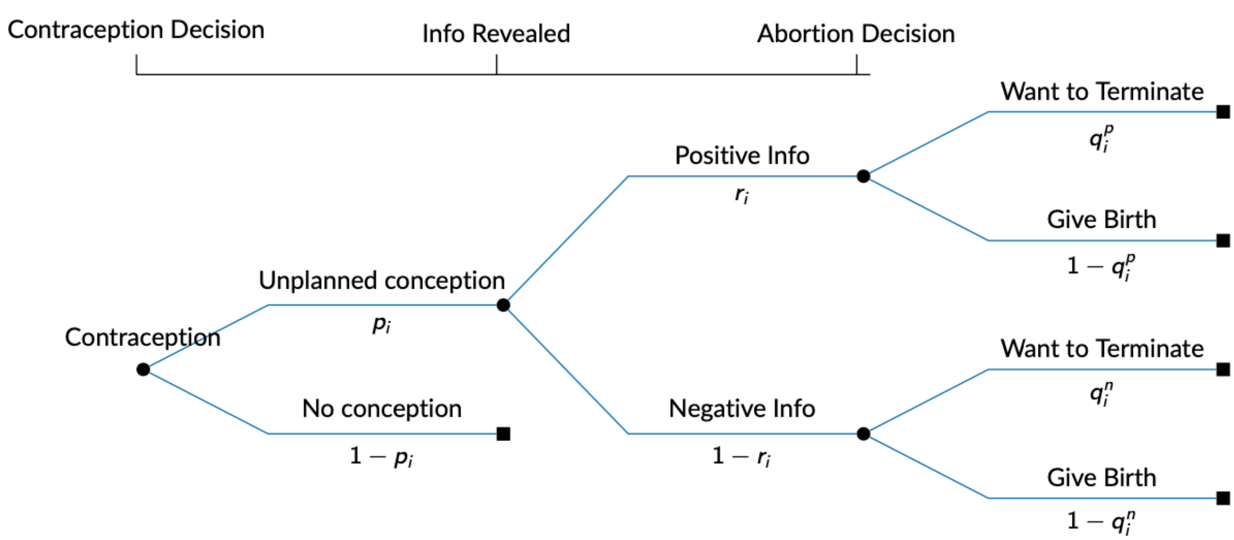
In this Section we describe the micro-foundations of the conceptual model presented in Subsection 5.2. We then show how this simple economic model maps to a statistical model of our reduced-form estimation, as well as provide us with alternative estimation methods.

D.1 Canonical 'Abortion as Insurance' Model (Levine & Staiger 2002)

Levine and Staiger (2002) present the 'abortion as insurance' model as a two-step decision process with incomplete information, as illustrated in Figure D1. In the model, a woman first makes a decision about contraception intensity p_i ($p_i \in [0, 1]$), which implies a conception will happen with probability $1 - p_i$. At this point, the woman has incomplete information about the conditions in which the birth will occur (i.e., whether the father wants the pregnancy or whether her family will support her). She assigns a probability r for receiving a good signal and with probability $1 - r$ a bad signal. Once conception is realized, the information is revealed, and the woman chooses whether to terminate ($q_i = 1$) or give birth ($q_i = 0$).

The theory predicts that incomplete information creates moral hazard in response to increases in abortion access (including a reduction in the cost), resulting in an increase in both abortions and births. A reduction in the abortion cost (i.e. monetary, physical, or psychological) will translate, by backward induction, to a lower contraception decision, resulting in more conceptions. A share of those added conceptions will receive positive information, resulting in more births; and a share will receive negative information, resulting in more abortions. The increase in abortions and births mechanically implies an increase in total conceptions.

Figure D1: 'Abortion as Insurance' Model (Levine & Staiger (2002))



Notes: This Figure illustrates the 'abortion as insurance' model defined by Levine and Staiger (2002), in which the abortion decision is set up as a two-step decision process with incomplete information. See full description above.

D.2 Updating the Model

We modify the Levine and Staiger (2002)'s 'abortion as insurance' model to incorporate credit constraints, as illustrated in Figure 3. Given the lack of evidence of the moral hazard channel presented in Subsection 5.1 we collapse the incomplete information and the abortion decision to one decision node. Then, we add another layer for the model capturing the combination of credit constraints and social stigma cost (which we refer to socially and financially constrained). This feature of the model captures the fact that credit-constrained young women from traditional or conservative households cannot rely on their social network as a safety net for the monetary cost of the abortion.

In the model, a woman first makes a decision on contraception intensity p_i ($p_i \in [0,1]$), which implies a conception will happen with probability $1 - p_i$ (decision I). At this point (decision II), the woman chooses whether to terminate ($q_i = 1$) or give birth ($q_i = 0$). Finally, conditional on wanting to terminate a pregnancy, a woman is either credit constrained ($z_i = 0$) or unconstrained ($z_i = 1$).[13] Below we setup each decision in the tree more formally.

D.2.1 Decision I: Contraception Intensity

Let V_i^j be the expected preferences of woman i over outcome $j \in \{noBirth, birth, abort\}$, and $C(p_i)$ be the contraception cost.[14] A woman i chooses p_i to maximize her expected utility:

$$\max_{p_i} U_i(p_i) - C(p_i)$$

Where:

$$\begin{aligned} U_i = &(1 - p_i) \cdot V_i^{noBirth} \\ &+ p_i \cdot q_i \cdot z_i \cdot (V_i^{noBirth} - V_i^{abort}) \\ &+ p_i \cdot q_i \cdot (1 - z_i) \cdot V_i^{birth} \\ &+ p_i \cdot (1 - q_i) \cdot V_i^{birth} \end{aligned}$$

D.2.2 Decision II: Abortion Decision

Let π^{abort} be the monetary cost of abortion, y_i woman i's disposable income, and $w_i^{abort} = \mathbb{1}\{$ Woman i has financial support for an abortion (family or government)$\}$. Following an unplanned conception, a woman choose to terminate her pregnancy if: $q_i \cdot z_i = 1$. Where:

$$\begin{aligned} q_i \equiv &\mathbb{1}\{V_i^{noBirth} - V_i^{abort} \geq V_i^{birth}\} \\ z_i \equiv &\mathbb{1}\{\pi^{abort} \leq y_i + w_i^{abort} \cdot \pi^{abort}\} \end{aligned}$$

Note that V_i^{abort} is the disutility from doing the abortion (mental, physical, social, and monetary), in absolute values.

[13] It is possible to switch the order of the tree (e.g., the woman first knows her type). However, the model's final derivations are agnostic regarding the timing of this decision tree.

[14] Incomplete information on V_i^{birth} at contraception time is the explanation for the (potential) moral hazard.

E Dynamic Treatment Effects: Event-study Relative To (Potential) Birth

Our detailed panel data on fertility, employment, and education allows us to go beyond average effects post-conception and examine the temporal dynamics of these effects, which can reveal a more nuanced story. For example, a dynamic approach allows us to see how long women delayed parenthood after having an abortion.

To do examine the temporal dynamics, we leverage the 2014 policy variation interacted with an event-study relative to timing of (potential) birth. This approach is an extension of the more standard event-study relative to potential birth that has become common in the "child-penalty" literature (Kleven et al., 2019b).

E.1 Empirical Strategy

In order to better understand the interaction we implement between the difference-in-difference and the event study, we begin by briefly presenting the standard event study design relative to the timing of potential birth known as the 'child-penalty' (Kleven et al., 2019b) and then explain our extension. Given a yearly panel for all women who conceived within the sample period, we denote the year in which she has her first child by c_i and index as $j = t - c_i$ for $j \in \{-3, 3\}$ years relative to the conception. This standard event study is estimated with the following regression:

$$y_{it} = \sum_{j \neq -1} \alpha_j \cdot \mathbb{1}\{j = t - c_i\} + \gamma_{a_{it}} + \gamma_t + \epsilon_{it} \tag{5}$$

where y_{it} denotes an outcome for woman i in year t and at event $j = t - c_i$ and $\mathbb{1}\{j = t - c_i\}$ represent the full set of event time dummies, $\gamma_{a_{it}}$ are age dummies, and γ_t are year dummies.[15] Event time dummy at $j = -1$ is omitted as the reference category, implying that the event time coefficients measure the impact of children relative to the year just before the first (potential) childbirth. By including a full set of age dummies, Equation 5 controls nonparametrically for underlying age-specific effects, and by including a full set of year dummies, we control nonparametrically for time trends, such as wage inflation and business cycles. It is possible to identify the effects of all three sets of dummies because, conditional on age and year, there is variation in event time which is driven by variation in the age at which individuals have their first child.

Next, we incorporate the policy variation (Equation 1) into this event study (Equation 5) Figure E1 illustrates the intuition behind the identification of this combined event study and difference-in-difference approach. Panels E1a and E1b show the event study of a given outcome y (e.g., employment) relative to the timing of a potential birth for a population of women who conceived (pooling abortion and births), split by conceptions that happened before the policy change (blue) and conceptions that happened after the policy change (orange). Panel E1a illustrates the event study for our treated group (age ≥ 20), while Panel E1b illustrates the corresponding event study for our control group (age < 20). The differences between the lines

[15] Unlike the age and time dummies estimated in Sections 4.1 and 6, which were the age at conception and year of conceptions, these are the age and year in which we observe women in the income and education panel.

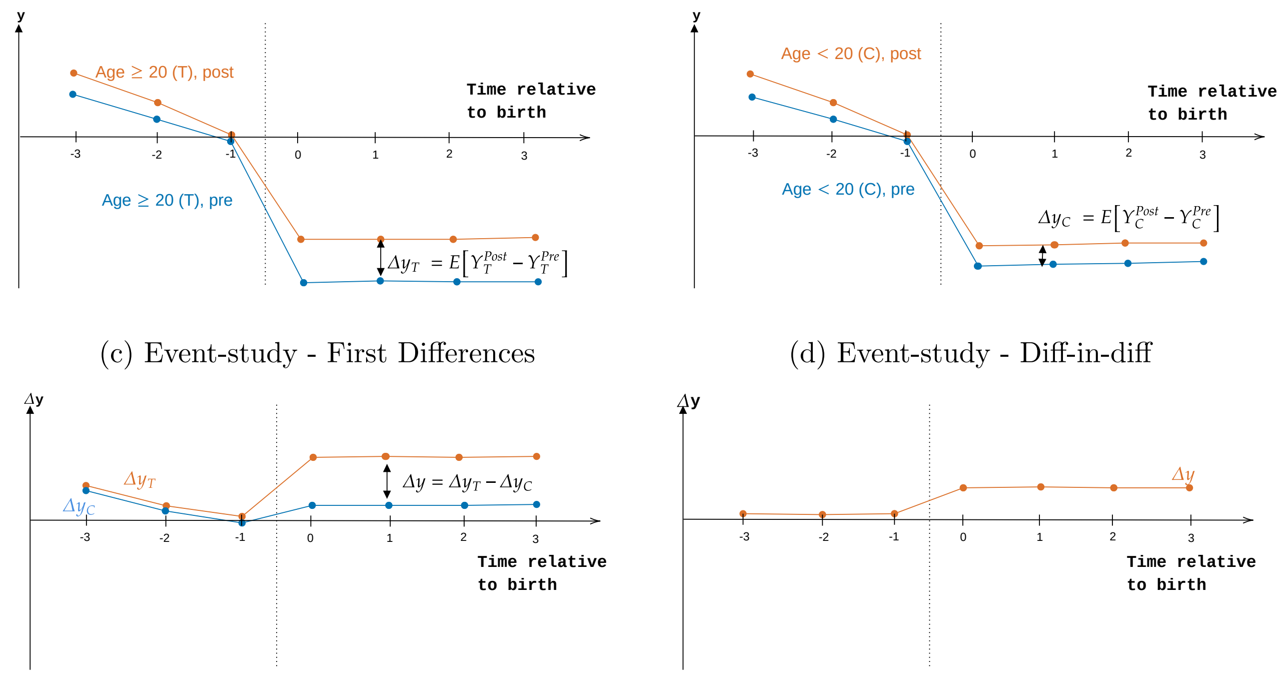
Figure E1: Identification Illustration - Event-study X Difference in Differences

(a) Event-study Treated

(b) Event-study Control

(c) Event-study - First Differences

(d) Event-study - Diff-in-diff



Notes: This figure illustrates the empirical strategy of Equation 6. Panels (a) and (b) show the conceptual event-study for a given outcome y (e.g., employment) relative to timing of potential birth for the population of women who conceived (pooling abortion and births), split by conceptions that happened pre-policy change (blue) and conceptions that happened post policy change (orange). Panel (a) illustrates the event-study for our treated group (age $\geq$ 20), while Panel (b) illustrates the corresponding event-study for our control group (age $<$ 20). The differences between the lines in (a) and (b) are the first differences as illustrated in Panel (c). The difference between the first-differences are illustrated in Panel (d).

in E1a and E1b are the first differences (post − pre) illustrated in Panel E1c. The difference between the first differences by treatment status is the event study difference-in-differences estimator illustrated in Panel E1d.

Figure E1 also illustrates the value of this combined approach. In the absence of any contemporaneous changes, the difference between the graphs in Panel E1a should be sufficient (i.e., the orange line on Panel E1c). However, as Panel E1b illustrates, the outcome could also be changing in the control group over time, perhaps due to a constant shift in attitudes (i.e., time trends). For example, if the outcome is employment and women are increasingly more expected to work post-birth, we will see the shift observed in Panel E1b regardless of the policy change. Therefore, to net out such effects, we take a third difference.

Given that the event study itself serves as a first difference from the dropped year (-1), this is essentially a triple difference estimator. In other words, these coefficients are conceptually equivalent to running separate triple differences for each year since conception relative to year -1. Though, they are not numerically equivalent since we put everything in one regression to get additional power to identify the year and age fixed effects.[16] Formally, we estimate:

[16] The implicit assumption is that age and year effects are fixed across the control and treatment group, before and after the policy change. Furthermore, note that age and year fixed effects are defined in a given calendaric

$$
\begin{aligned}
y_{it} = & \\
& \sum_{j \neq -1} \alpha_j^{Post \times T} \cdot \mathbb{1}\{j = t - c_i\} \cdot Post_{c_i} \cdot T_i \\
& + \sum_{j \neq -1} \alpha_j^{Post} \cdot \mathbb{1}\{j = t - c_i\} \cdot Post_{c_i} \\
& + \sum_{j \neq -1} \alpha_j^{T} \cdot \mathbb{1}\{j = t - c_i\} \cdot T_i \\
& + \sum_{j \neq -1} \alpha_j^{KLS} \cdot \mathbb{1}\{j = t - c_i\} + \gamma_{a_{it}} + \gamma_t \\
& + Post_{c_i} \cdot T_i + Post_{c_i} + T_i + \epsilon_{it}
\end{aligned} \tag{6}
$$

Similarly to Equation 1, *Post* is an indicator that the policy is in effect ($\mathbb{1}\{c_i \geq \text{Dec-2013}\}$) and T_i indicates woman i is eligible for the subsidy ($\mathbb{1}\{20 \leq age_{c_i}\}$). The rest of the terms are defined as in Equation 5.[17]

Here our parameters of interest are the $\alpha_j^{Post \times T}$ from the triple interaction of each event period with the double difference: $\mathbb{1}\{j = t - c_i\} \cdot Post_{c_i} \cdot T_i$. We interpret these estimates as the *additional* effect from avoiding an undesired birth due to the policy among the women who conceived.[18] This triple difference estimator represents the reduced form, or the intention-to-treat (ITT) effect of the policy. Given the policy change happened in 2014 and our tax and education data span until 2018, we take a time span of three years prior to three years post potential birth ($j \in \{-3, 3\}$), relative to one year before potential birth (the dropped category). While this approach allows us to examine interesting temporal dynamics, we note that with only three years of post-policy data, these are relatively short-term impacts.

E.2 Effect on Human Capital Investment and Labor Force Participation

E.2.1 Parenthood and Marriage

Why would access to free abortion affect women's career outcomes? It is well documented that parenthood imposes a 'penalty' on women's careers (e.g., Angelov et al. (2016); Kleven et al. (2019a,b); Eckhoff Andresen and Havnes (2019)). Therefore, we first establish that the increase in abortion due to the policy allowed women to avoid an undesired birth, and thus delay parenthood in the medium-term. For this purpose we define a binary parenthood outcome ($\mathbb{1}\{\text{Is a parent}\}$) and estimate Equation 6 on the sample restricted to socially and financially constrained women (as defined in Section 5.2).

Figure E2a shows a persistent decrease in the probability of entering into parenthood in the three years following potential birth. More specifically, among socially and financially

year, while $Post_{c_i}$ and T_i are defined relative to *conception* time, making them both separately identified.

[17] Given the yearly structure of the data we cannot include month fixed-effects, however, for robustness we include month-of-conception fixed effects which does not affect our results.

[18] 'Added' in the sense of an additional penalty for being *undesired* on top of the child penalty of a birth in general.

constrained women, the 2014 policy reduced the probability of being a parent by 14 percentage points in the three years following the index pregnancy (relative to a baseline share of 43% women who conceived in the year of potential birth). In Section 6 we run the IV equivalent of this strategy. We find that 84% of the compliers did not enter parenthood in the subsequent four years following the conception they aborted.

Avoiding an undesired birth might also reduce undesired marriage, because women who have an undesired pregnancy may end up marrying the father. The data seems to support that claim: Figure E2b presents a persistent decrease in the probability of getting married in the three years following potential birth. More specifically, Table 4 shows that among the population of socially and financially constrained unmarried 18-21 year-olds, the 2014 policy reduced the probability of getting married by 16 percentage points in the years following the index pregnancy (relative to a baseline of 22% in the year of potential birth).

E.2.2 Human Capital Investment & Labor Market Outcomes

We estimate Equation 6 on academic enrollment. We find that an increase in the probability of academic enrollment (Figure E2c). Specifically, the share of women who conceived and enrolled in an academic institution increased by 4 - 11.7 percentage points due to the elimination of the abortion cost (relative to a baseline of 4.4% in the year of potential birth). Note that the (insignificant) effect prior to conception can be attributed to the structure of the education system.[19] Furthermore, note the effect is not a continuation of these trends but a sharp break in trends during the year of potential birth, followed by an increase in enrollment that grows larger over time.

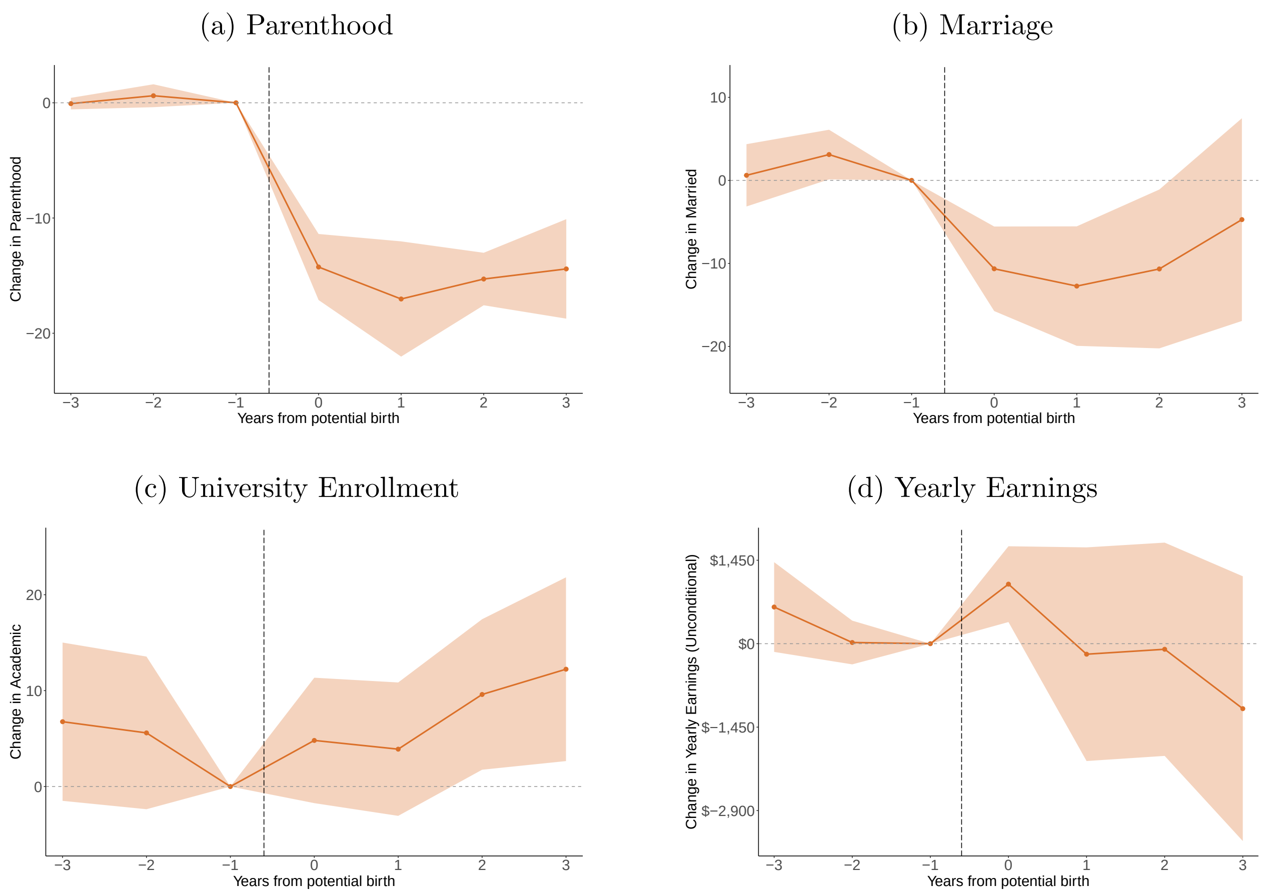
Does this investment in human capital translate into an increase in yearly earnings? Following the same empirical strategy we find a *temporary* increase in yearly earnings unconditional on working (see Figure E2d). Specifically, the yearly earnings (unconditional on working) of socially and financially constrained women increased in the year of potential birth by $1,020 due to the policy (out of a baseline at year of potential birth of $5,806).

This result emphasizes the benefit of the event study estimation relative to the standard reduced form: running the standard (non-dynamic) reduced form shows a negative effect (Table 4), while the event study demonstrates the dynamics: an initial increase in the year of potential birth, followed by decreases in the subsequent years.

Finally, we ask whether this investment in human capital translates into employment in better-paying jobs. To answer this question, we estimate Equation 6 on the sector-level wage-premium. Following Abowd et al. (1999), we estimate the sector-level wage-premiums by running a log-wage regression on individual and sector fixed effects (see further details in Appendix G). The results in Table 4 suggest an increase of 0.08 log-points (or five percentage points) in the wage-premium of the sector in which these women work. Although our three years of data following the 2014 policy somewhat limit our ability to examine the policy's full labor-market

[19] Our treated group (20-21 year-olds)are mechanically more likely to be enrolled three years prior to conception because they are at age (18) when some young people in Israel start higher education. Our control group (18-19 year-olds), who are only 15-16 years old three years prior to conception, are still in high-school. Thus, if academic enrollment for 18 year olds increases over time (because fewer are serving in the military) we should see this effect. These differences are partially residualized because of the inclusion of the age fixed effects. However, in the data our treatment group is never 15 or 16 years old. Consequently, these fixed effects are identified only from the control group. Furthermore, since fixed effects are essentially just demeaning the outcome, there is remaining variation in higher level moments (e.g., variance).

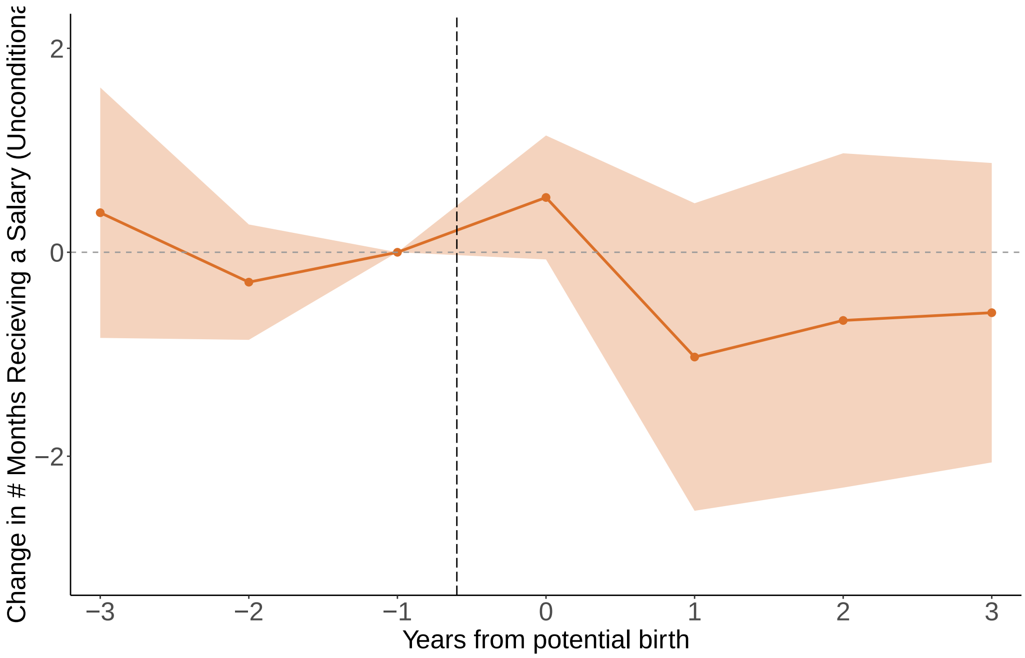
Figure E2: Abortion Access Decreases Entrance to Parenthood and Increased Human-Capital Investment



Notes: This figure presents the event study-DiD results for four outcomes: Panel (a) presents the results for the probability a woman is a parent, Panel (b) presents the results on the probability a woman is married, Panel (c) presents the results for the probability a woman is enrolled in an academic, 4-year university program, and Panel (d) presents the results for the woman's yearly earnings, unconditional on working. Each panel presents the results for the reduced form effect of the 2014 policy relative to year of potential birth as described in Equation 6. Each orange circle represents the treatment effect for the reduced form estimated, from three years prior until three years post potential birth timing, relative to one year prior to potential birth (the dropped year). The shaded regions mark 95% confidence intervals around each point estimate. The dashed vertical line is at 0, indicating an insignificant result (at the 5% level). The sample consists of unmarried women 18-21 year olds who conceived between 2009-2016, and are socially and financially constrained (religious and low-SES, see Section 5.2).

consequences, the temporal dynamics imply an investment in human capital and a shift towards better paying jobs over the short-to-medium term.

Figure E3: Initial Increase in Months Worked Due to the 2014 Policy



Notes: This figure presents the results for the reduced form effect of the 2014 policy relative to year of potential birth as described in Equation 6 on months worked per year. Each orange circle represents the treatment effect for the reduced form estimated, from three years prior until three years post potential birth timing, relative to one year prior to potential birth (the dropped category). The shaded region mark the 95% confidence interval around the point estimate. The dashed horizontal line is at 0, indicating an insignificant result. The sample consist of unmarried women 18-21 years old who conceived between 2009-2016, and are financially and socially constrained (see Section 5.2).

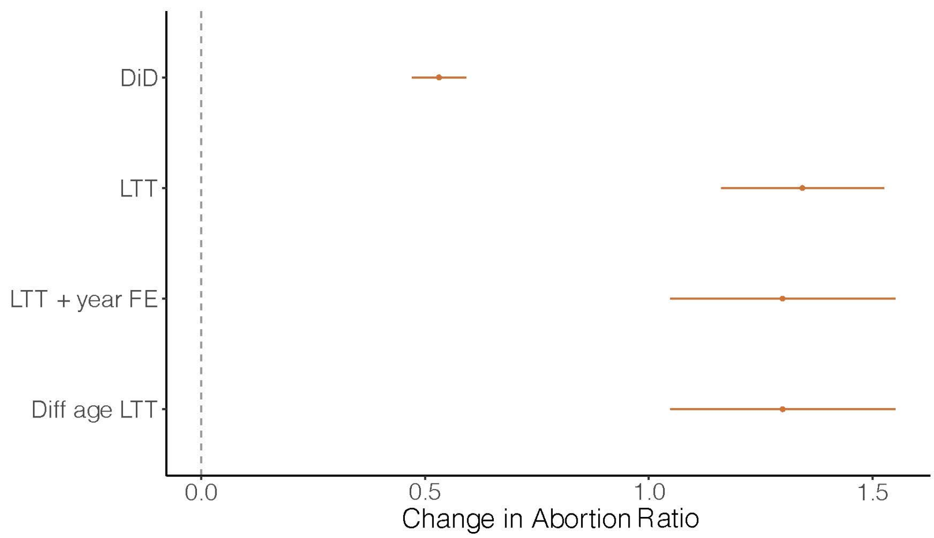
F Hassle Cost

To test for the effect of reduction in the hassle cost of navigating the health insurance system, we exploit another policy change that took place in May 2004, eliminating the need for proof of coverage for women aged 18 or below. Even when women do not have to pay for a health service, such as an abortion, in the Israeli health system they are required to provide the hospital with a payment guarantee from their HMO (health insurer) demonstrating that the procedure is covered. This requires obtaining the payment guarantee form upfront before seeking an abortion or paying out of pocket and going through the HMO to get reimbursed later. Motivated by the concern that the teenagers are unable to navigate the bureaucratic system, and would not want to involve their parents or extended family and social network for support, the Ministry of Health decided in 2004 to eliminate the requirement to obtain a payment guarantee for an abortion for women aged 18 and under.

We test for the effect of the policy and find an increase of 0.8-1.2 p.p in the abortion ratio due to the reduction in hassle cost. More specifically, we exploit the May 2004 policy change and comparing the effect of women aged 18-19 who conceived around the cutoff (18 years old). We restrict to conceptions between 2002-2008 (before the next change of policy that added 19 years old women to the subsidy). Figure F1 presents the results. These results suggest that for very young women who may find it difficult to navigate the bureaucratic and administrative complexities of the health care system necessary to have their abortion paid for by the state, eliminating the requirement to get a payment guarantee has pronounced effects. Note that this exact reduction in hassle cost did not take place on the 2014 policy, but removing the need to come up with the funding can be a reduction in hassle cost by itself.[20] We use the 2004 policy as some evidence for pronounced effects of a small reduction in hassle cost.

[20] Even if a woman doesn't have the money in her bank account but have a source of funding, still figuring out how to raise the money could be seen as a hassle cost.

Figure F1: Reduction in Hassle Cost Increased the Abortion Ratio



Notes: This figure presents difference-in-difference results for the effect of the May 2004 policy that eliminated the requirement to have a payment guarantee from the *kupa* for women below 19 on abortion ration. We exploit the May 2004 policy change and compare the effect of women aged 18-19 who conceived around the cutoff (18 years old). We restrict to conceptions between 2002-2008 (before the next change of policy that added 19 years old women to the subsidy). Each dot represents the treatment effect ($\delta \cdot Post_t \times T_i$) and the lines mark the 95% confidence interval around the point estimate. The dashed vertical line is at 0, indicating an insignificant result.

G Estimating Wage Premiums (AKM)

In this section we explain our estimation of the sectors' wage premiums in the spirit of Abowd et al. (1999) (AKM). To do so we use firm-level estimates on tax data spanning the entire economy (men and women) we have from a different project (Dobbin and Zohar, 2020). We then take these estimates in the firm level and collapse them to the sector level to match to the sector in which each woman is working pre- and post-potential birth.

To estimate the firm's wage premium we assume earnings follow a log-linear functional form:

$$w_{i,t} = \alpha_i + \psi_{J(i,t)} + \mathbf{X}_{i,t} \cdot \boldsymbol{\beta} + \epsilon^y_{i,t} \ , \tag{7}$$

where $w_{i,t}$ is log-earnings of individual i at time t, $\alpha_{i,t}$ is the *individual component of earnings*, $\psi_{J(i,t)}$ the *firm component of earnings*, $\mathbf{X}_{i,t}$ a set of time-varying covariates and $\epsilon^y_{i,t}$ an error term.

Next, we show that the data fits this framework well. In this specification, the *individual component* captures the earnings dispersion within the firm, while the *firm component* (firm earnings premium) captures the earnings dispersion across firms. For equation 7, we have:

$$\begin{aligned} w_{i,t} &= \alpha_i + \psi_{J[i,t]} + \mathbf{X}_{i,t} \cdot \boldsymbol{\beta} + \epsilon^y_{i,t} \ , \\ w_{i,t+1} &= \alpha_i + \psi_{J[i,t+1]} + \mathbf{X}_{i,t+1} \cdot \boldsymbol{\beta} + \epsilon^y_{i,t+1} \ . \end{aligned}$$

Taking first differences:

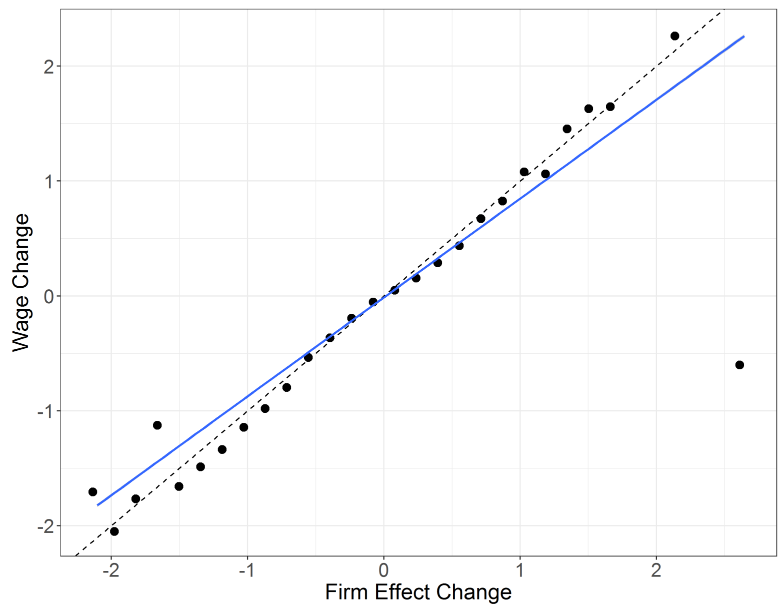
$$\Delta w_{i,t} - \Delta \mathbf{X}_{i,t} \cdot \boldsymbol{\beta} = \Delta \psi_{J[i,t]} + \Delta \epsilon^y_{i,t} \ .$$

In expectation:

$$\mathbb{E} \left[\Delta w_{i,t} - \Delta \mathbf{X}_{i,t} \cdot \boldsymbol{\beta} \right] = \mathbb{E} \left[\Delta \psi_{J[i,t]} \right] \ .$$

We take this restriction to the data by focusing on job switchers and comparing their earnings change against their firm-effect change. The results are in Figure G1. The solid blue line plots the best-fitting line estimated based on the micro-data. The dashed line plots the 45 degree line. Indeed, a percentage change in earnings is corresponds to a percentage change in wage premiums, as implied by the log-linear structure we use.

Figure G1: Earnings Change Corresponds to Firm Fixed Effect Change



Notes: This figure show how the magnitude of earnings changes relate to the change in firm-level pay for workers who switch annual dominant jobs. The earnings are the residualized annualized earnings in the last year at the previous job and in the first year at the new job (employer-to-employer transitions). We sort the job changers into 30 bins on the basis of the change in the firm effects. The circles plot the bin means. The solid line plots the best-fitting line estimated based on the micro-data. The dashed red line plots the 45 degree line. Data source: Israel's Social Security

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