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Not the right time for children: Unemployment, fertility, and abortion

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ABSTRACT

I analyze the effect of local unemployment rates on fertility rates, abortion rates, and the abortions to pregnancies ratio, combining population statistics and administrative data on induced abortions performed in Italy between 2004 and 2016. Using a shift-share instrument measuring labor demand, I exploit demand-driven shocks to unemployment. The findings indicate that both childbearing and abortions behave pro-cyclically, thus suggesting that changes in fertility rates arise not just from changes in planned pregnancy but also from a higher incidence of abortions. These effects are driven by women above 25 years old, and are particularly large in the 35–49 age group, while younger women are largely unaffected.

1. Introduction

Economic recessions impact individual fertility choices due to lower income and economic uncertainty. Bleak economic prospects can influence contraception effort, timing and number of children, and abortion decisions. Understanding the role of economic factors in fertility is crucial for planning healthcare spending and family policies, labor supply. Previous studies have suggested that in high-income countries childbearing tends to decrease as economic conditions worsen (Sobotka et al., 2018). The same attention however has not been dedicated to abortion behavior, where the existing evidence is mostly correlational. This study addresses this gap, documenting the causal relationship between abortions and local economic conditions, proxied by local unemployment rates. Moreover, it extends our understanding of the role of economic factors by considering both abortions and births in the same context. This is crucial because changes in childbearing intentions could be completely absorbed by contraception effort, thus affecting the selection into pregnancy while leaving abortions unchanged.

I use administrative, population-level data on abortions and province-level vital statistics to obtain a panel of fertility and abortion rates for Italian provinces between 2004 and 2016. I then exploit the variation in unemployment rates that occurred during the Great Recession and the Sovereign Debt Crisis to investigate the relationship between unemployment rates, induced abortions, and childbearing. To recover the causal effect of local economic conditions on fertility and abortion outcomes, I employ a shift-share instrumental variable strategy. The findings indicate that both childbearing and abortions behave pro-cyclically, thus suggesting that changes in fertility rates arise not just from changes in planned pregnancy but also from a higher incidence of abortions.

Italy provides a setting where public health insurance covers induced abortion and legislation is homogeneous across regions. This limits potential confounders compared to countries with substantial legislative heterogeneity like the United States. In this context, the monetary cost of an abortion is limited to travel costs, thus allowing to abstract away from the effect of changes in the relative price of abortions during economic downturns. Nonetheless, women still bear the social and personal cost of an abortion and face obstacles in accessing abortion services. To the extent that they are time invariant, these non-monetary costs can be accounted for by exploiting within-area variation.

I employ local unemployment rates as a proxy for both job loss and societal economic uncertainty, for instance through the perceived risk of unemployment. However, unemployment rates may also reflect changes in labor supply induced by fertility choices, perhaps related to unobserved changes in preferences (Hotz et al., 1997). I therefore construct a shift-share instrument in the spirit of Bartik (1991) and similar studies on the relationship between unemployment rates and fertility (Schaller, 2016; Aksoy, 2016). The instrument measures predicted local employment that is unrelated to changes in local labor supply, drawing geographical variation from the predetermined industry specialization and time variation from national employment. For identification, it leverages the employment shocks induced by international recessions, which vary by industry. The identification assumption then is that national employment shocks are independent from relevant province-by-industry unobservables, following Borusyak et al. (2022).

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Standard models of fertility suggest that the effect of a rise in the unemployment rate is ex-ante ambiguous. The negative income effect decreases desired fertility, while unemployment lowers the opportunity cost of childbearing and might thus increase desired fertility through a substitution effect (Becker, 1991). This change in childbearing intentions will affect contraception behavior, subject to the budget constraint; pregnant women will then choose between childbirth or abortion.[1] Therefore, the unemployment rate might affect the abortion rate through responses in both contraception and the propensity to abort conditional on being pregnant. Unemployment shocks can also indirectly affect abortion demand through their impact on other factors, including divorce rates (Schaller, 2013; González-Val and Marcén, 2018) or domestic violence (Anderberg et al., 2016; Bhalotra et al., 2021; Tur-Prats, 2019). Finally, individual characteristics such as age, socioeconomic status, and job characteristics can determine different opportunity costs of childbearing and different exposure to unemployment shocks.

I find that a one standard deviation increase in the unemployment rate decreases the fertility rate by 0.9 standard deviations and increases the abortion rate by 0.25 standard deviations; the propensity to abort conditional on being pregnant also increases by 0.37 standard deviations. This implies that a typical economic downturn, characterized by a 5 percentage point rise in the unemployment rate, translates into approximately 3 fewer births and 0.4 more abortions per 1000 women; conditioning only on pregnant women, the effect rises to 12 more abortions per 1000 pregnant women. Consequently, the overall impact of changes in abortion rates explains approximately one-eighth of the observed fluctuations in fertility rates.

Aggregate fertility responses potentially mask heterogeneous behavior across groups. The response to changes in unemployment rates is likely heterogeneous across age groups because women are at different points of their childbearing cycle and professional careers. Using age-specific instruments, I find that aggregate effects are driven by women above 25 years old and are particularly large in the 35–49 age group, while younger women are largely unaffected. Such differences in response are potentially mediated by a number of factors, including heterogeneity in childbearing intentions, parity, civil status, and career stage. Furthermore, I also explore the heterogeneity of responses across different geographical areas, where women are exposed to different labor market characteristics, social norms, and access to abortions. The results suggest that the change in national rates is driven by provinces in the Central and Northern regions, with some evidence of local labor market characteristics such as job irregularity playing a role.

This study contributes to the literature investigating the fertility response to economic shocks on the one hand, and the literature regarding the determinants of abortion demand on the other. Previous research has investigated the relationship between childbearing and economic fluctuations, both at the aggregate and individual level; Sobotka et al. (2018) reviews the literature concerning high-income countries. Although the evidence is not uniform, the majority of studies support a pro-cyclical relationship of births to economic fluctuations, though with relatively small and short-lived effects. Goldstein et al. (2013), Schneider (2015), Comolli (2017), and Matysiak et al. (2021) show that fertility decreases with higher unemployment rates. Schaller (2016) finds that a one percentage point increase in local unemployment induces between a 1.5 and 2.2 percent decrease in the fertility rate; Del Bono et al. (2012) estimate that female job loss due to plant closure reduces the number of children born by 5 to 10 percent in the short and medium-term. Previous literature has therefore suggested that exposure to higher local unemployment rates has a smaller impact on childbearing than individual unemployment status, also taking into consideration that plant closures induce a potentially prolonged decrease in income. The literature has also documented an effect on childbearing intentions. In the context of Italy, having a temporary job contract or being unemployed is associated with a reduction in childbearing intentions by between 15 and 10 percentage points (Modena and Sabatini, 2012; Modena et al., 2014); more generally in Europe, societal economic uncertainty is negatively correlated with short-term parenthood intentions (Fahlén and Oláh, 2018). The present study brings forward additional evidence of the pro-cyclical behavior of childbearing, where the size of the estimates is consistent with previous results by Schaller (2016). In addition, I consider the response of both births and abortions to the same shock. This approach recognizes that the analysis of abortion patterns cannot be separated from that of fertility, since the number of abortions ultimately depends on the number of pregnancies. Due to data limitations, this study does not speak to the relative effect of female or male unemployment shocks on fertility choices which are discussed by Huttunen and Kellokumpu (2016) and Schaller (2016).

Studies of abortion demand have mostly concentrated on changes in the cost of abortion, particularly in the United States, for instance due to legislative restrictions (Haas-Wilson, 1996; Bitler and Zavodny, 2001; Medoff, 2007; Myers and Ladd, 2020), abortion clinic closures (Fischer et al., 2018; Lindo et al., 2020), insurance coverage (Levine et al., 1996), or the diffusion of contraception methods (Ananat and Hungerman, 2012). Some research has studied the role of social welfare policies, such as income support (Snarr and Edwards, 2009) or child support enforcement (Crowley et al., 2012). González and Trommlerová (2023) find that a universal child benefit in Spain increased the birth rate both through an increase in conceptions and a decrease in abortions. Bárdits et al. (2023) investigate the role of dismissal protection in fertility responses to employment shocks, showing births are timed to protect women from mass layoffs, while abortions increase if the firm closes and protection is lost. Specifically, the role of unemployment has been largely overlooked: although it is included as one of many potential determinants of abortion in the classic models of Medoff (1997) and Blank et al. (1996), these do not account for its potential endogeneity with respect to fertility choices. Lima et al. (2016) find that the abortion ratios in 2010–2012 exceeded the predicted trend across several European countries, suggesting the economic recession and austerity policies as potential determinants. However, their study only relies on temporal variation, and the authors cannot make a causal statement nor address the potential heterogeneous response across countries that experienced the recession with different intensity and timing. In this paper, I overcome these issues and fill the gap left by previous research by taking a causal approach and focusing on a single country, Italy.

The remainder of the paper is organized as follows: Section 2 describes the institutional framework of abortions in Italy; Section 3 presents the data; Section 4 illustrates the empirical strategy. Section 5 presents the main results, while Section 6 covers robustness checks. Section 7 concludes.

2. Institutional framework

The Italian National Health Service is based on the principle of free universal coverage and follows a decentralized model where regions are responsible for the organization and provision of care. Abortion in Italy is regulated by Law 194 of 1978 and since then the Italian National Institute of Health (Istituto Superiore di Sanità) has maintained a surveillance system for legally induced abortions, based on quarterly reporting by the regional health authorities.

According to this regulation, all women are eligible to request the voluntary interruption of a pregnancy during the first 90 days of gestation. Beyond this 90 days limit, only therapeutic abortions are

[1] While standard economic models often assume that agents have perfect control over their fertility, here I account for the stochastic nature of fertility, i.e. the possibility of unintended pregnancies. Therefore, women are faced with three subsequent choices: their childbearing intentions, contraception effort, and abortion or childbearing choice.

permitted, i.e. abortions motivated by medical concerns. A woman seeking an abortion must first obtain a certificate attesting to the pregnancy from either her general practitioner, a private physician, or a public family clinic; parental or judicial consent is required for minors. With the exception of urgent cases, there is a mandatory seven-day period of reflection after the certificate. Induced abortions can be performed either in public hospitals, free of charge, or in authorized private clinics; more than 90% take place in public hospitals (Ministero della Salute, 2016).

Although the interruption of a pregnancy for elective reasons is legally allowed, access to abortion services is not always straightforward. Law 194/78 also regulates the practice of conscientious objection, granting healthcare personnel the right to refuse to participate in procedures aimed at terminating a pregnancy, except in cases where they are deemed necessary to save the woman's life. (Autorino et al., 2020) show that a higher prevalence of objecting professionals is associated with a higher share of women having an abortion outside the region and longer waiting times. In 2016, 71% of gynecologists were objectors,[2] over 8% of all abortions were sought by women out of their region of residence, and around 13% out of their province of residence (Ministero della Salute, 2016). To account for this inter-regional mobility, I consider abortions based on the province of residence of the woman, rather than the province where they occurred.

There is substantial geographical variation in terms of both abortion and childbearing behavior, influenced by both demand and supply factors. Fig. 1 illustrates the variation in childbearing, abortion, and doctor's conscientious objection rates in 2004. A visual inspection suggests a positive correlation between these patterns and political liberal views. Specifically, abortion rates are high in the North-West and North-Center areas and low in the conservative North-East and Sicily, while the opposite applies to fertility rates. This correlation with local views is not a threat to the analysis, as these persistent cultural factors are accounted for by province fixed effects. On the other hand, there are also areas where both abortions and childbearing rates are high, especially in the Central and Southern regions, which might indicate a higher conception rate is associated with a higher proportion of abortions out of pregnancies. Moreover, there are substantial differences in the prevalence of conscientious objection, presented in Panel (c). Higher rates of objection are observed in more conservative areas such as the North-West and Southern regions. The coexistence of high abortion rates and a higher proportion of objecting doctors in these areas suggests that factors beyond cultural and supply considerations contribute to abortion demand. This study focuses on the elasticity of abortion demand to changes in unemployment rates, which might be confounded by constraints on the supply side. To address this potential issue, any systematic time-invariant differences in the supply of abortion services across provinces are captured by province dummies. The robustness analysis further accounts for time-varying elements, such as changes in doctors' objection preferences over time.

Over the years, several regulatory changes have occurred: the legalization of emergency contraception pills in 2000 and 2012[3]; the introduction of medication abortion in 2009; the availability of emergency contraception pills without a medical prescription from 2015. The role of these regulatory changes as potential confounders is addressed in Section 6.

3. Data and descriptive statistics

In this section, I describe the main sources of data used in this paper and sample selection. Moreover, I define the dependent variables and discuss their evolution over time. Further details on the construction of variables are provided in Appendix E.

3.1. Data and construction of the sample

I employ yearly data on Italian provinces between 2004 and 2016 provided by the Italian National Institute of Statistics (ISTAT). The geographical unit of the analysis is a province (NUTS III division), with boundaries fixed to 2004 to keep geography constant.[4] The final sample includes 1339 observations, corresponding to a balanced panel of 103 provinces across the years 2004–2016.

To construct province-level dependent variables, I combine two sources of data: population statistics at the province level and administrative data on induced abortions. Population statistics include data on population and live births from the General Register Office, disaggregated by the province of residence and age of the mother. Data on abortions pertain to administrative data on voluntary interruptions of pregnancies collected by ISTAT.[5] This covers all legal abortions performed in Italy, both in public and private facilities. For each procedure, the medical staff compiles a standardized form (module D.12) with details on the procedure and socio-demographic characteristics of the woman. In particular, this form records the age, citizenship, area of residence, marital status, and reproductive history of the woman. Appendix B.2 discusses measurement error in the abortion data, including covert abortions, and presents several robustness checks.

Local economic conditions are proxied by the provincial unemployment rate, which reflects both the share of unemployed agents and the perceived level of uncertainty in the local labor market. The main variable of interest considered is the total unemployment rate, covering both genders across all working-age individuals, but I also consider age-specific unemployment rates. To address concerns of endogeneity between unemployment rates and fertility and abortion outcomes, I construct an instrument based on supply-side employment data by industry from the regional accounts, as described in Section 4.

The analysis focuses on Italian women of childbearing age, i.e. between 15 and 49 years old. I thus restrict the sample to women of Italian nationality and born in Italy,[6] as women of foreign nationality might react differently to economic factors due to differences in socio-economic composition, cultural values, labor market exposure, and access to health services (Spinelli et al., 2006).[7] Most importantly, different cultural norms of foreigners regarding abortion would not be captured by province indicators, which only proxy for the local social cost of abortion in Italy, thus introducing bias in the estimation. The incidence of abortions among foreign citizens is indeed almost three times higher than among Italian citizens (Ministero della Salute, 2016) and is not comparable to the abortion rate of any Italian region, as highlighted by Fig. A.1. In addition, I consider only interruptions of pregnancies taking place before the statutory 90 days limit since

[2] Additionally, in 2016 49% of anaesthesiologists and 44% of non-medical staff were conscientious objectors.

[3] Emergency contraception pills were legalized in Italy in 2000 if based on Levonorgestrel, and in 2012 if based on Ulipristal acetate.

[4] Seven new provinces were established during the observation window: 4 in 2006 and another 3 in 2010. I absorb each of these new provinces into their parent province; the correspondence is one-to-one except for Olbia Tempio, which I assign to Sassari since it was composed of 24 municipalities from Sassari and only 2 from Nuoro. Therefore, from an initial sample of 110 provinces, I remain with 103 provinces. In addition, in 2009 seven municipalities moved from the Pesaro to the Rimini province; results are robust to dropping these provinces, as reported in Table B.3.

[5] Data analysis was conducted at the Laboratory for Elementary Data Analysis (Laboratorio per l'Analisi dei Dati Elementari) of ISTAT, in compliance with legislation concerning the protection of statistical secrecy and personal data.

[6] Foreign-born women with Italian citizenship constitute around 6% of the sample of Italian women in the abortion data.

[7] In addition, the prevalence of illegal abortion might be higher among foreigners, related to irregular migration status, thus making measurement error systematically different for this group. Pieroni et al. (2023) show how changes in immigrants' legal status following the 2007 EU enlargement affected voluntary abortions.

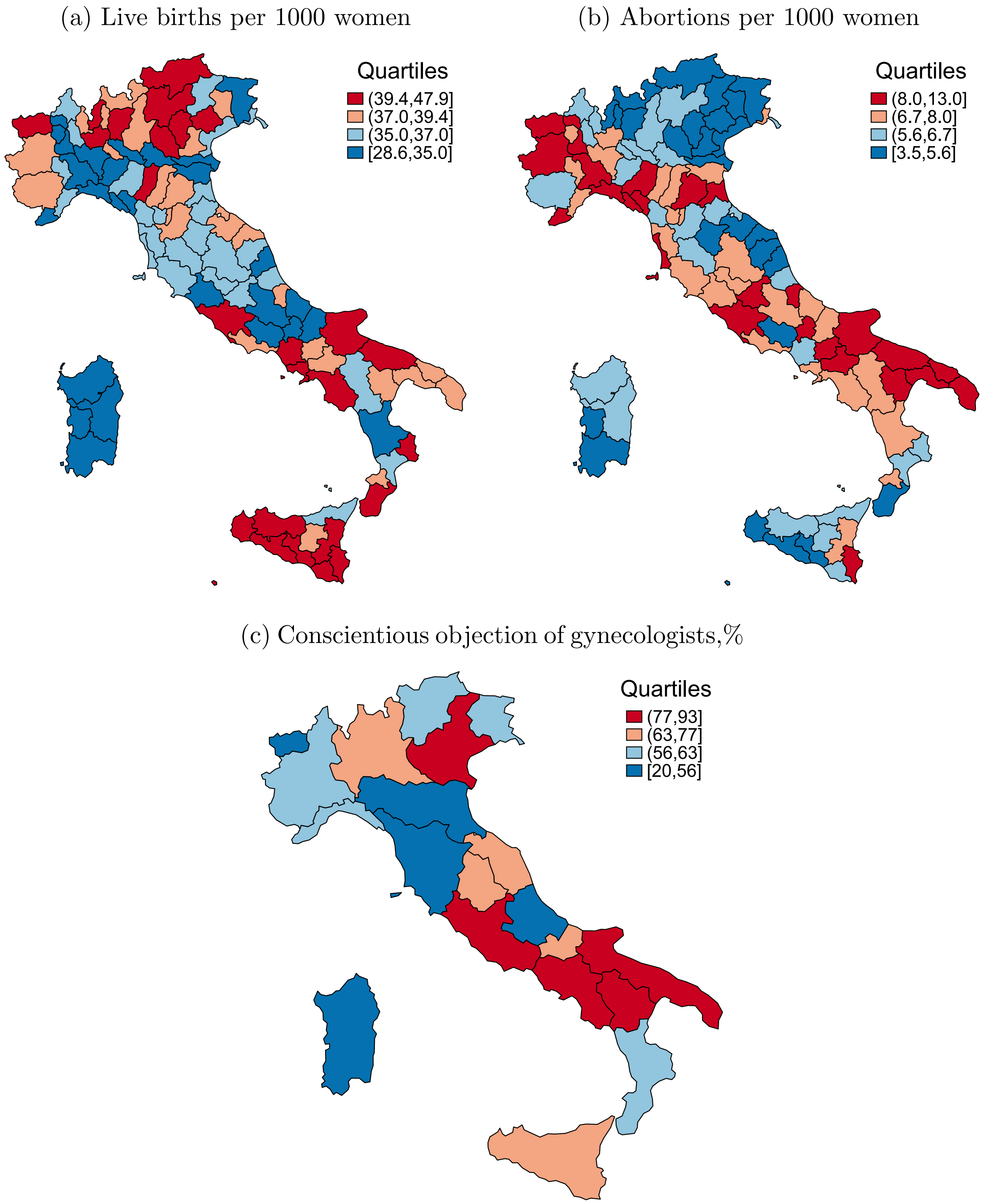

Fig. 1. Fertility and abortion in 2004.

abortions performed after this date should respond primarily to medical concerns.[8]

[8] After 90 days from conception, abortions are allowed only in case of serious risk to the woman's life or severe fetal malformation. However, it is also possible that part of these abortions is related to economic conditions, as previous literature has established an increase in miscarriages (Bruckner et al., 2016) and better infant health outcomes (Dehejia and Lleras-Muney, 2004) in times of high unemployment rates. What effect economic fluctuations might have on fetal malformations and mothers' medical state remains an open question for future research, out of the scope of this analysis.

3.2. Dependent variables

This study considers three dependent variables: the general fertility rate, the abortion rate, and the abortion ratio. The general fertility rate (henceforth GFR) measures the average number of births in a year for every 1000 women who are in their childbearing years, i.e. between 15 and 49 years old. Similarly, the abortion rate indicates the incidence of abortions in the population of reproductive-aged women, i.e. the number of abortions per 1000 women in their childbearing years by the province of residence. Focusing on the province of residence rather than the province of abortion establishes a closer connection between local economic conditions and abortion choice, on top of accounting for cross-province migration to access abortion services (see Section 2). A rise in the abortion rate can be the result of both a higher pregnancy rate (as a result of reduced contraception, for instance) or a larger share of unwanted pregnancies, keeping the pregnancy rate constant. For this reason, I also consider the abortion ratio, which measures the propensity to abort conditional on being pregnant, thus capturing also changes in the pregnancy rate. This ratio is computed as the share of abortions over pregnancies, where the number of pregnancies is proxied by the sum of live births and abortions.[9] As a result, the abortion ratio is increasing in the abortion rate and decreasing in the general fertility rate, as follows:

$$Ab.ratio = \frac{Abortions}{Pregnancies_{000}} \approx \frac{Ab}{Ab + Livebirths} = \frac{Ab.rate}{Ab.rate + GFR} \tag{1}$$

These aggregate measures are affected by the age structure of the population; I therefore replicate the analysis using age-specific rates. For example, I construct group-specific fertility rates by dividing the number of births by the appropriate at-risk population, i.e. women aged 15–49 in the relevant demographic group.

3.3. Descriptive statistics

The evolution of the main variables of interest at the national level is presented in Fig. 2. The general fertility rate, reported in Panel 2(a), is relatively stable in the first half of the sample and reaches its peak in 2008, at around 39 live births per 1000 women of childbearing age. It then starts to rapidly decrease, down to 34 live births per 1000 women of childbearing age in 2016. The abortion rate is decreasing throughout the observation period, but at a slower rate from 2009 onwards; by 2016, it stands at around 5 abortions per 1000 women in their childbearing years. Panel 2(b) reports the evolution of the abortion to pregnancies ratio and the unemployment rate. The abortion ratio initially decreases steadily, but it almost flattens out from 2008 as a result of the sudden decrease in childbearing and the slower decrease in abortions. Finally, the national unemployment rate shows substantial variation over time, confirming the years between 2008 and 2014 to be a period of prevailing economic instability and underlining the double-dip nature of the recession. The impact of the recessions under consideration varied across geographical areas in terms of timing, but the total variation in unemployment rates was dispersed across the country.[10]

For summary statistics of the province-level data and the underlying abortion individual-level data, see Table A.1 in Appendix A.

4. Empirical strategy

The goal of the analysis is to study the relationship between childbearing, abortion, and local economic conditions, proxied by the unemployment rate. To analyze different aspects of fertility choice, I consider multiple dependent variables: the general fertility rate, the abortion rate, and the abortion ratio. When studying fertility rates I lag the unemployment rate by one year to consider the unemployment rate in the year of conception.[11] The baseline specification is therefore the following:

$$Fert_{p,t} = \alpha_0 + \beta_0 Unempl_{p,t-1} + \gamma_{0,p} + \delta_0 year + \epsilon_{0,p,t} \tag{2}$$

$$Ab_{p,t} = \alpha_1 + \beta_1 Unempl_{p,t} + \gamma_{1,p} + \delta_1 year + \epsilon_{1,p,t} \tag{3}$$

$$Ab.ratio_{p,t} = \alpha_2 + \beta_2 Unempl_{p,t} + \gamma_{2,p} + \delta_2 year + \epsilon_{2,p,t} \tag{4}$$

where $Fert_{p,t}$, $Ab_{p,t}$ and $Ab.ratio_{p,t}$ are the outcomes for province p in year t, $Unempl_{p,t}$ is the local unemployment rate, and γ_p are province dummies. Standard errors are clustered at the province level, and province boundaries are fixed to 2004.

This baseline specification is modeled after Schaller (2016) and controls for province fixed effects and a linear time trend, which capture the confounding effect of unobserved time-invariant local characteristics and linear trends over time that are common to all provinces. For instance, province dummies capture social attitudes towards abortion, i.e. the social cost of induced abortion, and time-invariant supply-side constraints in abortion services. Moreover, they also account for time-invariant measurement error in both the dependent and independent variables such as mismatches between the actual and official place of residence; the average cross-province employment mobility; and systematic misreporting of births and induced abortions. The linear time trend removes any aggregate variation given by national trends, therefore it accounts for agents' linear expectations of the evolution of the economy as well as the historically decreasing trend in fertility outcomes depicted in Fig. 2.[12]

Using local unemployment rates as a proxy of local economic conditions captures the effect of both individual job displacement and economic uncertainty, as the rate of unemployment correlates with the perceived risk of unemployment (Anderberg et al., 2016), the job separation rate, and it is negatively related to the job-finding rate. Although unemployment rates can understate the magnitude of a recession by not accounting for discouraged workers, they are the best available proxy to capture changes in the labor market conditions at this level of disaggregation. In addition, the unemployment rate is useful in exploring fertility behavior because it is less likely to be endogenous to childbearing or abortion choices than other variables such as own wages.

However, the results of the OLS estimation cannot be interpreted as causal for a number of reasons. First, there is an issue of reverse causality: an increase in fertility (abortions) can induce more women to drop out (stay in) the labor force (Bloom et al., 2009; Kalist, 2004). Unemployment rates are therefore correlated with fertility-induced changes in the labor supply. Moreover, recent literature shows that fertility indicators lead economic recessions by several quarters (Buckles et al., 2021). Second, there are omitted variables that affect both fertility and unemployment, such as the age composition of the labor force or unobserved preferences (Hotz et al., 1997). A simultaneous increase in childbearing intentions and a decrease in female labor force participation will increase the fertility rate while decreasing the denominator of the unemployment rate, thus leading to upward biased OLS estimates. Conversely, a change in preferences that induces both a decrease in abortions and labor force participation will lead to downward biased OLS estimates.

To address these concerns of endogeneity, I construct an instrument for the local unemployment rate following the approach developed by Bartik (1991) and employed by Bound and Holzer (2000), Autor and Duggan (2003), Schaller (2016). The instrument averages national employment across industries, using local industry employment shares as weights, to produce a measure of predicted local employment that is unrelated to changes in local labor supply. More specifically, this variable interacts predetermined industrial composition within a province with national time-varying industry employment levels, as follows:

$$B_{p,t} = \sum_{k=1}^{K} \chi_{p,k,t_0} E_{-p,k,t} \tag{5}$$

[9] I define the abortion ratio as the number of abortions per 1,000 pregnancies, in line with the definition by the Guttmacher Institute, while the CDC defines it as the number of abortions per 1,000 live births.

[10] This is further discussed in Appendix D.

[11] By definition, three-quarters of the conceptions in year t will be realized in year $t+1$ due to the 9 months gestation period. Since monthly data on live births is not available, I approximate the time of conception with the year preceding the birth. Results are robust to considering the contemporaneous unemployment rate.

[12] A sensitivity analysis of the results to alternative time specifications is reported in Section 6, including province-specific time trends and a quadratic time polynomial.

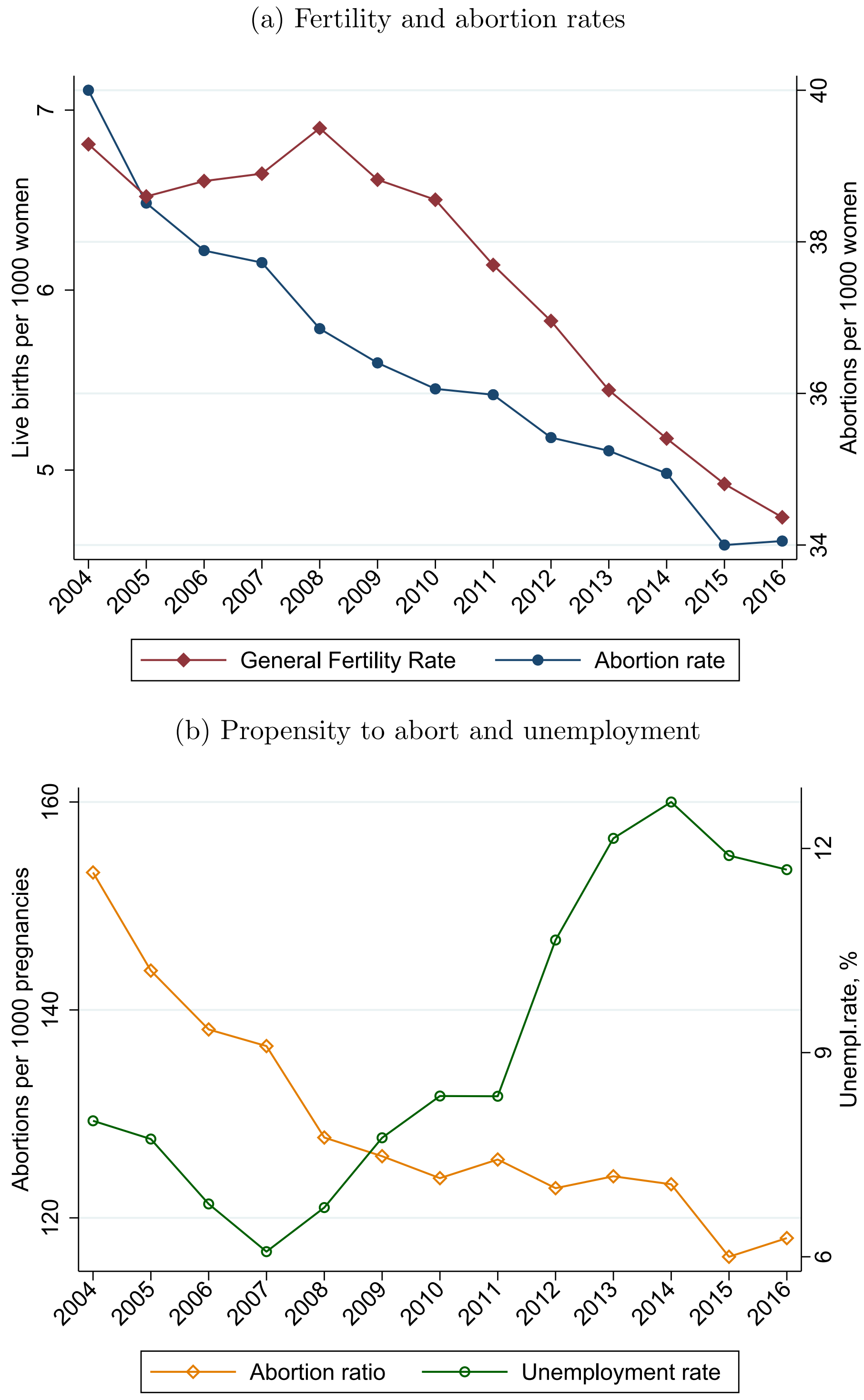

Fig. 2. Main variables.

where χ_{p,k,t_0} is the employment share of sector k in province p and base period t_0, and $E_{-p,k,t}$ is number of people employed in sector k, period t in Italy, excluding province p.

The instrument measures the predicted employment level for each province. It captures variation driven by changes in the national economy but differing across provinces due to predetermined differences in industry employment distribution. Therefore, if specific industries are hit by a shock, this will reflect particularly on provinces that specialize in those industries. To address finite sample bias coming from using own-observation data, I calculate each industry's national employment excluding own-province employment. Fig. 3 illustrates the first stage relationship graphically, showing a strong negative correlation between the observed unemployment rate and the predicted employment level.[13]

Therefore, the instrument identifies demand-driven changes in the unemployment rate: by keeping industry composition constant over time and exploiting temporal variation originating only from the national employment level, it abstracts from changes in local labor supply

[13] Appendix C.2 of the Appendix further reports the point estimates for the first stage relationship and the relative Kleinebergen-Paap F-statistic, which rejects the null hypothesis of under-identification.

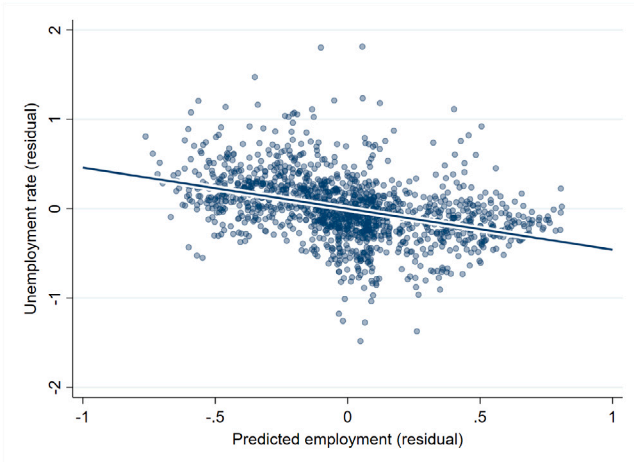

Fig. 3. First stage relationship.
Notes: The figure plots the relationship between residualized unemployment rates and predicted employment as measured by the shift-share instrument, once removed the linear time trend and province fixed effects.

and sorting into industries. Recent methodological literature has underlined that identifying variation in shift-share designs can stem from either the national shifts or the local shares (Jaeger et al., 2018; Adao et al., 2019; Goldsmith-Pinkham et al., 2020; Borusyak et al., 2022). In the present setting, the local industrial composition is likely to be endogenous, because local industrial composition determines opportunities for female employment and thus correlates with the unobserved labor supply changes and childbearing preferences. In addition, other unobserved shocks, such as the expansion of female-dominated industries, are likely to affect outcomes through the same mixture of exposure shares. I thus follow the interpretation suggested by Borusyak et al. (2022), where identification relies on the exogenous assignment of national industry employment shocks. This translates into assuming that national employment shocks are independent from relevant province-by-industry unobservables. Fig. 4 illustrates the variation of employment levels over time across the industries, where the grayed areas highlight the years of recession. Employment in public administration remains quite stable over time. The Great Recession induces a contraction in employment, particularly in the manufacturing, real estate, and financial sectors. In 2013, we observe another contraction that affects most industries, but especially the construction sector.

As a robustness check, I construct alternative versions of the instrument, by changing either the shares, the reference year employed for the shares, or the shift variable. In addition, I also employ a version of the instrument based on data from the Labor Force Surveys, allowing both to expand the number of industries considered and to refine some measurement error in industry employment. The construction of these alternative measures and the corresponding regression estimates are discussed in Appendix C.3.

In addition, I replicate the analysis focusing on age-specific fertility and abortion outcomes and age-specific instruments. The general fertility and abortion rates are affected by the underlying age composition of the population, as childbearing decisions significantly vary over a woman's lifecycle (Hotz and Miller, 1988; Del Boca and Sauer, 2009). Thus, if the female population of a province is predominantly young, we would expect the fertility rate to be higher than in a province with a predominantly old female population. Moreover, changes in the labor market heterogeneously impact different age groups; for instance, younger women might be more inclined than older women to postpone childbearing when facing adverse job prospects. To construct age-specific versions of the instrument, I adjust the employment shifts by the contemporaneous employment share of each age group at the national level, similarly to Schaller (2016): thus, the age-specific instrument rescales the predicted local level of employment according to the national age composition of employment.[14]

5. Results

This section first discusses the overall effect of changes in unemployment rates on fertility and abortion outcomes and later considers the heterogeneity of this response depending on women's age and area of residence, i.e. local labor market characteristics and social values.

Table 1 reports the main results of the paper for the three dependent variables of interest. Columns (1) and (4) analyze the general fertility rate, reporting respectively the OLS and IV estimates. The OLS results indicate that the unemployment rate has a negative and statistically significant effect on fertility rates, where a one standard deviation increase in the unemployment rate reduces the GFR by 0.21 standard deviations (95% CI [−0.31, −0.11]). The IV estimation confirms the direction of the effect and yields larger estimates in magnitude: increasing the unemployment rate by one standard deviation translates into a reduction in the fertility rate by 0.95 standard deviations (95% CI [−1.13, −0.78]). The reported Kleibergen–Paap LM under-identification test and the Kleibergen–Paap F statistic strongly reject the hypothesis of a weak instrument.

Columns (2) and (5) report the results for the abortion rate. The abortion rate increases with the unemployment rate, and again the IV estimates are larger than the OLS ones. The IV estimates indicate that a one standard deviation increase in the unemployment rate brings about a 0.25 standard deviation change in the abortion rate (95% CI [0.12,0.39]).

Finally, columns (3) and (6) present the main results for the abortion to pregnancies ratio. The OLS coefficients indicate that a one standard deviation increase in unemployment is associated with a 0.18 standard deviation change in the propensity to abort conditional on being pregnant (95% CI [0.11, 0.26]). This effect however more than doubles when we move to the IV estimation: a one standard deviation increase in unemployment is now associated with a 0.38 standard deviation change in the abortion ratio (95% CI [0.23,0.52]). Overall, the propensity to abort conditional on being pregnant increases when unemployment rises, as a result of both the reduced number of births and the increased number of abortions.

Coefficient estimates from the IV analysis are larger, in absolute value, than the OLS estimates, consistently with the expected reverse-causality bias discussed in Section 4: lower fertility induces higher female labor supply, and thus lower unemployment rates, generating attenuation. Moreover, measurement error in unemployment rates could also be causing OLS coefficients to be biased downward (in absolute terms). The estimated effects indicate that for a one standard deviation increase in the unemployment rate, the general fertility rate responds with almost a one standard deviation decrease while the abortion rate increases by a quarter of a standard deviation.[15] In percentage terms, a 1 percentage point change in the unemployment rate induces a 1.7% reduction in the fertility rate of the following year. This aligns with findings by Schaller (2016), who estimates a decrease in fertility between 1.6 and 2.2%. In general, individuals experiencing unemployment spells are expected to be more affected due to direct income

[14] Formally, the age-specific instrument corresponds to the following: $B_{a,p,t} = \frac{E_{a,IT,t}}{E_{IT,t}} \sum_{k=1}^{K} \chi_{p,k,t_0} E_{-p,k,t}$.

[15] These results are robust to the use of initial female population weights, or initial births and abortions weights. Weighting for the female population accounts for the distribution of the population of interest, i.e. Italian women of childbearing age, across provinces. Weighting for the initial magnitude of births and abortions accounts for the relevance of each province in contributing to the national fertility outcomes. Such estimates are available from the author upon request.

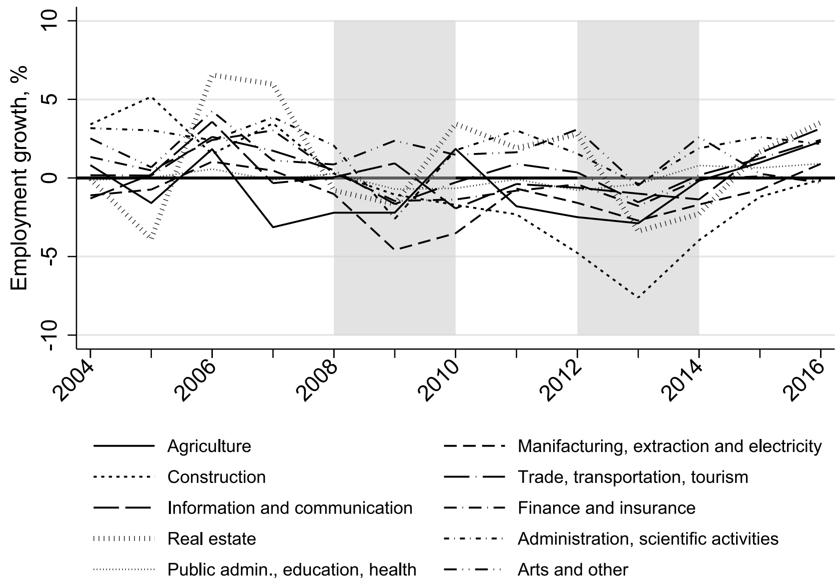

Fig. 4. Employment growth.

Table 1
Main specification - standardized variables.

	OLS			IV		
	GFR (1)	Ab.rate (2)	Ab.ratio (3)	GFR (4)	Ab.rate (5)	Ab.ratio (6)
$Unempl_{t-1}$	−0.212*** (0.052)			−0.956*** (0.089)		
$Unempl_t$		0.144 *** (0.032)	0.187*** (0.038)		0.253*** (0.069)	0.375*** (0.074)
Observations	1236	1339	1339	1236	1339	1339
R^2	0.581	0.561	0.353			
KP LM p-value				0.000	0.000	0.000
KP F-stat				360.5	285.5	285.5

Standard errors in parentheses. All regressions include province fixed effects and a linear time trend.
* $p < .05$, ** $p < .01$, *** $p < .001$.

shocks, while women indirectly impacted by economic uncertainty may respond to a lesser extent. This entails that the estimated effect of the unemployment rate on aggregate fertility outcomes should be smaller than that of individual unemployment on individual fertility. In fact, the effects estimated in this study are relatively limited compared to the effect of job displacement (Del Bono et al., 2012; Huttunen and Kellokumpu, 2016), also taking into consideration that plant closures induce a prolonged decrease in income.

Previous literature has also shown that fertility rates respond heterogeneously to male and female unemployment rates (Aksoy, 2016; Huttunen and Kellokumpu, 2016; Schaller, 2016), both as a proxy of income and substitution effects respectively. Similarly, abortion rates could respond differently to male and female-specific labor market shocks. In the current setting, there is insufficient independent variation between female and male unemployment rates, which tend to co-move as shown in Appendix Fig. D.1. Consequently, estimates separating the effect of changes in female and male unemployment rates will be uninformative, leaving this question open for future research.

5.1. Age heterogeneity

In this section I explore the heterogeneity of response by age, dividing women of childbearing age into three groups: from 15 to 24; from 25 to 34; and from 35 to 49 years old. Different age groups might experience different substitution effects because labor force participation, wage levels, career expectations, and the probability of having a stable partner change with age. Moreover, by focusing on age groups, I can differentiate the responses of women at different points of their childbearing cycle, with varying childbearing intentions, contraceptive use patterns, and number of previous children.

Previous literature on fertility has painted a varied picture: Ananat and Hungerman (2012), Goldstein et al. (2013), Schneider (2015), Comolli (2017) find childbearing of younger groups to be the most responsive to changes in the unemployment rate, an effect driven by first births. Schaller (2016) and Del Bono et al. (2012) instead find older age groups to be more responsive; Comolli (2017) finds women in their late thirties (35–39) to be the second most affected group. On the one hand, young women might be the most responsive, as youth unemployment was most affected by the recession and fertility plans can be revised more easily at younger ages (Goldstein and Cassidy, 2014); on the other hand, they are more likely to use contraception.[16] Older women instead are more likely to be in a stable relationship and less likely to use contraception, even when not intentionally seeking a pregnancy (Loghi and Crialesi, 2017); however, economic recessions might affect both divorce and marriage rates (Schaller, 2013; González-Val and Marcén, 2017). Women between 25 and 34 are in the prime of their childbearing years and professional careers, and therefore

[16] In fact, survey data suggests that, conditional on being sexually active, the share of women using contraception is higher among women younger than 25 (Loghi and Crialesi, 2017). Survey data also indicates that women between 18 and 25 years old are the main users of emergency contraception (Bastianelli et al., 2005, 2016).

most likely to postpone childbearing. Changes in the behavior of this group will be particularly relevant because the incidence of births is highest in this age group (see Table A.1). Women above 35 years old are closest to the end of their reproductive life, but they are also more likely to already have children and face a stronger trade-off between quality and quantity of children.[17] Finally, the recession might have brought about a reduction in births through lower recourse to assisted reproductive treatments such as IVF, as these are only partially covered by public health insurance. Since these treatments are used predominantly by women older than 35, such changes are going to affect disproportionately the group of women above 35 years old.[18]

Fig. 5 replicates the analysis relating age-specific dependent variables to the corresponding age-specific unemployment rates; the corresponding coefficients are reported in Appendix Table B.1. Age-specific unemployment rates reflect the conditions that women in each age group face in the labor market, particularly if the labor market is segmented.[19] For the IV estimation, I construct age-specific instruments that measure the predicted employment level for each age group, as described in Section 4.

The OLS estimates in panel (a) confirm that the fertility rate tends to respond negatively to increases in unemployment rates while abortion outcomes increase with unemployment. Moreover, these estimates suggest that the response of fertility (and abortion) rates to changes in age-specific unemployment rates is (inversely) U-shaped over the life cycle. While the estimates are neither statistically nor economically significant for young women, they are significant for older women, peaking for women in their thirties and declining (in absolute terms) for women in their forties.

The IV estimates in panel (b) similarly point to an increasing response across the age profile of women. Similarly to results in Table 1, the IV estimates are larger in magnitude than the corresponding OLS estimates. Higher youth unemployment slightly increases the fertility rates of younger women, while higher unemployment for the middle-aged group largely decreases their fertility rates.[20] Specifically, a one standard deviation increase in youth unemployment brings about a marginally significant 0.2 standard deviation increase in the teen fertility rate (95% CI [0.05,0.34]), while the effect is not statistically different from zero for women in their twenties. The response is most marked for women in their thirties: a one standard deviation increase in their unemployment rate decreases the general fertility rate of women aged 30–34 and 35–39 by 0.8 and 1.4 standard deviations respectively (95% CI [−1.14, −0.55] and [−1.65,−1.19]). In addition, also women in their forties significantly reduce their fertility rates by almost half a standard deviation. Regarding age-specific abortion rates, the IV analysis confirms that unemployment rates do not have a statistically significant impact on the abortion choice of women in the youngest age groups, while they play a role for older women. A one standard deviation increase in the respective age-specific unemployment rates brings about a 0.4 standard deviation increase in the abortion rate of women in the 25–29 age group (95% CI [0.07,0.68]), which rises up to a 0.6 standard deviation increase for women in the 40–44 age group (95% CI [0.32, 0.82]). Interestingly, the response of abortion rates is not statistically significantly different from zero for women in the 30–34 age group, suggesting that omitted variables such as childbearing preferences play a larger role for this group. Finally, IV estimates for abortion ratios confirm the small effect on the abortion propensity of young women and the positive effect for older women, which follows naturally from the decrease in fertility rates and increase in abortion rates attested by the previous plots. The increase in the propensity to abort conditional on pregnancy concentrates among women between 35 and 44 years old, amounting to almost a one standard deviation increase following a 1 standard deviation increase in their unemployment rate.

Overall, a common thread emerges from this analysis: unemployment rates have a statistically significant effect on childbearing and abortion, and their relevance for fertility choice is increasing over the life cycle. Notably, unemployment rates have a limited positive effect on young women's fertility and no statistically significant impact on the abortion rates of women in the youngest age group, while they decrease fertility rates and increase abortion rates of older women. The fact that the response of childbearing to changes in unemployment is increasing in magnitude across the age profile is consistent with Schaller (2016) and Del Bono et al. (2012). Since the average age at first birth in Italy is 32 (Istat, 2017), the reduction of births from women aged between 25 and 35 is likely to translate into a postponement of first births. The large response of women between 35 and 49 years old also suggests a potential permanent reduction in fertility since these women are closer to the end of their reproductive cycle. This strong reduction in births might be explained by a decrease in marriage rates, an increase in divorce rates, or a change in the willingness to invest in the quality of children when financial resources are more scarce or uncertain. The change in births for the oldest age group also captures changes in takeup of assisted reproductive treatment; however, given the limited diffusion of IVF procedures[18], any changes in the use of these treatments are likely only a marginally contributing factor to the observed decrease in births.

5.2. *Geographic heterogeneity*

This section considers the heterogeneity in the response to changes in unemployment rates distinguishing by geographical area. Because of the differences in labor market characteristics, access to abortions, and female empowerment, we can expect the reaction of fertility and abortion to show substantial variation across geographic areas.

Fig. 6 presents the standardized effects disaggregating by geographical area; the corresponding estimates are reported in Appendix Table B.2. All variables are standardized at the national level to preserve the comparability of the estimates. The estimated coefficients have consistent signs across geographical areas, but notable differences in magnitudes. In the South, fertility and abortion outcomes show lower responsiveness to changes in the unemployment rate. The IV estimation results indicate that the general fertility rate responds similarly in Central and Northern provinces, while the reaction of the abortion rate is driven by Central Italy. Moreover, the IV estimates for the abortion rate and ratio are not significantly different from zero in the Southern provinces.

I explore two potential mechanisms behind this heterogeneity, which in particular might explain the low response from Southern areas. First, in provinces with a stronger aversion to abortion, higher social costs might lead to lower sensitivity to economic factors. Marie and Zwiers (2023) show that demand and supply-side religious views can influence contraceptive use and thus fertility. These factors could also play a role in abortion decisions, affecting both the level and sensitivity to economic factors of fertility and abortion outcomes.

[17] In 2016, the average age of the mother at first birth in Italy was 32 (Istat, 2017). Hofmann and Hohmeyer (2013) show that couples with children respond significantly to economic concerns by reducing fertility, while childless couples do not.

[18] In 2016 the average age of women resorting to assisted reproduction was 37, and these treatments accounted for 3% of live births in Italy (Ministero della Salute, 2018)

[19] Relating age-specific dependent variables to the overall unemployment rate yields similar results, available from the author upon request.

[20] Estimates relating age-specific outcomes to the total unemployment rate confirm that differences in response are not driven by variations in the units of unemployment measurement. Moreover, while the estimates reported an increase over the age profile, the standard deviation of age-specific unemployment rates decreases across the age groups (Table A.1). Therefore translating the coefficients into a comparable unit change in unemployment would only magnify the finding that the response of fertility and abortion increases across age categories.

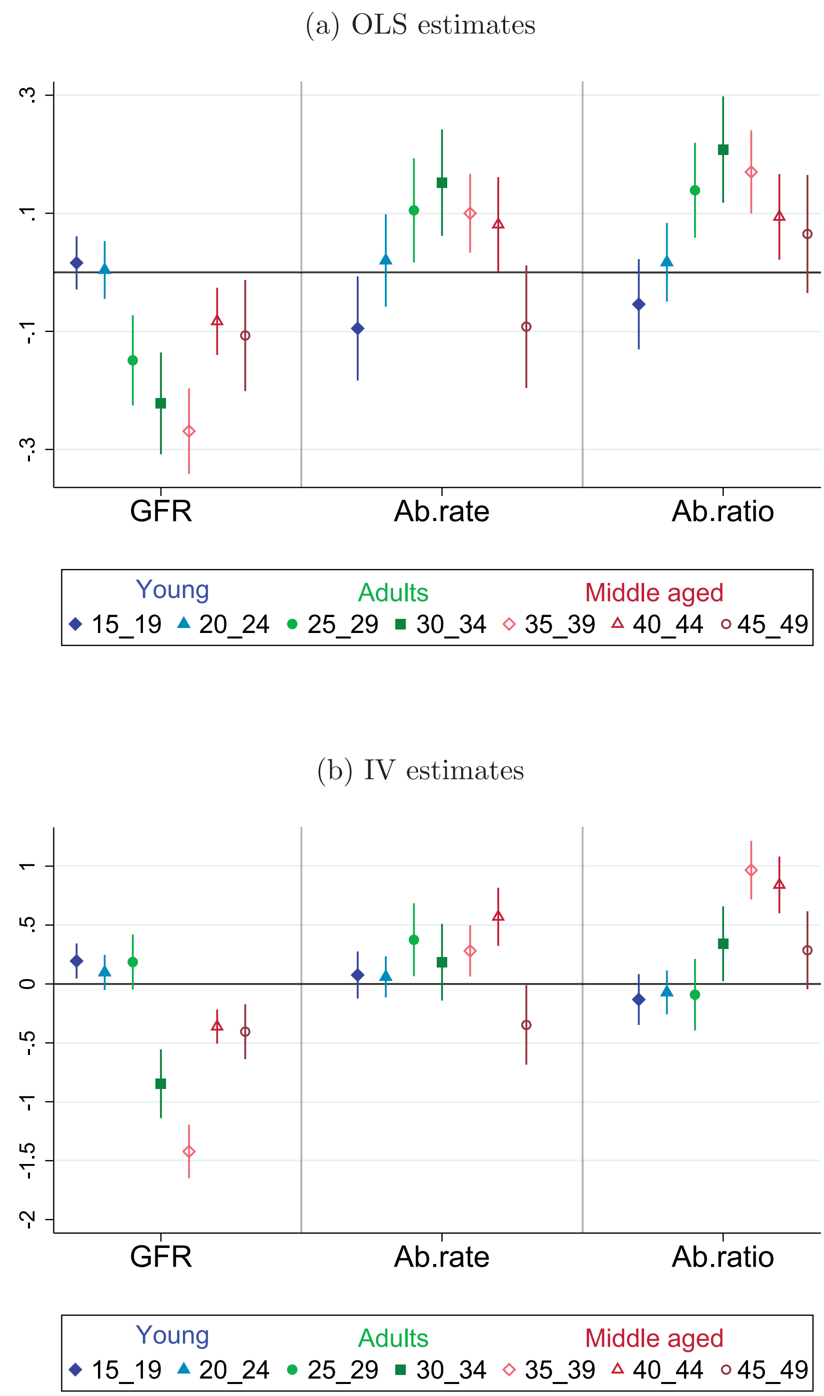

Fig. 5. Regression estimates - age heterogeneity. Notes: 95% CI reported.

While the effect of local social norms on outcome levels is captured by province fixed effects, I investigate the potential lower sensitivity to economic factors using historical views on abortion in the 1981 referendum against abortion legalization.[21] Second, higher informality in the labor market might translate into less salience of unemployment rates for agents' choices or higher uncertainty of their employment status.

In Table 2, I proxy for these two channels using historical views on abortion and the share of irregular workers at the province level. The coefficients for the response to the unemployment rate are stable across specifications and close to the baseline estimate, particularly in the IV estimation. Columns (1), (4), and (7) suggest that local views against abortion do not have a statistically significant effect on fertility or abortion choice; though the OLS coefficient for the fertility rate is marginally significant and positive, this is not confirmed by the IV estimation. The share of irregular workers is a significant factor in determining the fertility and abortion rate, as shown by the IV estimates in columns (3) and (6): areas with more irregularity in the labor market have a higher incidence of both births and abortions. The

[21] More specifically, I consider the share of favorable votes to the referendum question opposing the legalization of induced abortion introduced by Law 194/78.

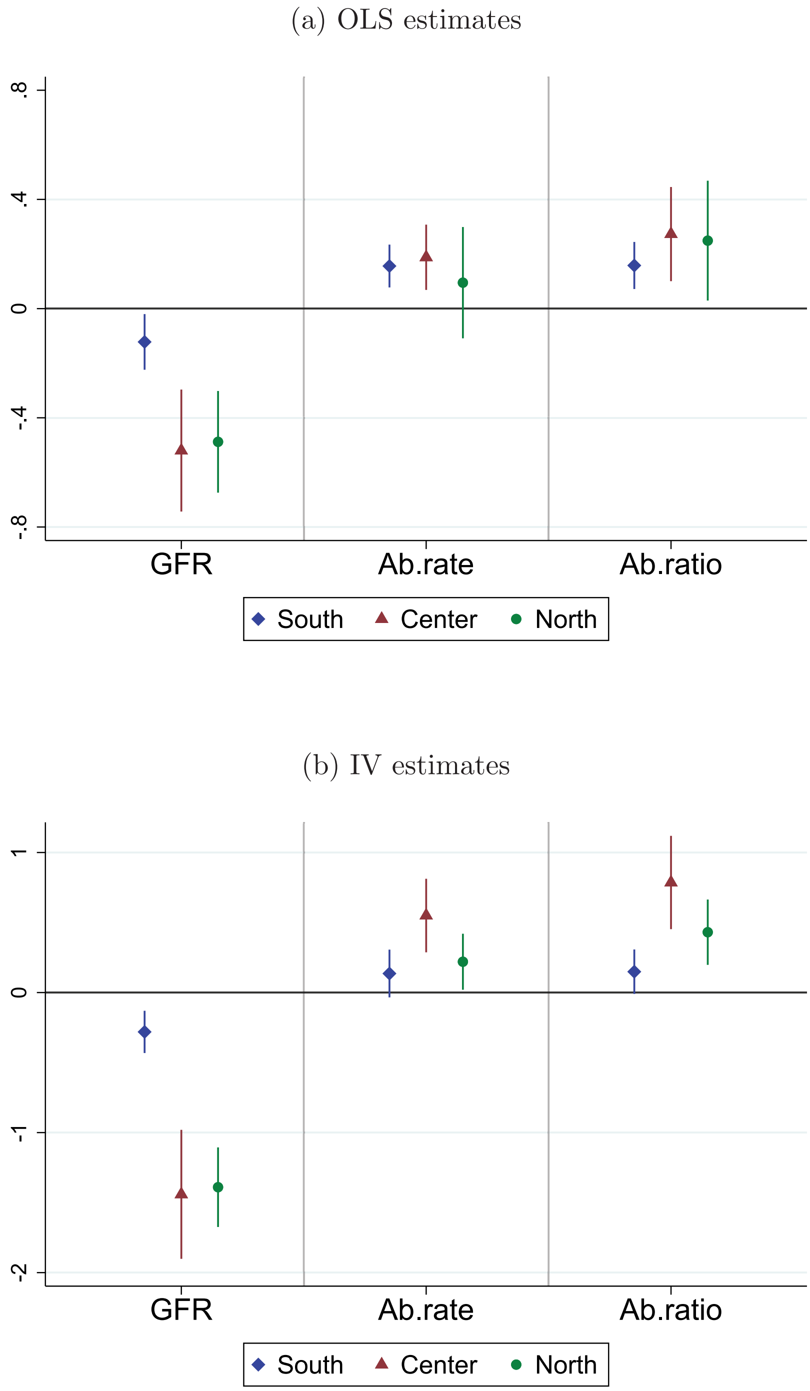

Fig. 6. Regression estimates - geographical heterogeneity. Notes: 95% CI reported.

coefficients on the interaction term between unemployment and the share of irregular workers suggest that a context of labor informality makes unemployment rates less relevant for fertility and abortion choices. Labor informality could therefore contribute to the null effect in Southern provinces, but the limited size of the estimates suggests that this factor alone does not fully explain the result.

6. Robustness checks

This section discusses several robustness checks, presented graphically in Figs. 8 and 7; the corresponding regression estimates are reported in Table B.3 of the Appendix. Additional checks are reported in the Appendix, in particular addressing measurement error in the abortion data in Appendix B.2.

Fig. 7 explores the sensitivity of results to alternative specifications of time. The linear time trend in the baseline specification removes the monotonic relationship between the outcome variables and time, thus leaving potential non-linear behavior in the residual term. The second specification considers province-specific linear time trends; since health services are administered at the regional level, provinces in different regions might have been following different trends. The third specification considers a quadratic time trend, and last specification includes time dummies for the introduction of potentially confounding national policies: the availability of medication abortion from 2009, a baby

Table 2
Heterogeneity - attitudes towards abortion and informal labor markets.

	GFR			Ab. rate			Ab. ratio		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					OLS				
Unempl	−0.215***	−0.361***	−0.343***	0.143***	0.143***	0.150**	0.189***	0.203***	0.189***
	(0.048)	(0.050)	(0.050)	(0.032)	(0.050)	(0.048)	(0.036)	(0.050)	(0.048)
Ref.$_{1981}$*Unempl	0.128*		0.085	0.017		0.030	−0.066		−0.057
	(0.053)		(0.050)	(0.043)		(0.044)	(0.044)		(0.046)
%Irreg		0.132	0.154		0.120	0.129		0.128	0.115
		(0.10)	(0.98)		(0.083)	(0.080)		(0.084)	(0.081)
%Irreg *Unempl		0.126***	0.108**		−0.002	−0.009		−0.017	−0.004
		(0.035)	(0.037)		(0.031)	(0.032)		(0.0323)	(0.035)
					IV				
Unempl	−0.957***	−0.917***	−0.916***	0.250***	0.225***	0.216***	0.380***	0.335***	0.342***
	(0.089)	(0.079)	(0.078)	(0.069)	(0.063)	(0.062)	(0.073)	(0.067)	(0.066)
Ref.$_{1981}$*Unempl	0.018		−0.005	0.039		0.054	−0.054		−0.038
	(0.049)		(0.046)	(0.039)		(0.038)	(0.045)		(0.046)
%Irreg		0.216***	0.214**		0.107	0.122*		0.108	0.097
		(0.065)	(0.067)		(0.063)	(0.062)		(0.062)	(0.062)
%Irreg *Unempl		0.267***	0.267***		−0.045	−0.054		−0.080**	−0.074*
		(0.029)	(0.030)		(0.029)	(0.029)		(0.031)	(0.031)
Observations	1236	1236	1236	1339	1339	1339	1339	1339	1339
R^2	0.589	0.597	0.602	0.562	0.562	0.564	0.356	0.353	0.358
KP LM p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
KP F-stat	176.918	254.518	169.645	139.050	197.915	132.467	139.050	197.915	132.467

Standard errors in parentheses. All regressions include province fixed effects and a linear time trend. With GFR as the dependent variable, all time varying controls are lagged by one year. The local unemployment rates are instrumented for using a leave-one-out Bartik instrument based on the number of employed individuals in each sector, using 2003 weights. The R^2 refers to the OLS estimation.
* $p < .05$, ** $p < .01$, *** $p < .001$.

bonus policy from 2013, the labor market reform, baby bonus policy, and availability of emergency contraception without prescription from 2015.[22] Due to data limitations, there is not sufficient within-period variation to sustain the analysis with yearly time dummies, particularly in the 2SLS estimation.[23] Notably, the point estimates are rather stable across specifications, though in some cases with a loss of precision.

Fig. 8 plots regression estimates for each of the dependent variables for a number of robustness checks, using respectively OLS and IV estimation. The graph first presents the baseline coefficients of Table 1 and then compares them to alternative specifications. The second specification focuses on a restricted sample excluding the provinces of Pesaro, Rimini, and the region of Puglia. This allows to more precisely keep geographical boundaries constant over time since in 2009 seven municipalities belonging to these provinces changed province. In addition, I exclude the region of Puglia because in 2008 it introduced a policy of free hormonal contraception for specific groups, which might have affected abortion behavior. The third specification replicates the analysis of the abortion dependent variables controlling for the regional share of objecting gynecologists. While systematic time-invariant differences in the supply of abortion services across provinces are captured by province dummies, changes in the preferences of doctors over time are a potential source of variation in abortion supply. Indeed, the percentage of objectors in Italy has increased over time, from below 60% in 2002 to over 70% in recent years (Autorino et al., 2020). The robustness of results to these alternative specifications suggests that the results are not driven by changes in geography nor in the availability of abortion services.

7. Conclusion

This paper explores the relationship between local unemployment rates and fertility choice, focusing on both childbearing and abortion behavior. The literature has been aware of the role of economic conditions in childbearing, suggesting that in most cases fertility behaves procyclically. However, the same attention has not been dedicated to understanding the response of abortion behavior, particularly using causal methods.

I investigate the response of births and abortions to changes in the unemployment rate in the context of the Great Recession and the Sovereign Debt Crisis in Italy. The empirical results suggest that as unemployment conditions worsen the general fertility rate decreases, while the abortion rate and the propensity to abort conditional on pregnancy increase. Thus, both childbearing and abortion behave procyclically. This evidence indicates that pro-cyclical behavior of fertility rates arises not just from changes in planned pregnancy but also from a higher incidence of abortions.

Further analysis indicates that the reaction is increasing over the age profile, where younger women exhibit little to no adjustment in their fertility behavior but older women show significant changes. Dynamic models of fertility predict that transitory fluctuations in wages might affect the timing of births, inducing a postponement of fertility, but they will not impact expected total fertility in the presence of perfect capital markets and certainty (Happel et al., 1984; Hotz et al., 1997). The reaction of women in the 24–34 age group can be reasonably interpreted as a fertility postponement, as the median age at first birth is 32. However, the reaction of women aged between 34 and 49 is even stronger. These women are the most likely to already have children but are also closer to the end of their reproductive life. A strong reaction from this group might therefore signal changes in their demand for additional children or having children at all. Future research can delve into this question and explore whether fertility and abortion responses vary by parity, i.e. whether the changes affect the intensive or extensive margin of fertility, and whether the total realized fertility of women exposed to recessions in their last reproductive years is affected.

Moreover, this study sheds light on the heterogeneous response to changes in unemployment rates across different regions in Italy. The findings indicate consistent signs but notable variations in magnitudes across regions, with the South displaying lower responsiveness in both fertility and abortion outcomes to changes in the unemployment rate. Further results suggest that labor market informality can diminish the

[22] For a detailed description of these policies, see Appendix B.2.
[23] Furthermore, recent literature has underlined how the use of TWFE fundamentally rests on modeling assumptions, rather than being a nonparametric estimation strategy (Imai and Kim, 2021). Though this research has focused on cases of binary treatment, the use of TWFE is likely to be just as problematic with continuous treatment such as the one under consideration here.

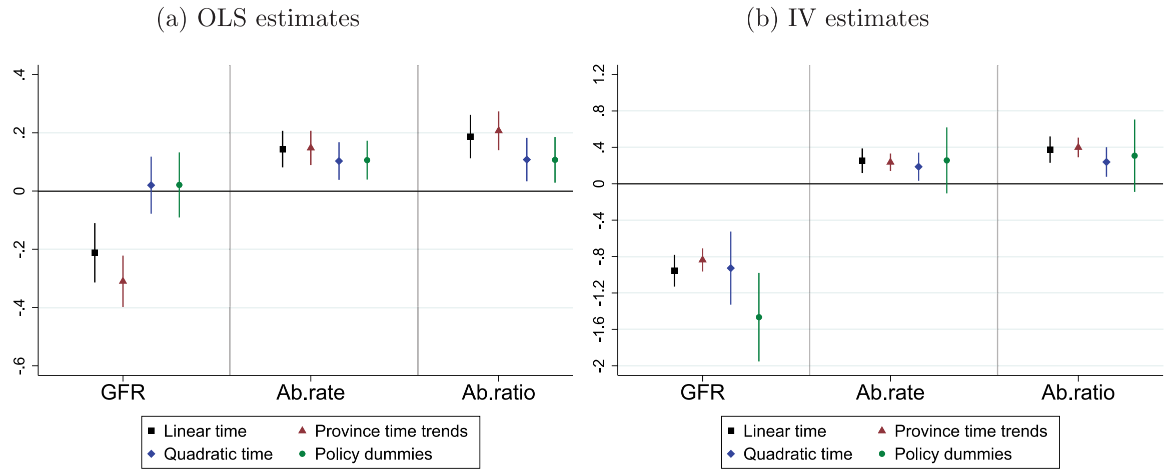

Fig. 7. Time specification.

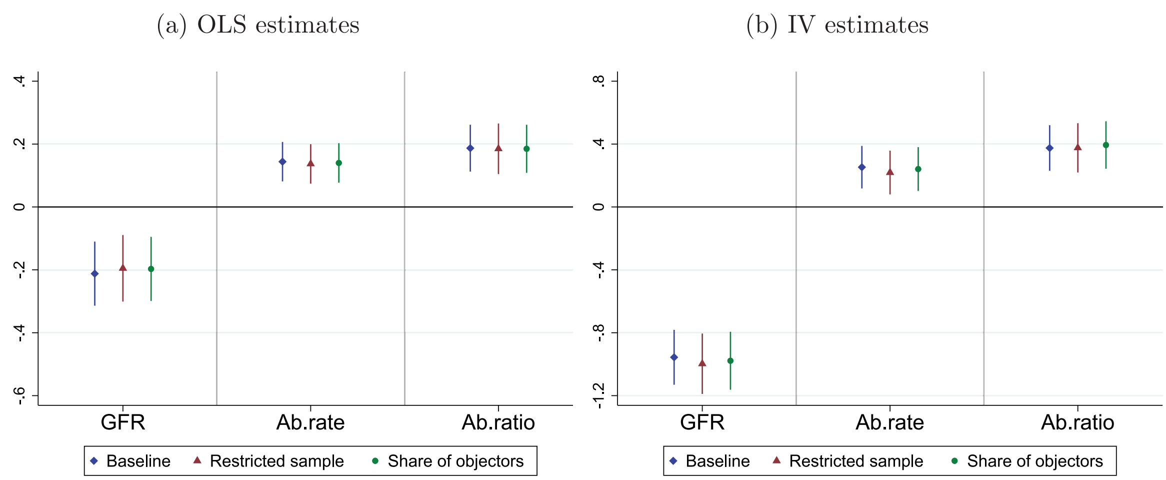

Fig. 8. Robustness checks. Notes: 95% CI reported.

relevance of unemployment rates for fertility and abortion choices, while there is no evidence of a heterogeneous response based on local anti-abortion views.

This study provides supportive evidence that both fertility and abortion rates are influenced by economic fluctuations, an important fact to be taken into account by policymakers. The procyclical behavior of these outcomes emphasizes the importance of implementing policies that address the impact of the business cycle on reproductive decisions. These can include child subsidies, guaranteed access to abortion and family planning services, and subsidized contraception during unemployment spells. The latter has recently been introduced in some regions, and future research can explore the impact of such policies and provide valuable insights for improving social and health outcomes. The findings also have implications regarding the allocation of public resources, both financial and human, in particular in the presence of budget cuts to the health sector.

Data availability

The data that has been used is confidential.

Acknowledgments

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Appendix A. Sample selection and summary statistics

Table A.1 reports the summary statistics for the province-level data. On average, the fertility rate is highest for women between 25 and 35 years old, but this age class also corresponds to the highest abortion rate. On the other hand, once accounted for the number of estimated pregnancies, the propensity to abort is much higher for the younger (15–24) and older (35+) women. This means that childbirth and abortion are relatively less frequent in the younger and older female population because these age classes have fewer pregnancies but, in the event of a pregnancy, they are more likely to abort. The mean unemployment rate throughout the sample is around 9%, and it is particularly high for the younger section of the population, with a mean of 29%. Moreover, the share of irregular workers is 12% on average, with a maximum of 25%.

Table A.1
Summary statistics - province data.

	Mean	Median	SD
Fertility rate by age group, per 1000 women			
GFR	36.21	35.96	3.35
GFR 15–24	12.83	11.19	5.16
GFR 25–34	76.82	76.32	7.82
GFR 35–49	22.55	22.43	2.81
Abortion rate by age group, per 1000 women			
Ab.rate	5.43	5.28	1.57
Ab.rate 15–24	6.65	6.42	2.13
Ab.rate 25–34	7.74	7.60	2.28
Ab.rate 35–49	3.42	3.32	1.03
Abortion ratio by age group, per 1000 pregnancies			
Ab.ratio	129.83	127.51	33.14
Ab.ratio 15–24	350.98	356.16	94.88
Ab.ratio 25–34	91.84	90.56	26.08
Ab.ratio 35–49	132.66	126.09	40.78
Unemployment rate by age group			
Unempl	9.37	8.12	5.26
Unempl 15–24	29.20	28.05	14.25
Unempl 25–34	13.35	11.08	8.60
Unempl 35–64	6.09	5.39	3.43
Additional economic indicators			
Empl. rate	57.88	62.30	9.79
Irregular workers, %	12.68	10.77	4.57
Real GDP per capita (000 €)*	22.13	22.36	5.84
Real Value Added per capita (000 €)*	19.89	20.06	5.17
Referendum on abortion law (1981)			
Yes votes, %	29.19	29.54	5.78
Observations		1,339	
Provinces		103	

Notes: The table provides within-cell means for the 103 provinces used in the baseline specification. The share of irregular workers is measured at the regional level.
* GDP and Value Added are CPI adjusted to 2004 Euros.

Fig. A.1 provides an additional discussion of the significant differences between Italian and foreign women who have an abortion in Italy, plotting the different incidence of abortions for these two populations. Panel (a) underlines a substantial difference between these two populations: on average, we observe 6 abortions per 1000 Italian women and 32 abortions per 1000 foreign women. Importantly, the abortion rate of foreigners is not comparable to that of any Italian region. The higher abortion rate among foreigners suggests different cultural norms apply to this population, indicating for example either more liberal views about abortion or lower use of contraception of this group compared to natives. Panel (b) shows that the difference between the abortion rate means of these two groups is statistically significantly different from zero in all regions.

Finally, Fig. A.2 shows the geographic variation in the propensity to abort conditional on pregnancy and in local views against abortion, proxied by the local vote share against the liberalization of abortion in the historical referendum of 1981. Note that the propensity to abort in Fig. A.2(a) reflects closely the variation in the abortion rate in Fig. 1(b). In the North and Center areas, there is a clear negative relation between these two variables. Interestingly, in Southern Italy the propensity to abort seems to go hand in hand with conservatism towards abortion, confirming the pattern already present in Fig. 1 regarding abortion rates and constraints in the supply of abortion services.

Appendix B. Additional results

B.1. Tables corresponding to graphical evidence

This section tabulates the main regression estimates, corresponding to the graphs shown in the main text. Table B.1 reports the estimates plotted in Fig. 5, analyzing the response of age-specific outcomes to changes in age-specific unemployment rates. Table B.2 reports the estimates plotted in Fig. 6, analyzing the heterogeneity of estimates across geographical areas.

Table B.3 reports the estimates for several robustness checks for each dependent variable; these are the same presented graphically in Fig. 8. Results are fundamentally unchanged, as coefficients remain close to the benchmark in Table 1. Columns (1–2) include time dummies for the introduction of potentially confounding national policies: the availability of medication abortion from 2009, a baby bonus policy from 2013, the labor market reform, baby bonus policy, and availability of emergency contraception without prescription from 2015. These policies are described in detail in the next subsection. Columns (3–4) add regional linear time trends to the baseline specification of Eq. (2). Since health services are administered at the regional level, different regions might have been following different trends. Columns (5–6) present the results when excluding from the sample the provinces of Pesaro, Rimini, and the region of Puglia. Finally, columns (7–8) control for the share of objecting gynecologists for the abortion dependent variables, accounting for potential changes in the preferences of doctors over time.

B.2. Robustness checks

This section reviews various robustness checks, first reporting the estimates of alternative specifications considered in the main text and second discussing potential measurement error in the abortion dependent variables.

Description of confounding national policies

This section discusses in detail the potentially confounding national policies implemented between 2004 and 2016, which are accounted for in the alternative specification of Table B.3.

Birth allowances and childcare vouchers have been implemented in different forms and magnitudes over the years, as they are approved and fine-tuned yearly in the national budget plan. Although demographic experts have not reached a consensus as to whether such short-term measures are effective in redressing the population imbalance, they might have provided temporary support for mothers in

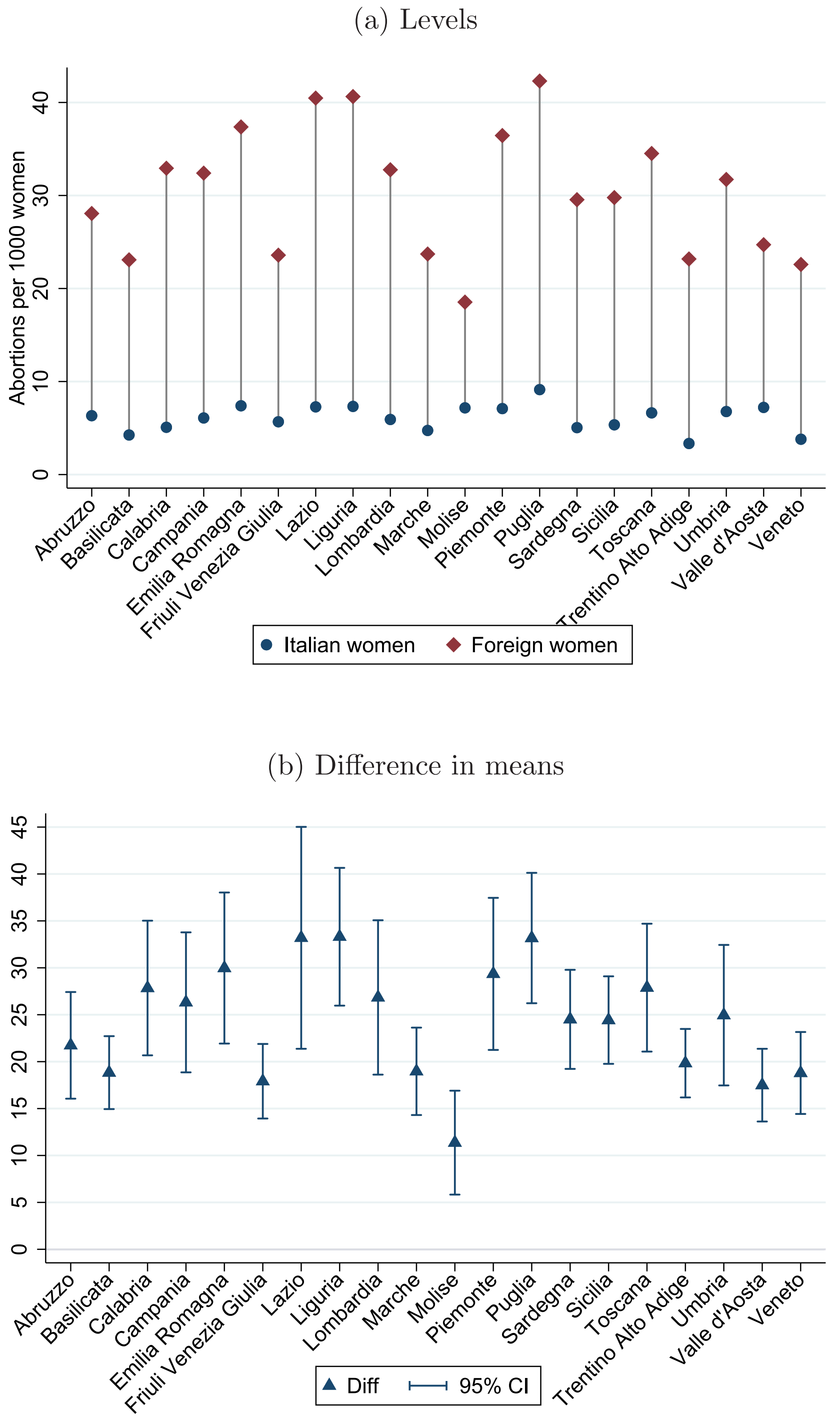

Fig. A.1. Abortion rates by nationality.
Notes: Panel (a) reports the time averaged abortion rate for Italian and foreign women of childbearing age in each region. Panel (b) reports the results from a t-test comparison of the two group means, allowing for unequal variances.

uncertain economic times (Drago et al., 2011; Malak et al., 2019). In 2009, an initiative was launched to offer loans at a subsidized rate for households with newborns; then in 2012 the government introduced a monthly voucher for babysitting and childcare services targeted at full-time working mothers with babies born between 2013 and 2016. Moreover, in 2014 a means-tested benefit for medium-low income family households with newborns was approved.

In addition, the 2014 Jobs Act (Law 78/2014) implemented a wide-ranging reform of the Italian labor market, reducing firing costs and eliminating restrictions on the use of temporary contracts with the aim of reducing unemployment and labor market dualism. The changes induced by the reform might have affected both the unemployment rate and job security and consequently fertility and abortion through unobserved factors.

Finally, some regulatory changes interested the abortion procedure directly. Medication abortion was introduced in 2009, but its use remains rather limited and there is no evidence that it has induced an increase in abortion incidence. The share of medication abortions

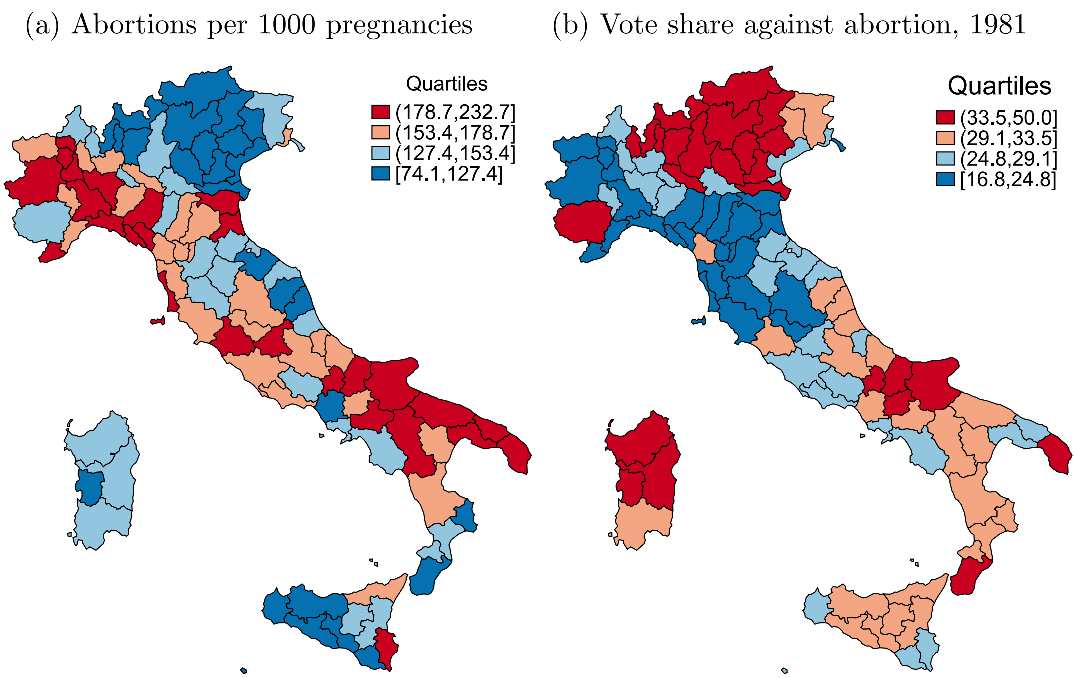

Fig. A.2. Geographic variation in abortion attitudes.
Notes: Panel (a) reports the abortion ratio in 2004 for each province. Panel (b) reports the share of votes opposing the legalization of induced abortion in the referendum vote of 1981.

Table B.1
Age-specific outcomes and age-specific unemployment rates.

(a) OLS

Age group	(1) 15–19	(2) 20–24	(3) 25–29	(4) 30–34	(5) 35–39	(6) 40–44	(7) 45–49	(8) Obs.
GFR	0.016 (0.023)	0.004 (0.025)	−0.149*** (0.039)	−0.222*** (0.044)	−0.269*** (0.037)	−0.083** (0.029)	−0.107* (0.048)	1236
Ab.rate	−0.095* (0.045)	0.020 (0.040)	0.105* (0.045)	0.152** (0.046)	0.100** (0.034)	0.081* (0.041)	−0.092° (0.053)	1339
Ab.ratio	−0.054 (0.039)	0.017 (0.034)	0.139** (0.041)	0.208*** (0.046)	0.170*** (0.036)	0.094* (0.037)	0.065 (0.051)	1339

(b) IV

Age group	(1) 15–19	(2) 20–24	(3) 25–29	(4) 30–34	(5) 35–39	(6) 40–44	(7) 45–49	(8) Obs.
GFR	0.195* (0.076)	0.098 (0.076)	0.186 (0.119)	−0.847*** (0.149)	−1.422*** (0.116)	−0.362*** (0.074)	−0.405*** (0.119)	1236
Ab.rate	0.076 (0.102)	0.060 (0.089)	0.375* (0.158)	0.185 (0.166)	0.281* (0.111)	0.570*** (0.126)	−0.348* (0.172)	1339
Ab.ratio	−0.132 (0.110)	−0.072 (0.095)	−0.091 (0.155)	0.341* (0.162)	0.966*** (0.127)	0.841*** (0.123)	0.286° (0.169)	1339 subfigure

Standard errors in parentheses. All regressions include province fixed effects and a linear time trend.
$^{\circ}p < .10$, $^{*}\ p < .05$, $^{**}\ p < .01$, $^{***}\ p < .001$.

increased over time, from 5% in 2010 to 16% in 2016, particularly in the North and Center regions (Ministero della Salute, 2016). Moreover, from 2015 emergency contraception can be purchased in pharmacies without a medical prescription, except in the case of minors.[24]

To control for such policies, in Table B.3 I replicate the analysis including an indicator variable for years from 2009 onwards to account for the availability of medical abortion; a time dummy from 2013 onwards to account for the kindergarten voucher; and a time dummy for years from 2015 onwards to capture the effect of the baby bonus policy, the labor market reform and emergency contraception access.

Measurement error in the abortion data

Table B.4 presents estimates for the abortion outcomes taking into account various sources of measurement error in the abortion data. Although the filing of Modello D12 is compulsory for all induced abortions performed in public or private hospitals, data can be missing for a number of reasons.

First, the data only reflects the number of legally performed abortions. According to the Ministry of Health, covert abortions accounted for only 20% of total abortions by Italian women in 2016 (Ministero della Salute, 2016), thus the data captures the vast majority

[24] This regulatory change initially applied to the Ellaone, Stromalidan, and Escapelle emergency contraception drugs; the same provision was extended to Norlevo in 2016.

Table B.2
Geographic heterogeneity - standardized.

	GFR			Ab.rate			Ab.ratio		
	South	Center	North	South	Center	North	South	Center	North
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					OLS				
$Unempl_{t-1}$	−0.122*	−0.520***	−0.488***						
	(0.052)	(0.111)	(0.093)						
$Unempl_t$				0.156***	0.188**	0.095	0.158***	0.273**	0.249 *
				(0.040)	(0.061)	(0.104)	(0.044)	(0.088)	(0.112)
					IV				
$Unempl_{t-1}$	−0.281***	−1.411***	−1.390***						
	(0.077)	(0.235)	(0.145)						
$Unempl_t$				0.136	0.559***	0.220	0.149	0.786***	0.431 **
				(0.087)	(0.134)	(0.102)	(0.081)	(0.170)	(0.119)
Observations	432	252	552	468	273	598	468	273	598
R^2	0.539	0.523	0.673	0.477	0.742	0.574	0.284	0.530	0.350
KP LM p-value									
KP F-stat	192.9	83.88	274.6	145.8	67.35	259.0	145.8	67.35	259.0

Standard errors in parentheses. All regressions include province fixed effects and a linear time trend. The local unemployment rates are instrumented for using a leave-one-out Bartik instrument based on the number of employed individuals in each sector, using 2003 weights. The R^2 refers to the OLS estimation.
* $p < .05$, ** $p < .01$, *** $p < .001$.

Table B.3
Robustness.

(a) GFR

	Policy dummies		Province time trends		Restricted sample	
	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)
Unempl_{t-1}	0.021	−1.146***	−0.310***	−0.837***	−0.190***	−0.970***
	(0.057)	(0.248)	(0.045)	(0.052)	(0.053)	(0.096)
Observations	1236	1236	1236	1236	1164	1164
R^2	0.630		0.737		0.570	
KP F		108.564		656.199		328.6
KP LM pval		0.000		0.000		0.000

(b) Abortion rate

	Policy dummies		Province time trends		Restricted sample		% Objectors	
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Unempl_t	0.093**	0.307*	0.148***	0.236***	0.127***	0.253**	0.120***	0.260***
	(0.035)	(0.134)	(0.033)	(0.042)	(0.034)	(0.082)	(0.033)	(0.077)
Obs	1339	1339	1339	1339	1261	1261	1301	1301
R^2	0.561		0.684		0.546		0.536	
KP F		101.869		688.869		254.420		255.389
KP LM pval		0.000		0.000		0.000		0.000

(c) Abortion ratio

	Policy dummies		Province time trends		Restricted sample		% Objectors	
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Unempl_t	0.093*	0.349*	0.207***	0.399***	0.168***	0.394***	0.162***	0.397***
	(0.040)	(0.142)	(0.037)	(0.047)	(0.041)	(0.086)	(0.038)	(0.081)
Obs	1339	1339	1339	1339	1261	1261	1301	1301
R^2	0.368		0.498		0.328		0.328	
KP F		101.869		688.869		254.520		255.389
KP LM pval		0.000		0.000		0.000		0.000

Standard errors in parentheses. All regressions include province fixed effects and a linear time trend. * $p < .05$, ** $p < .01$, *** $p < .001$
The restricted sample in columns (1–2) excludes the provinces of Pesaro-Urbino, Rimini, and the region of Puglia from the sample.
The policy dummies used in columns (3–4) control for the introduction of medication abortion from 2009, a baby bonus policy from 2013, a labor market reform in 2015, and emergency contraception without prescription from 2015.
The local unemployment rates are instrumented for using a leave-one-out Bartik instrument based on the number of employed individuals in each sector, using 2003 weights.

of abortions. Moreover, covert abortions are partially measured by miscarriages, accounted for in Table B.4.

Second, women can choose not to share their personal data, in which case the abortion is recorded in the data but all information regarding the province of residence and birth is omitted. This appears to be a minor concern since only 1% of observations do not report the province of residence. Third, there are inconsistencies in the transmission of data from regions to the National Statistics Agency.

Table B.4
Measurement error in abortion data.

	Incomplete region-years				Miscarriages			
					Italian		All	
	OLS (1)	IV (2)	OLS (3)	IV (4)	OLS (5)	IV (6)	OLS (7)	IV (8)
Ab.rate	0.130**	0.394***						
	(0.041)	(0.076)						
Ab.ratio			0.169***	0.563***	0.175***	0.380***	0.170***	0.405***
			(0.048)	(0.081)	(0.037)	(0.073)	(0.037)	(0.072)
Obs	1094	1094	1094	1094	1339	1339	1339	1339
$\bar{Y}$	5.54	5.54	132.84	132.84	116.80	116.80	114.45	114.45
R^2	0.639		0.439		0.355		0.381	
KP LM		0.000		0.000		0.000		0.000
KP F-stat		238.972		238.972		285.542		285.542

Columns (5–6) measure the estimated number of pregnancies including miscarriages of Italian women by the province of abortion; columns (7–8) consider the number of miscarriages by the province of residence, but independently of citizenship.
The local unemployment rates are instrumented for using a leave-one-out Bartik instrument based on the number of employed individuals in each sector, using 2003 weights.
Standard errors in parentheses. All regressions include province fixed effects and a linear time trend.
* $p < .05$, ** $p < .01$, *** $p < .001$.

Table C.1
Sectors used for the Bartik instrument.

NACE code, Rev. 2	Sector
A	Agriculture, forestry and fishing
B–E	Mining and quarrying; manufacturing; electricity, gas, steam and air-conditioning supply; water supply, sewerage, waste management and remediation
F	Construction
G–I	Wholesale and retail trade, repair of motor vehicles and motorcycles; Transportation and storage; Accommodation and food service activities
J	Publishing, audiovisual and broadcasting activities; Telecommunications; IT and other information services
K	Financial and insurance activities
L	Real estate activities
M–N	Professional, scientific and technical activities; Administrative and support service activities
O–Q	Public administration and defence, compulsory social security; Education; Human health services; Residential care and social work activities
R–U	Arts, entertainment and recreation; Other services

Notes: Industries are aggregated to match the industry data available at the province level.

Underreporting the number of abortions leads to a downward measurement error in the abortion rate and the abortion ratio. Comparing the number of reported abortions to official regional estimates, which integrate the incomplete data with hospital discharge data, I find that incomplete reporting affects 17% of region-year cells and that incomplete observations have on average 5% of total abortions missing. Incomplete data affects almost all regions at least once in the observed period, but only Campania and Sicilia report incomplete data for most years in the sample.[25] Columns (1–4) of Table B.4 exclude region-year cells with incomplete data from the analysis. The estimates maintain their statistical significance and are larger in magnitude than in the full sample: a one standard deviation increase in the unemployment rate brings about a 0.38 standard deviation increase in the abortion ratio and a 0.54 standard deviation increase in the abortion ratio.

Finally, the abortions to pregnancies ratio proxies the number of pregnancies only with live births and abortions. Not accounting for stillbirths and miscarriages leads to an upward biased estimate of the abortion ratio and non-random measurement error insofar as the number of miscarriages is related to economic instability, for instance through maternal stress (Bruckner et al., 2016). Columns (5–8) of Table B.4 report the estimation results when including the number of spontaneous abortions in the calculation of the abortion ratio. Columns (5–6) measure the estimated number of pregnancies including miscarriages of Italian women by the province of abortion; columns (7–8) consider the number of miscarriages by the province of residence, but independently of citizenship. Since both corrections remain subject to some measurement error, the preferred proxy for the number of pregnancies remains the sum of live births and induced abortions. Coefficients remain statistically significant at the 0.1% confidence level and suggest that the effect of a 1 standard deviation increase in the unemployment rate ranges between 0.38 and 0.41 abortion ratio standard deviations.

[25] The region-years affected by incomplete data transmission, according to ISTAT, are: Abruzzo (2009, 2012), Basilicata (2009, 2014), Calabria (2008), Campania (2002–2003, 2005–2014), Friuli-Venezia Giulia (2005, 2006), Liguria (2013), Lombardia (2014), Marche (2014), Molise (2005), Puglia (2012, 2013), Sardegna (2008, 2009, 2013–2015), Sicilia (2004–2012, 2014, 2015), Umbria (2010–2012), Veneto (2015, 2016).

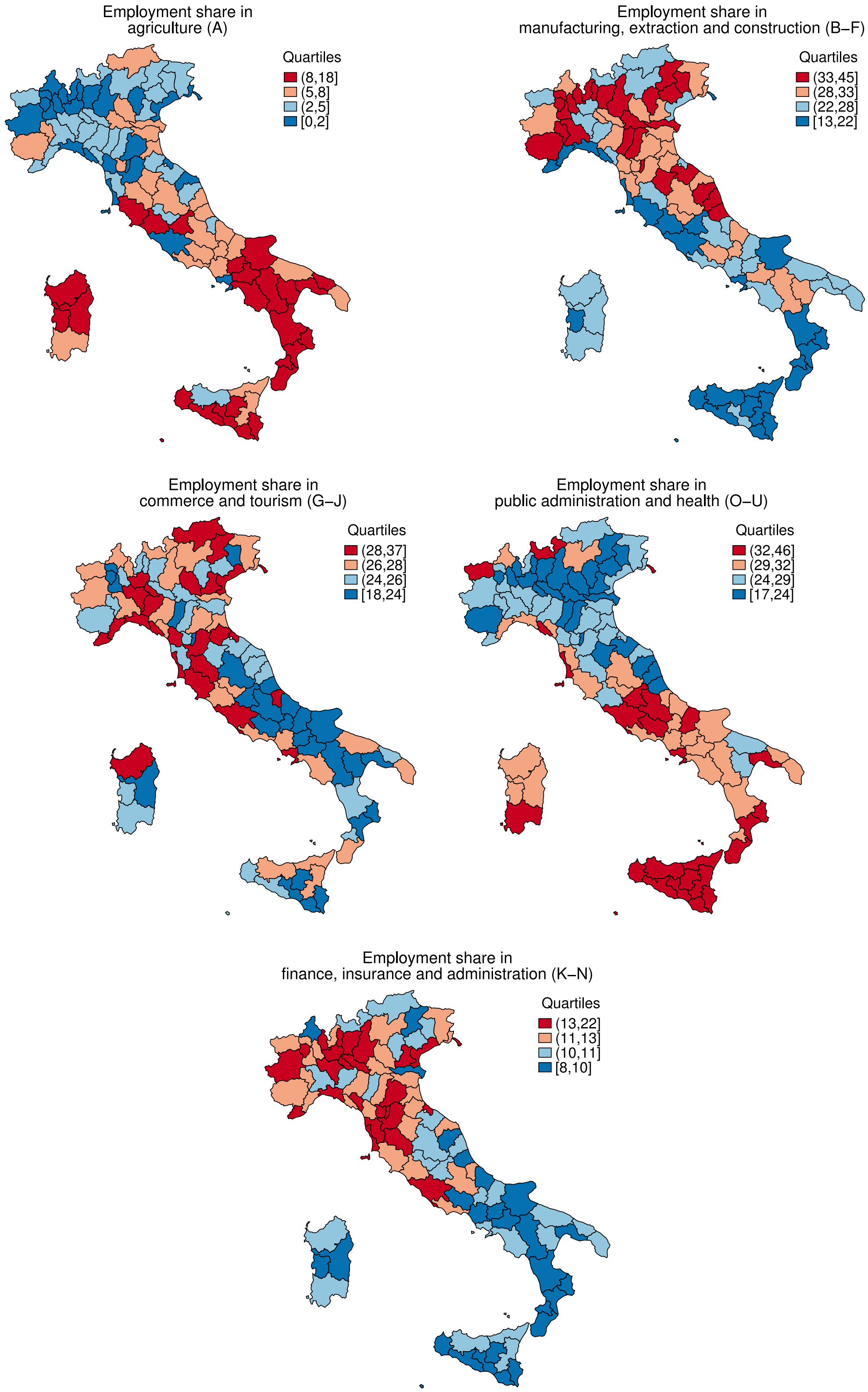

Fig. C.1. Industry shares in 2003.

Table C.2
First stage estimates - standardized.

(a) Main specification

	GFR (1)	Ab.rate & Ab.ratio (2)
	$Unempl_{t-1}$	$Unempl_t$
$Bartik_{t-1}$	−0.511***	
	(0.027)	
$Bartik_t$		−0.460***
		(0.027)
Observations	1236	1339
KP LM p-value	0.000	0.000
KP F-stat	360	285

(b) Age-specific rates

	GFR			Ab.rate & Ab.ratio		
	15–24 (1)	25–34 (2)	35–49 (3)	15–24 (4)	25–34 (5)	35–49 (6)
$Bartik_{t-1}$	−0.889***	−0.659***	−1.008***			
	(0.090)	(0.072)	(0.069)			
$Bartik_t$				−0.740***	−0.479***	−0.871***
				(0.069)	(0.065)	(0.066)
Observations	1236	1236	1236	1339	1339	1339
KP LM p-value	0.000	0.000	0.000	0.000	0.000	0.000
KP F-stat	97	82	213	114	54	176

Standard errors in parentheses. All regressions include province fixed effects and a linear time trend.
The local unemployment rates are instrumented for using a leave-one-out Bartik instrument based on the number of employed individuals in each sector, using 2003 weights.
* $p < .05$, ** $p < .01$, *** $p < .001$.

Table C.3
Alternative Bartik estimators.

(a) Second stage

	E_$empl_{03}$	E_$empl_{00}$	E_wp_{03}	VA_$empl_{03}$	E_LFS_$empl_{04}$	U_LFS_$empl_{04}$
	Dependent variable: GFR					
$Unempl_{t-1}$	−0.956***	−0.974***	−1.124***	−0.777***	−0.421***	−0.604***
	(0.089)	(0.095)	(0.106)	(0.064)	(0.075)	(0.075)
	Dependent variable: Ab. rate					
$Unempl_t$	0.253***	0.254***	0.277***	0.192***	0.165*	0.136*
	(0.069)	(0.074)	(0.076)	(0.049)	(0.065)	(0.065)
	Dependent variable: Ab. ratio					
$Unempl_{t-1}$	0.375***	0.365***	0.421***	0.335***	0.335***	0.300***
	(0.074)	(0.079)	(0.082)	(0.053)	(0.074)	(0.07)
Observations	1339	1339	1339	1339	1236	1236

(b) First stage

	(1) E_$empl_{03}$	(2) E_$empl_{00}$	(3) E_wp_{03}	(4) VA_$empl_{03}$	(5) E_LFS_$empl_{04}$	(6) U_LFS_$empl_{04}$
	$Unempl_{t-1}$					
$Bartik_{t-1}$	−0.511***	−0.524***	−1.642***	−0.575***	−1.144***	0.437***
	(0.027)	(0.030)	(0.101)	(0.025)	(0.077)	(0.026)
Observations	1236	1236	1236	1236	1133	1133
KP LM pvalue	0.000	0.000	0.000	0.000	0.000	0.000
KP F-stat	360.531	310.931	264.509	537.680	222,017	275,393
	$Unempl_t$					
$Unempl_t$	−0.460 ***	−0.465***	1.500***	−0.567***	−1.149***	0.411***
	(0.027)	(0.030)	(0.098)	(0.025)	(0.074)	(0.025)
Observations	1339	1339	1339	1339	1236	1236
KP LM pvalue	0.000	0.000	0.000	0.000	0.000	0.000
KP F-stat	285.542	242.509	233.017	526.748	241.946	274.045

Standard errors in parentheses. All regressions include province fixed effects and a linear time trend.
* $p < .10$, ** $p < .05$, *** $p < .01$
Columns (1)–(3) refer to the leave-one-out Bartik instrument measured using the employment level, weighting industries by their employment share in 2003, 2000 and the working population in 2003, respectively. Column (1) corresponds to the preferred instrument used in the main analysis. Column (4) employs a leave-one-out Bartik instrument measured using industry real value added, weighting industries by their employment share in 2003. Columns (5) and (6) use as share the employment share in 2004 for ATECO industries (2002 classification, 2 digit level) and the leave-one-out number of employed and unemployed as shifts, respectively.

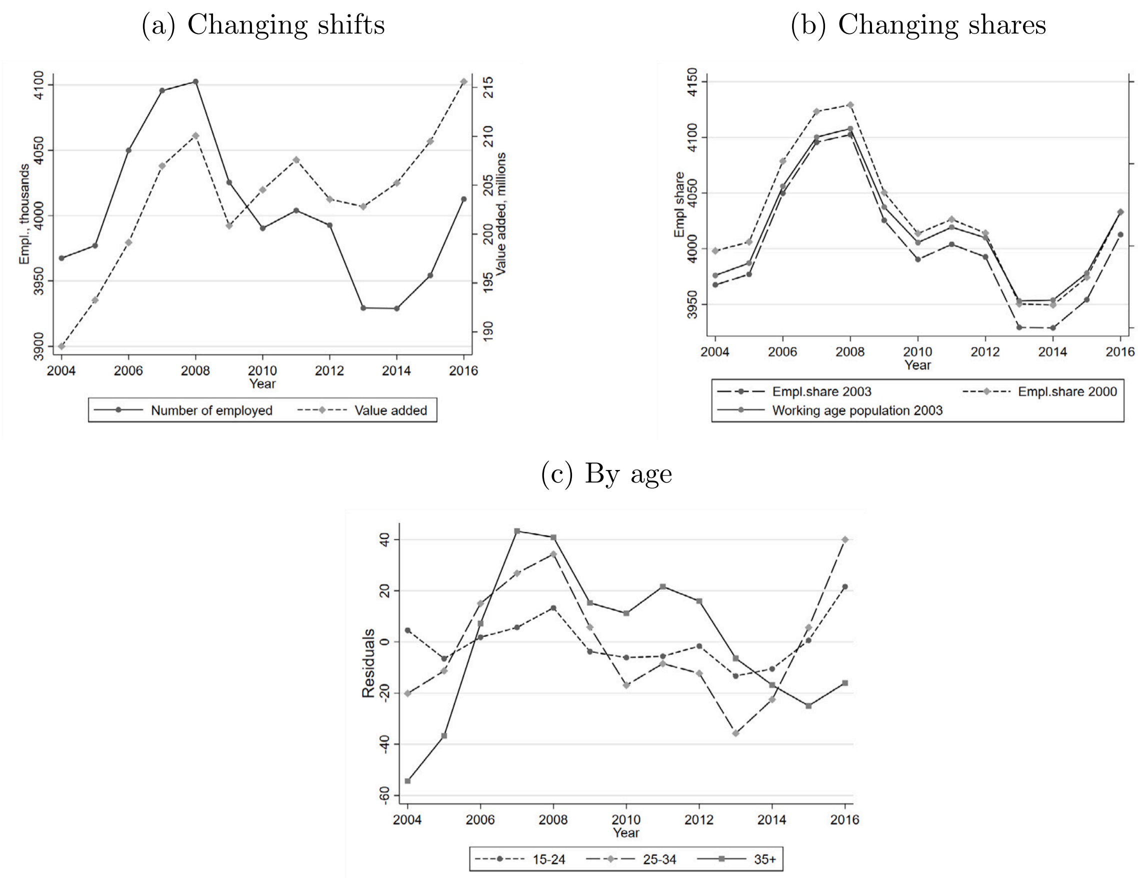

Fig. C.2. Bartik instrument alternatives - average province.
Notes: the figure plots the predicted value of different versions of the instrument for the average province. Panel C.2(a) shows the predicted number of employed individuals and the predicted value added, while Panel C.2(b) shows the predicted number of employed individuals using different shares. Panel C.2(c) plots the deviations of age-specific instruments from the linear time trend.

Appendix C. Bartik instrument

C.1. Industry sectors

Table C.1 reports the list of sectors employed to construct the shift-share instrument used in the main analysis. In total, these are 10 sectors from the ATECO 2007 classification, which corresponds to the NACE Rev.2. Some industries are only available as a group, for example the manufacturing and extraction industries. Additional robustness checks reported in Table C.3 employ an alternative version of the instrument which uses a more refined industry classification (2002 ATECO classification 2 digit level), thus increasing the number of industries considered to around 60.

Fig. C.1 presents the initial geographical variation of the local employment shares for different industries, i.e. the shares used to compute the Bartik instrument. There is substantial variation across provinces in the importance of each sector for the local labor market, with Northern and Central provinces concentrating on industrial production and services, and Southern areas concentrating on public administration and agriculture.

C.2. First stage relationship

The first stage estimates the relationship between the province unemployment rate and the constructed shift-share instrument, which measures the predicted province employment level. If the labor force were constant, changes in the number of employed individuals would perfectly determine changes in the unemployment rate, so the instrument and the unemployment rate would be almost inversely proportional. In the presence of changes in labor force participation, the level of employment will be informative of the unemployment rate insofar as changes in the labor force are not fully absorbed by only employment or only unemployment.

Table C.2 reports the first stage estimates of the IV analysis, for both the main specification and the age-specific instruments. The unemployment rates are negatively correlated with the employment level predicted by the Bartik instrument; Fig. 3 in the main text presents the first stage relationship graphically. The large Kleibergen–Paap F-statistic informs us that the instrument is relevant, and the Kleibergen–Paap LM test rejects the null hypothesis of under-identification.

C.3. Alternative Bartik instruments

This section presents alternative versions of the shift-share instrument and discusses their performance in the main specification (2). For details on the construction of each measure, see the Data appendix below.

Fig. C.2 presents the evolution over time of different versions of the Bartik instrument, taking the population-weighted average instrument and changing the initial shares, the contemporaneous shifts, or adjusting the shifts to be age-specific. The benchmark measure used in the main analysis considers the number of employed in each industry, weighted by the industry employment share in 2003. Panel C.2(a) illustrates what happens when using different shifts, specifically comparing the index based on employment level versus value-added. Notably,

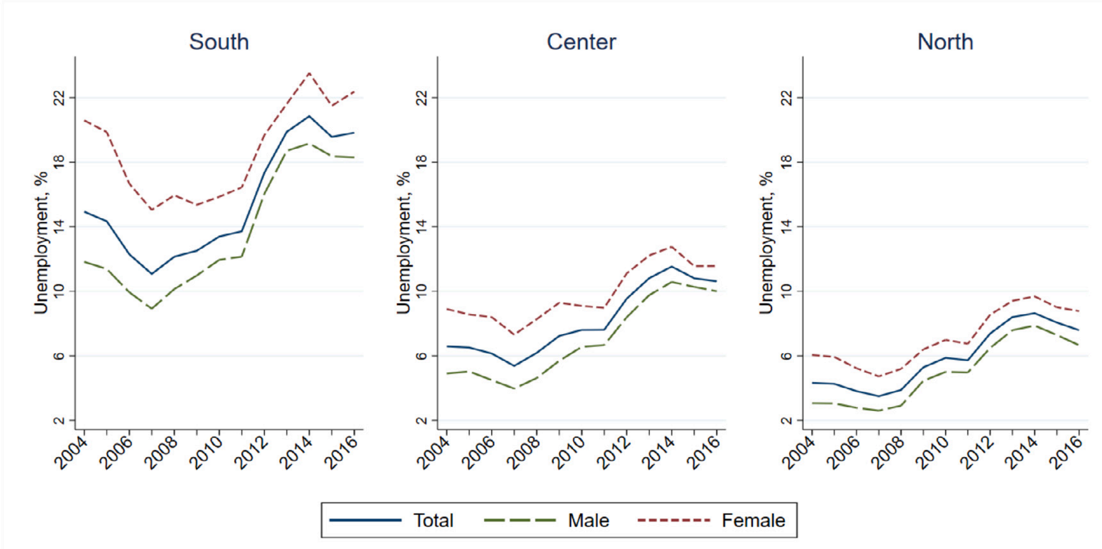

Fig. D.1. Unemployment by macro-areas and gender.
Notes: The figure plots the unemployment rate for the average (population weighted) province in the Center-North and South macro-areas.

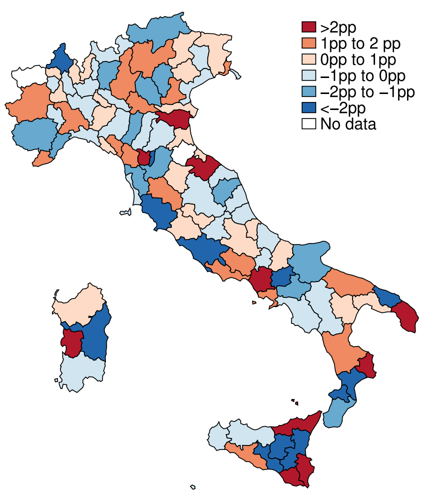

Fig. D.2. Changes in unemployment.
Notes: The map presents the change in the unemployment rate before and after the recession. I regress the unemployment rate on region-specific time trends and calculate the average residuals for the pre-recession period 2004–2007 and the post-recession period 2015–2016. I then subtract the pre-recession average from the post-recession average.

the instrument based on value-added rebounds quickly after the crisis, while employment remains depressed. A quick comparison with Fig. 2(b) suggests that the value-added-based index is less reflective of the extent of the crisis in the labor market. Moreover, value-added data is only available at the national level, so the resulting instrument cannot feature a leave-one-out correction. Panel C.2(b) compares alternative instruments by changing the initial provincial shares, both in terms of content and base year, which results in a simple rescaling of the instrument. I consider as benchmark share the province-industry employment share, as in Schaller (2016), and alternatively the ratio of local industry employment to the working-age population, as in Brown et al. (2018).[26]

Finally, Panel C.2(c) illustrates the age-adjusted instruments, where I adjust the employment shifts by the contemporaneous employment share of each age group at the national level, respectively. By construction, these adjusted instruments are going to reflect the evolution over time of the employment share of each age group, which between 2004 and 2016 increased for individuals between 35 and 65 years old and correspondingly decreased for the younger groups. The figure thus plots the deviations of the age-specific instruments from a linear time trend, which evolve quite similarly across groups and are the relevant source of variation for this analysis given the benchmark specification of Eq. (2).

Note that the benchmark instrument is computed using supply-side employment data on workers participating in the production process in each industry province, thus not accounting for residents who work outside of the province while including non-residents who work in the province. Data from the Labor Force Surveys indicates that in the period analyzed only 9% of workers were employed out of their province of residence, thus reassuring us of the limited nature of this measurement error. In any case, this does not threaten the analysis as long as the unobserved cross-province employment migration patterns are not correlated with childbearing or abortion. Finally, the results are robust to using an alternative instrument that measures employment by the province of residence extracted from the Labor Force Survey, as reported in Table C.3.

Table C.3 reports on how estimates vary when employing different versions of the Bartik instrument. The first column reports the estimates using the benchmark instrument, which combines the industry-specific local employment shares in 2003 with leave-one-out national employment, while columns 2 and 3 report the estimates using different shares to construct the instrument, respectively local employment shares in 2000 and the share of working age population in 2003; these are the same alternative instruments plotted in Fig. C.2(b). Column 4 employs a leave-one-out Bartik instrument measured using industry real value added, weighting industries by their employment share in 2003. Columns 5 to 6 report alternative versions of the instrument, changing both the shifts and shares. Specifically, they use the employment share in 2004 for industries according to the 2002 ATECO classification (2

[26] Unfortunately, data on the working-age population is not available at the province level for the year 2000. Moreover, the ratio of local industry employment to the working-age population by definition does not sum up to 1 within a province. This motivates the choice of the employment share as the benchmark in the main analysis.

Table E.1
Data appendix.

Variable	Source	Method
General fertility rate; population and live births	ISTAT, General Register Office	The GFR measures the number of live births for 1000 women in their childbearing age (15-49), considering only women of Italian nationality.
Abortion rate	ISTAT , Rilevazione sulle Interruzioni Volontarie di Gravidanza (Laboratorio ADELE)	The abortion rate measures the number of voluntary abortions for 1000 women in their childbearing age (15-49), considering only women of Italian nationality.
Abortion ratio	ISTAT, Rilevazione sulle Interruzioni Volontarie di Gravidanza (Laboratorio ADELE)	The abortion ratio measures the share of voluntary abortions over the number of estimated pregnancies in childbearing age, i.e. the sum of voluntary abortions and live births, considering only women of Italian nationality.
Miscarriages	ISTAT, Indagine sulle dimesse dagli istituti di cura per aborto spontaneo	Number of spontaneous abortions by province of abortion, 2010–2016.
Unemployment rate	ISTAT, Rilevazione sulle Forze di Lavoro	Computed using employment and unemployment counts from the LFS
Share of irregular workers	ISTAT, European System of Accounts	Share of employed individuals with irregular work per region, 2004–2016.
Referendum 1981	Ministry of Internal Affairs, Historical Archive of Elections	Share of yes votes to question 5 of the 17/05/1981 Referendum.

Table E.2
Bartik instrument, data and formulas.

Data and source	Method	Formula
Employment Bartik Employment by industry and province of work - ISTAT, Regional accounts	I take a weighted average of the national-level number of employed individuals added in each of the 10 sectors considered (see Table C.1), where the weights are the local employment shares of each industry in 2003. The employment data measures the number of workers by place of work, i.e. it measures the workers that participate in the production process in each province: it does not account for residents that work outside of the province, while it includes non-residents that work in a firm located in the province. An alternative measure uses the number of employed by industry weighted by the share of working age population employed in each industry and province in 2003.	$B_{p,t} = \sum_{k=1}^{K} \chi_{p,k,t_0} E_{-p,k,t}$ $\chi_{p,k,t_0} = \frac{E_{p,k,t_0}}{E_{p,t_0}}$ Alternative weight: $\chi_{wp,p,k,t_0} = \frac{E_{p,k,t_0}}{Pop_{15-65,p,t_0}}$
Value Added Bartik VA by industry, Italy-ISTAT	I take a weighted average of the national-level real value added in each of the 10 sectors considered (see Table C.1), where the weights are the local employment shares of each industry in 2003. Real value added by industry is measured in thousands of 2004 Euros.	$B_{AV,p,t} = \sum_{k=1}^{K} \chi_{p,k,t_0} VA_{IT,k,t}$
Age-specific Bartik	For the age-specific Bartik instrument, I rescale the standard Bartik instrument by the share of employed of each age class at the national level over the total number of employed.	$B_{a,p,t} = \psi_{a,IT,t} \sum_{k=1}^{K} \chi_{p,k,t_0} E_{-p,k,t}$ $\psi_{a,IT,t} = \frac{E_{a,IT,t}}{E_{IT,t}}$

digit level), thus increasing the number of industries considered to around 60, and the leave-one-out number of employed and unemployed as shifts, respectively. Overall the sign of the second-stage estimates remains consistent when employing alternative versions of the instrument, though the is some variability in the size of the estimates: the effect on the fertility rate ranges between −0.4 and −1.1 of a standard deviation; the estimate on the abortion rate ranges between 0.1 and 0.3 s.d.; the estimate on the abortion rate ranges between 0.3 and 0.4 standard deviations.

Appendix D. Geographic variation in unemployment

This section describes the geographical differences in the Italian labor market in terms of unemployment and the effect of the economic crisis in different areas.

Fig. D.1 illustrates both the geographic variation and evolution over time of unemployment rate, respectively for the average provinces of Southern, Central, and Northern Italy. Two patterns emerge from this graph: first, the female unemployment rate is consistently higher than that of males regardless of geographic location; second, the average unemployment rate becomes smaller as we move towards the North. In particular, the labor market of Southern provinces is characterized by a higher unemployment rate, partially due to a higher prevalence of inactivity, irregular contracts, and low female labor force participation (De Philippis et al., 2022). Moreover, unemployment rates in the Southern provinces seem to respond less strongly to the Great Recession in 2008–2010, while in the Center and North they increase in correspondence to both crises. This can be partially traced back to the different industrial composition of these macro-areas, as the export-oriented manufacturing industry that was greatly impacted by the Great Recession concentrates in Northern provinces, as illustrated by Fig. C.1. Moreover, the response of labor supply to the recession was also geographically heterogeneous: in Southern Italy, the increase in the unemployment rate during the first recession was partly offset by increased inactivity, while in the Central and Northern regions most of the reduction in employment translated directly into a rise of unemployment (Consiglio Nazionale dell'Economia e del Lavoro, 2011).

Fig. D.2 maps the before-after change in unemployment rates across provinces. Although Northern provinces show a slight prevalence of increased unemployment rates, the map does not reveal a clear North-South divide in terms of the labor market effects of the crises. Therefore, while different areas experienced impacts at different times, the total variation in unemployment rates was dispersed across the country.

Appendix E. Data appendix

See Tables E.1 and E.2.

References

Adao, R., Kolesár, M., Morales, E., 2019. Shift-share designs: Theory and inference. Q. J. Econ. 134 (4), 1949–2010.

Aksoy, C.G., 2016. The effects of unemployment on fertility: evidence from England. BE J. Econ. Anal. Policy 16 (2), 1123–1146.

Ananat, E.O., Hungerman, D.M., 2012. The power of the pill for the next generation: oral contraception's effects on fertility, abortion, and maternal and child characteristics. Rev. Econ. Stat. 94 (1), 37–51.

Anderberg, D., Rainer, H., Wadsworth, J., Wilson, T., 2016. Unemployment and domestic violence: Theory and evidence. Econ. J. 126 (597), 1947–1979.

Autor, D.H., Duggan, M.G., 2003. The rise in the disability rolls and the decline in unemployment. Q. J. Econ. 118 (1), 157–206.

Autorino, T., Mattioli, F., Mencarini, L., 2020. The impact of gynecologists' conscientious objection on abortion access. Soc. Sci. Res. 87, 102403.

Bárdits, A., Adamecz-Völgyi, A., Bisztray, M., Weber, A., Szabo, A., 2023. Precautionary Fertility: Conceptions, Births, and Abortions around Employment Shocks. IZA Discussion Paper.

Bartik, T.J., 1991. Who Benefits from State and Local Economic Development Policies?. W.E. Upjohn Institute for Employment Research, Kalamazoo, Michigan.

Bastianelli, C., Farris, M., Benagiano, G., 2005. Reasons for requesting emergency contraception: a survey of 506 Italian women. Eur. J. Contracept. Reproduct. Health Care 10 (3), 157–163.

Bastianelli, C., Rosato, E., Farris, M., Benagiano, G., 2016. Emergency contraception: A survey of 1773 women. Eur. J. Contracept. Reproduct. Health Care 21 (6), 455–461.

Becker, G.S., 1991. A Treatise on the Family. Harvard University Press.

Bhalotra, S., Britto, D., Pinotti, P., Sampaio, B., 2021. Job displacement, unemployment benefits and domestic violence. IZA Discussion Paper, No. 14543.

Bitler, M., Zavodny, M., 2001. The effect of abortion restrictions on the timing of abortions. J. Health Econ. 20 (6), 1011–1032.

Blank, R.M., George, C.C., London, R.A., 1996. State abortion rates the impact of policies, providers, politics, demographics, and economic environment. J. Health Econ. 15 (5), 513–553.

Bloom, D.E., Canning, D., Fink, G., Finlay, J.E., 2009. Fertility, female labor force participation, and the demographic dividend. J. Econ. Growth 14 (2), 79–101.

Borusyak, K., Hull, P., Jaravel, X., 2022. Quasi-experimental shift-share research designs. Rev. Econom. Stud. 89 (1), 181–213.

Bound, J., Holzer, H.J., 2000. Demand shifts, population adjustments, and labor market outcomes during the 1980s. J. Labor Econ. 18 (1), 20–54.

Brown, D., De Cao, E., et al., 2018. The impact of unemployment on child maltreatment in the United States. Institute for social and economic research.

Bruckner, T.A., Mortensen, L.H., Catalano, R.A., 2016. Spontaneous pregnancy loss in Denmark following economic downturns. Am. J. Epidemiol. 183 (8), 701–708.

Buckles, K., Hungerman, D., Lugauer, S., 2021. Is fertility a leading economic indicator? Econ. J. 131 (634), 541–565.

Comolli, C.L., 2017. The fertility response to the Great Recession in europe and the United States: Structural economic conditions and perceived economic uncertainty. Demogr. Res. 36 (1), 1549–1600.

Consiglio Nazionale dell'Economia e del Lavoro, 2011. Rapporto sul mercato del lavoro 2010–2011.

Crowley, J.E., Jagannathan, R., Falchettore, G., 2012. The effect of child support enforcement on abortion in the united states. Soc. Sci. Q. 93 (1), 152–172.

De Philippis, M., Locatelli, A., Papini, G., Torrini, R., 2022. Italian economic growth and the north-south gap: historical trends and future projections in light of the recent demographic scenarios. Bank of Italy Occasional Paper (683).

Dehejia, R., Lleras-Muney, A., 2004. Booms, busts, and babies' health. Q. J. Econ. 119 (3), 1091–1130.

Del Boca, D., Sauer, R.M., 2009. Life cycle employment and fertility across institutional environments. Eur. Econ. Rev. 53 (3), 274–292.

Del Bono, E., Weber, A., Winter-Ebmer, R., 2012. Clash of career and family: Fertility decisions after job displacement. J. Eur. Econom. Assoc. 10 (4), 659–683.

Drago, R., Sawyer, K., Shreffler, K.M., Warren, D., Wooden, M., 2011. Did Australia's baby bonus increase fertility intentions and births? Popul. Res. Policy Rev. 30 (3), 381–397.

Fahlén, S., Oláh, L.S., 2018. Economic uncertainty and first-birth intentions in Europe. Demographic Research 39 (October), 795–834.

Fischer, S., Royer, H., White, C., 2018. The impacts of reduced access to abortion and family planning services on abortions, births, and contraceptive purchases. J. Public Econ. 167, 43–68.

Goldsmith-Pinkham, P., Sorkin, I., Swift, H., 2020. Bartik instruments: What, when, why, and how. Amer. Econ. Rev. 110 (8), 2586–2624.

Goldstein, J.R., Cassidy, T., 2014. A cohort model of fertility postponement. Demography 51 (5), 1797–1819.

Goldstein, J., Karaman Örsal, D.D., Kreyenfeld, M., Jasilioniene, A., 2013. Fertility Reactions to the "Great Recession" in Europe. Demogr. Res. 29 (July), 85–104.

González, L., Trommlerová, S.K., 2023. Cash transfers and fertility: How the introduction and cancellation of a child benefit affected births and abortions. J. Hum. Resour. 58 (3), 783–818.

González-Val, R., Marcén, M., 2017. Divorce and the business cycle: a cross-country analysis. Rev. Econ. Household 15 (3), 879–904.

González-Val, R., Marcén, M., 2018. Unemployment, marriage and divorce. Appl. Econ. 50 (13), 1495–1508.

Haas-Wilson, D., 1996. The impact of state abortion restrictions on minors' demand for abortions. J. Hum. Resour. 140–158.

Happel, S.K., Hill, J.K., Low, S.A., 1984. An economic analysis of the timing of childbirth. Popul. Stud. 38 (2), 299–311.

Hofmann, B., Hohmeyer, K., 2013. Perceived economic uncertainty and fertility: Evidence from a labor market reform. J. Marriage Family 75 (2), 503–521.

Hotz, V.J., Klerman, J.A., Willis, R.J., 1997. The economics of fertility in developed countries. In: Handbook of population and family economics, Vol. 1. Elsevier, pp. 275–347.

Hotz, V.J., Miller, R.A., 1988. An empirical analysis of life cycle fertility and female labor supply. Econometrica 91–118.

Huttunen, K., Kellokumpu, J., 2016. The effect of job displacement on couples' fertility decisions. J. Labor Econ. 34 (2), 403–442.

Imai, K., Kim, I.S., 2021. On the use of two-way fixed effects regression models for causal inference with panel data. Political Anal. 29 (3), 405–415.

Istat, 2017. Natalità e fecondità della popolazione residente.

Jaeger, D.A., Ruist, J., Stuhler, J., 2018. Shift-share instruments and the impact of immigration. NBER working paper.

Kalist, D.E., 2004. Abortion and female labor force participation: Evidence prior to Roe v. Wade. J. Labor Res. 25 (3), 503–514.

Levine, P.B., Trainor, A.B., Zimmerman, D.J., 1996. The effect of medicaid abortion funding restrictions on abortions, pregnancies and births. J. Health Econ. 15 (5), 555–578.

Lima, J.M., Reeves, A., Billari, F., McKee, M., Stuckler, D., 2016. Austerity and abortion in the European Union. Eur. J. Public Health 26 (3), 518–519.

Lindo, J.M., Myers, C.K., Schlosser, A., Cunningham, S., 2020. How far is too far? New evidence on abortion clinic closures, access, and abortions. J. Hum. Resour. 55 (4), 1137–1160.

Loghi, M., Crialesi, R., 2017. La salute riproduttiva della donna. Lett. Stat. - Temi.

Malak, N., Rahman, M.M., Yip, T.A., 2019. Baby bonus, anyone? Examining heterogeneous responses to a pro-natalist policy. J. Popul. Econ. 32 (4), 1205–1246.

Marie, O., Zwiers, E., 2023. Religious Barriers to Birth Control Access. IZA discussion papers.

Matysiak, A., Sobotka, T., Vignoli, D., 2021. The Great Recession and fertility in Europe: A sub-national analysis. Eur. J. Popul. 37 (1), 29–64.

Medoff, M.H., 1997. A pooled time-series analysis of abortion demand. Popu. Res. Policy Rev. 16 (6), 597–605.

Medoff, M.H., 2007. Price, restrictions and abortion demand. J. Family Econ. Issues 28 (4), 583–599.

Ministero della Salute, 2016. Relazione del Ministro della Salute sulla attuazione della legge contenente norme per la tutela sociale della maternità per l'interruzione volontaria di gravidanza (Legge 194/78).

Ministero della Salute, 2018. Relazione del Ministro della Salute sulla attuazione della legge contenente norme in materia di procreazione medicalmente assistita (Legge 40/2004).

Modena, F., Rondinelli, C., Sabatini, F., 2014. Economic insecurity and fertility intentions: The case of Italy. Rev. Income Wealth 60 (S1), 233–255.

Modena, F., Sabatini, F., 2012. I would if I could: Precarious employment and childbearing intentions in Italy. Rev. Econ. Household 10 (1), 77–97.

Myers, C., Ladd, D., 2020. Did parental involvement laws grow teeth? The effects of state restrictions on minors' access to abortion. J. Health Econ. 71, 102302.

Pieroni, L., Roig, M.R., Salmasi, L., Turati, G., 2023. Legal Status and Voluntary Abortions by Immigrants. Dipartimento di Economia e Finanza, Università Cattolica del Sacro Cuore.

Schaller, J., 2013. For richer, if not for poorer? Marriage and divorce over the business cycle. J. Popul. Econ. 26 (3), 1007–1033.

Schaller, J., 2016. Booms, busts, and fertility testing the Becker model using gender-specific labor demand. J. Hum. Resour. 51 (1), 1–29.

Schneider, D., 2015. The great recession, fertility, and uncertainty: Evidence from the United States. J. Marriage Family 77 (5), 1144–1156.

Snarr, H.W., Edwards, J.A., 2009. Does income support increase abortions? Soc. Choice Welf. 33 (4), 575–599.

Sobotka, T., Skirbekk, V., Philipov, D., 2018. Economic Recession and Fertility in the Developed World. Popul. Develop. Rev. 37 (2), 267–306.

Spinelli, A., Forcella, E., Di Rollo, S., Grandolfo, M., 2006. L'interruzione volontaria di gravidanza tra le donne straniere in Italia. Rapporto ISTISAN 06/17, Istituto Superiore di Sanità.

Tur-Prats, A., 2019. Family types and intimate partner violence: A historical perspective. Rev. Econ. Stat. 101 (5), 878–891.