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Abstract

Despite the growing incidence of cesarean deliveries (CDs), procedure costs and benefits continue to be controversially discussed. In this study, we identify the effects of CDs on subsequent fertility and maternal labor supply by exploiting the fact that obstetricians are less likely to undertake CDs on weekends and public holidays and have a greater incentive to perform them on Fridays and days preceding public holidays. To do so, we adopt high-quality administrative data from Austria. Women giving birth on different days of the week are pre-treatment observationally identical. Our instrumental variable estimates show that a non-planned CD at parity 0 decreases lifecycle fertility by almost 13.6 percent. This reduction in fertility translates into a temporary increase in maternal employment.

JEL Classification: I12, J13, J11, J22, J21.
Keywords: Cesarean delivery, cesarean section, fertility, female labor supply.

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1 Introduction

In recent decades, the incidence of cesarean deliveries (CDs)[1] has substantially increased in almost all OECD member countries. On average, CD rates (i. e., the percentage of all CDs per live births) increased from 14 percent in 1990 to 20 percent in 2000 and almost 28 percent in 2013 (OECD, 2015). By contrast, the recommended CD rate by the *World Health Organization* is between 10 and 15 percent (World Health Organization, 2015). It seems indisputable that this upward trend in CD rates can only be partly explained by changes in the incidence of medical indications.[2] The rapidly growing number of CDs raises concerns about the effects of non-medically indicated CDs.[3] A CD is more costly than a vaginal birth and in many OECD countries, the average cost of a CD is twice or more than that of a vaginal birth (Koechlin et al., 2010). The surge in the number of CDs poses a moral hazard problem in health insurance markets and places an unwarranted burden on healthcare systems (Gibbons et al., 2010). In addition, non-medically indicated CDs may exert direct detrimental effects on the mother or child. In this study, we focus on the far-reaching consequences of CDs on maternal outcomes. More specifically, we investigate to what extent non-medically indicated CDs affect subsequent fertility and maternal labor supply.

How do CDs affect subsequent fertility? By undergoing a CD, mothers are subject to the risk of reducing their subsequent fertility involuntarily or voluntarily. A CD could cause later physiological reproductive constraints. While surgical procedures have improved over time, CD mothers are still at a higher risk of stillbirth, and ectopic pregnancy due to placental abnormalities (O'Neill et al., 2013, 2014). A CD could also affect fertility through an infection at the wound site or pelvic adhesions (Murphy et al., 2002). On the other hand, CD mothers could voluntarily reduce fertility through different channels. More specifically, the first CD is an indication of subsequent ones and this is associated with a higher risk of complications in future pregnancies (Daltveit et al., 2008). Observational studies show that maternal morbidity progressively increases with the number of CDs (Silver et al., 2006). Second, the recovery period after birth may be longer for CD mothers than those who had a normal vaginal delivery. This may increase the psychological problems after childbirth and reduce the willingness to have another child (Declercq et al., 2006, Weisman et al., 2010). Third, tighter constraints caused by the first child could reduce a CD mother's fertility. A large number of observational studies highlight the negative associations between CD and child outcomes.[4] Assuming that this association captures a causal relationship, CD mothers may decide to spend more time and resources on their first born and forgo having another child.

[1] A CD is the delivery of a child through surgical incisions made to the mother's abdomen and uterus (*ICD-10 Procedure Coding System*: 10D00Z0).

[2] Possible complementary explanations are older women experiencing first-time motherhood, increased *in-vitro* fertilization (OECD, 2013), malpractice liability concerns (Currie and MacLeod, 2008), reductions in CD risk, doctors' and patients' scheduling convenience (OECD, 2013, Brown III, 1996), and changes in patient preferences (Sachs et al., 1999).

[3] There is clear consensus that CD improves maternal and child health outcomes when medically indicated. Among others, breech presentation; complications of labor such as fetal distress, cord prolapse, placenta praevia, and other complications; and pre-existing conditions (e.g., certain cases of HIV infection) are widely accepted conditions that require a CD (NICE, 2011).

[4] These studies relate CDs to respiratory problems (Thavagnanam et al., 2008), allergies (Koplin et al., 2008), obesity (Goldani et al., 2011), diabetes (Cardwell et al., 2008), and cognitive and impaired emotional development (Martins and Gaffan, 2000). A study by Almgren et al. (2014) evidenced that CD may cause negative epigenetic changes in babies' DNA. The evidence from available design-based approaches is more mixed. Jensen and Wüst (2015) find a positive effect by CDs in the case of breech positions, where the APGAR score increases and healthcare utilization reduces in the first year of life. In contrast, Costa-Ramón et al. (2018) find some negative effects on Apgar Scores, however, with no medium-term consequences. Recently, Card et al. (2018) find a more nuanced picture with

Negative association between CDs and subsequent fertility Both macro and micro data show a negative association between CD and subsequent fertility. Figure 1 depicts the correlation between CD rates and total fertility rates for 23 European countries. The negative correlation of -0.68 is significant at the 99.9 percent level. Likewise, the dominant finding in observational studies based on micro data is that fertility is delayed and reduced after a CD compared to a vaginal delivery (Gurol-Urganci et al., 2013, O'Neill et al., 2013). Although illustrative, this evidence — much like the above cross-country correlation — does not establish a causal relationship given the endogenous allocation to mode of delivery. A comparison of cesarean delivering mothers (henceforth, CD mothers) and non-CD mothers is most likely biased given the presence of factors promoting CD and affecting the outcomes under consideration.

The potential causal link between CD and fertility has received little attention in the economic literature. A notable exception is Norberg and Pantano (2016), who exploit several data sources and adopt various empirical strategies to explore the importance of maternal choices to identify the (negative) association between CD and subsequent fertility. Although the authors do not claim to identify a causal effect, they provide suggestive evidence for maternal responses as an important aspect of the link between CD and subsequent fertility: CD for an index pregnancy is associated with a lower incidence of future live births, higher self-reported infertility, miscarriage of future pregnancies, and higher overall rate of active contraception. Recently, Card et al. (2018) use an instrumental variables (IV) approach to identify the causal effect of CD for low risk first births in the US. They use the relative distance from a mother's home to hospitals with high and low CD rates as their source of exogenous variation. The identifying assumption is that this variable is not correlated with any unobserved determinants of their outcome variables.[5] The focus of their study is children's and maternal health outcomes, but they also examine the probability of a second birth up to 4 years after first birth. They find no clear effects on fertility, but conclude that their '*design is underpowered to detect modest sized fertility effects*' as they only have access to five years of data.

Our empirical analysis is based on high quality administrative data from Austria. We use the Austrian Birth Register linked with a employer-employee matched dataset. We observe all births between 1995 and 2007, along with information on mode of delivery, health indicators for newborns, and mother's socioeconomic information. We focus on nulliparous mothers and match information on their subsequent fertility and maternal labor supply up to 2013. This allows us to contribute in three ways to the literature on the causal effects of a CD on subsequent fertility. First, we provide meaningful estimates of the effect of CD on the probability of a second child up to six years after birth. Second, we are able to inform about the total fertility effects. In particular, we restrict the sample to births earlier than 2004 to estimate the effects up to ten years after the first birth. Third, we investigate whether the effects are heterogeneous by maternal education.

CDs and maternal labor supply A CD may influence maternal labor supply through potential effects on child health, maternal health or fertility. These channels work in opposite directions. First, mothers may reduce their labor supply to take care of their sicker children. Second, a

some positive and some negative effects for low risk first births, while Costa-Ramón et al. (2019) find avoidable CDs to increase only asthma from early childhood.

[5]In an early version of this paper, we have also experimented with using regional variation in CD rates as a source of exogenous variation. This alternative IV approach provides the same qualitative results, however, with a larger effect size. We concluded that this alternative approach is less credible in the Austrian setting (at least with the data at our hand) and excluded it from the paper.

longer physical and mental recovery may influence the time it takes for mothers to come back to work after delivery, and any long-lasting effects on maternal health may translate into lower participation and/or wages. Last, fertility reductions may translate into higher employment.

Bailey (2006) estimates the fertility and labor market effects of another technology — the introduction of the oral contraception pill — and finds that it reduced teenage pregnancy, increased female labor force participation and the number of working hours substantially. One could expect smaller positive effects from a CD if the procedure affects subsequent fertility. The existing causal evidence on the effects of a third child on employment (Angrist and Evans, 1998, Maurin and Moschion, 2009) suggests that the effects of subsequent children on maternal employment are relatively small and disappear in the long-run. Our paper then also contributes to the scarce literature on the effects of birth-related technologies on maternal labor market behavior.

Establishing a causal relationship We analyze whether a non-medically indicated CD affects subsequent fertility behavior and maternal labor supply. To establish a causal link, we exploit three facts. First, obstetricians have considerable control over whether a delivery is vaginal or cesarean. Second, hospitals face capacity shortages for surgical procedures on weekends and public holidays. Third, obstetricians have a preference for leisure on weekends. As a result, despite nature's almost uniform distribution of births across time, one finds differential likelihoods of CDs across days of the week. This holds true after excluding planned CDs, which are always scheduled on working days. There is a 5 percentage point lower likelihood of a CD on weekends and public holidays and a 1 percentage point higher likelihood on Fridays and working days preceding public holidays. We argue that the differential likelihood of a CD for expectant mothers admitted to a hospital on a certain weekday is driven by obstetricians' demand for leisure and not correlated with factors related to the mother or child. Accordingly, we exploit this variation in CDs across weekdays using an IV approach.

The first IV is the delivery on a weekend or public holiday and the second is delivery on a Friday or any other working day before a public holiday. Our identifying assumption is that the weekday of admission is not correlated with the unobserved determinants of subsequent fertility or labor supply. While this assumption cannot be fundamentally tested, it is reassuring that women admitted on different days of the week are observationally identical. Their unadjusted means of socioeconomic characteristics are indistinguishable, they have the same average labor market history, and equal health expenditure prior to birth. One potential concern regarding the exclusion restriction of the first (but not second) IV is the quality of medical services provided on weekends or public holidays. Due to lower staffing and availability of clinical services, medical quality could be lower. This could lead to the subsequent reduction of fertility. While there is little empirical evidence on reduced quality in neonatal units on weekends, such an effect would provide us with a more conservative result. We find very similar results for both IVs.

Preview of main results Our main result is that non-medically indicated CDs (triggered by hospitals' free surgical capacity or obstetricians' demand for leisure) reduce subsequent fertility permanently and increase maternal labor supply over a period of about six years. The probability of a second birth decreases by about 18.7 percentage points and the likelihood of a third birth by 6.2 percentage points after 10 years. This translates into a 13.6 percent reduction in lifecycle fertility. The resulting increase in maternal employment (over a period of 6 years after the first birth) generates an additional maternal labor income of about 17 percent. The estimated treatment effects on fertility are very similar by maternal age and education. By contrast, labor supply effects are driven by highly-educated mothers. All estimated treatment effects are highly statistically significant and similar across both IV approaches.

The remainder of this paper is organized as follows. Section 2 discusses our research design. Section 3 presents the estimation results and several sensitivity checks. Section 4 summarizes and concludes the paper.

2 Research design

In this section, we first discuss the institutional background and present our data along with descriptive statistics. Then, we argue that the supply determination of CDs provides a source of exogenous variation. Once we specify the source of supply determination, we exploit and discuss the assumption that allows for the identification of a causal CD effect on subsequent maternal behavior. Next, we present our econometric estimation strategy.

2.1 Institutional setting

Austria has a Bismarckian-type healthcare system that is predominantly funded by employment-based social insurance contributions. Patients hold mandatory health insurance administered through nine *regional health insurance funds* ("Gebietskrankenkassen"), which cover private employees and their dependents, and 16 social security institutions that provide health insurance for specific occupational groups such as farmers, civil servants, and self-employed persons.[6] More than 99 percent of the Austrian population is insured with free access to (almost all) healthcare services. Health insurance covers all expenditures associated with sickness and maternity for both in- and outpatient healthcare services. In particular, medical expenses during pregnancy and the costs of delivery are fully covered, irrespective of mode of delivery.[7]

Expectant mothers can freely choose a hospital for delivery, although most mothers select the nearest one with a maternity department, particularly when they expect a spontaneous delivery. CDs are available across Austrian hospitals and a planned CD can be scheduled in any preferred obstetrics department.[8] Mothers who have voluntary and private supplemental health insurance can also choose their treating obstetrician and midwives. In contrast to several (Anglo Saxon) healthcare systems, all treating doctors, midwives, and nursing staff are permanently employed by the chosen hospital. According to the duty hours of healthcare staff members, the largest capacity of hospital staff is available during weekdays. However, even if there is a shortage of staff during the weekends and public holidays, a complete surgical team must be available 24 hours a day and seven days a week in every obstetrics department. This infrastructure guarantees that unplanned CDs can be performed at all times.

[6] The *regional health insurance funds* cover approximately 75 percent of the Austrian population.

[7] Under the sickness funds, pregnant women are given access to a comprehensive mother–child care program introduced in 1974. This program comprises five basic prenatal health exams for the expectant mother and her unborn baby. Pregnant women receive these services free of charge. A strong financial incentive to participate in this program is the eligibility for several family (birth) benefits, which is tied in with the utilization of these prenatal medical examinations.

[8] The lack of quality rankings for Austrian hospitals indicates weak competition among clinics. As a result, potential quality differences between hospitals that offer CDs do not trigger much attention.

Typically, a team of doctors, midwives, and nursing staff works closely together in the maternity ward. Uncomplicated birth procedures are carried out by midwives and nursing staff, and the responsible obstetrician is regularly informed about the progress of the birth. If complications arise, the doctors (obstetrician, surgeon, and anaesthetist) decide if and when a CD is performed. The operation itself is then carried out by the obstetrician and/or surgeon, and the anaesthetist. Usually the personnel work in shifts. Therefore, midwives and obstetricians, who take over a case on admission to hospital, do not necessarily follow up on it until delivery. Understandably, the surgical team does not have strict shift changes. If they perform a CD (earlier), they may be able to end their shift on time and prevent the planned shift change from being exceeded.

2.2 Data and sample restrictions

Our empirical analysis is based on two administrative datasets from Austria. The first source of data is the *Austrian Birth Register* (ABR). This comprises all hospital and all home births which take place within Austria. All hospitals and midwives have a legal obligation to submit a specified form for each birth to *Statistik Austria*. The individual-level records contain information on birth characteristics such as date, place of birth, mode of delivery, gestation length, and birth weight. This information is complemented by maternal socioeconomic characteristics such as age, educational attainment, marital status, occupation, and religious denomination. Different births to the same mother can be linked by a unique mother identifier. Data from the ABR are available for 1971–2007, but information on mode of delivery is only available from 1995 onwards.

The second source of data is the *Austrian Social Security Database* (ASSD). These are administrative records to verify pension claims and structured as a matched employer–employee dataset. The database allows us to observe mothers' labor market status and the date of all live births in the period between 1973 and 2014. For each mother, we observe on a daily basis where she is employed, along with her occupation, experience, and tenure. Information on earnings is provided per year and per employer. The limitations of the data are top-coded wages and the lack of information on (contracted) working hours (Zweimüller et al., 2009).

The incidence of CDs substantially increased over the observation period. The CD rate increased from about 12 percent in 1995 to 27 percent in 2007 (see Appendix Figure C.1). Similar to other developed countries, there is no specific development in Austria that can explain the rise in CD rates by an increase in medical conditions. A more plausible explanation is that either the demand or supply for non-medically justified (or clinical gray area) CDs has increased over time. In our empirical analysis, we examine all first-time Austrian[9] mothers who had a single inpatient birth between 1995 and 2007 and analyze their subsequent fertility and labor market behavior until 2013. We focus on singleton deliveries because multiple deliveries (1.5 percent between 1995 and 2007) are not comparable in terms of subsequent fertility. We further exclude outpatient births since a CD always takes place in a hospital. Outpatient births are a rare phenomenon in Austria: between 1995 and 2007, the average annual rate of outpatient births among first-time mothers was a mere 1.2 percent. To abstract from planned CDs, we focus on births with a gestation length of 39 weeks or more.

[9]We focus on Austrian mothers because non-Austrian ones are more likely to leave the country or give birth abroad. Since the ABR only records births that took place in Austria, there is potential for a measurement error in non-Austrian mothers' subsequent fertility.

Table 1 provides descriptive statistics for a list of mother and child characteristics by treatment status. Column (1) refers to mothers with vaginal delivery and column (2) to CD mothers. We find significant differences: CD mothers are on average 1.7 years older, have somewhat higher educational qualifications, are more likely to be married (plus 2.2 percentage points), and are less likely to be Catholic (minus 1.9 percentage points). There are smaller differences in birth outcomes. The gestation age is identical and birth weights are comparable.[10] However, female newborns are substantially under-represented among CD mothers (46 percent). Table 2 compares outcome variables by treatment status. In terms of fertility, CD mothers are less likely to have a second (and third) child within a period of six years after their first birth. Six years after the first, CD mothers are more than 8 percentage points less likely to have a second child. CD mothers are more likely to be employed before the first birth and even more so after. We conclude that CD and non-CD mothers substantially differ before treatment.

2.3 Classification of reasons for CD

As discussed in the introduction, CDs are performed for various reasons. Some of these reasons are also correlated with subsequent maternal behavior. Thus, a naïve estimation of the effects of a CD for the first birth on subsequent behavior is most likely to give us biased estimates. For the sake of argument, we conceptually distinguish the following cases:

- *Medically indicated CD*: A CD is performed because of medical indications that impede a vaginal birth. An obstetrician is most likely to perform a CD if there are absolute medical indications that obligate him/her to do so. The patient is expected to follow the obstetrician's recommendation.

Other cases are influenced by either an obstetrician's subjective consideration or incentives and potentially by the patient's preferences. We distinguish between non-medically indicated and clinical gray area CDs:

- *Non-medically indicated CD*: A CD is performed without medical indication and is either recommended by the obstetrician or requested by the mother.
  - *Supply-determined CD*: The CD is performed on the obstetrician's recommendation. Obstetricians may act in their own (or their employers') interest and suggest a CD even though there is no medical indication. The potential motives are to increase revenue or avoid prolonged working hours.
  - *Demand-determined CD*: The CD is performed as per the patient's request. Women may prefer a CD because of scheduling convenience or to avoid anxiety and labor pain associated with vaginal birth.
- *Clinical gray area CD*: Whether a CD is necessary is not always obvious. Often obstetricians must choose between a CD and vaginal delivery using their subjective judgment and patients are most likely to follow their doctors' advice. We hypothesize that CDs in this category are overwhelmingly supply and not demand determined.

[10]These results are influenced by our sample construction (we excluded observations with a gestation length shorter than 39 weeks).

Medical indications that always require a CD may be either exogenous or endogenous. Women with pre-existing conditions (e.g., selected cases of HIV infections) must opt for a CD and their future fertility is most likely to differ from that of other women. Other medically indicated CDs may be driven by circumstances that are not correlated with future fertility (e.g., breech position). Non-medically indicated CDs may also be endogenous or exogenous. These CDs are endogenous if women requesting for one tend to have differing fertility behavior. Most importantly, in the absence of sorting, supply-determined CDs are by definition exogenous to the mother. The empirical relevance for such CD types is, however, difficult to assess because doctors have a strong incentive to conceal such unethical behavior.[11]

The incidence of non-medically supply-determined CD is expected to be higher when a doctor (or his/her employer) is increasingly driven by profit-maximizing behavior. Clinical gray area CDs (where we also expect the supply side to be the driving force) may result from a fully altruistic action solely guided by the Hippocratic Oath and/or profit-maximizing behavior. Thus, whether a woman undergoes a CD is determined by the doctor and/or hospital's characteristics. However, these supply-side characteristics are unlikely to affect women's fertility behavior, provided the women do not select themselves to some of these characteristics.

Obviously, one cannot directly distinguish between the different types of CDs in observational data. Nevertheless, we suggest two alternative IV estimation strategies to isolate variations in non-medically indicated and clinical gray area supply-determined CDs.

2.4 Quantitative importance of supply-determined CDs

There are different ways to assess the importance of supply-determined CDs. The most common empirical strategy is to examine the variation in the incidence of CDs across regions, hospitals, and providers while implicitly assuming that one can control for differences in medical indications and demand-side factors. The dominant finding in the literature is that practice patterns substantially vary and non-medically supply-determined factors explain the recent dynamics in CD rates.[12] These differences can be explained by variations in physicians' risk aversion, altruism, patience, and other individual attitudes as well in legal regulations or reimbursement for different types of deliveries.[13]

[11] In contrast, there is anecdotal evidence for non-medically demand-determined CDs. Austrian women have reported that their request for a CD was fulfilled without any discussion, even in the absence of a medical indication, whereas medical guidelines mandate that doctors offer women a planned CD only if they demand for one after a discussion and support (NICE, 2011).

[12] Evidence for substantial differences across regions, hospitals, and providers is available for the United States (Menacker and Hamilton, 2010), Canada (Hanley et al., 2010), the United Kingdom (Bragg et al., 2010), France (FHF, 2008), Switzerland (OFSP, 2013), and several Latin American countries (Belizán et al., 1999). For instance, Kozhimannil et al. (2013) report that among 593 US hospitals nationwide, CD rates for low-risk pregnancies varied from 2.4 to 36.5 percent in 2009. Epstein and Nicholson (2009) study the variation across physicians. Using physician data for the states of Florida and New York, they conclude that physicians' idiosyncratic treatment styles explain nearly 30 percent of the variation in CDs (after controlling for a large set of relevant patient characteristics).

[13] Gruber and Owings (1996), Grant (2009), Triunfo and Rossi (2009) provide theoretical and empirical evidence that changes in doctors' reimbursement for different modes of delivery influence treatment intensity. Dubay et al. (1999), Currie and MacLeod (2008) analyze the impact of a medico-legal framework on physicians' inclination to perform CDs and find that physicians are able to adjust their behavior to changes in the legal environment.

2.4.1 Variation in CDs across weekdays

We identify supply-determined CDs by exploiting variations in the incidence of CDs across weekdays. We focus on births of nulliparous Austrian mothers for 1995–2007 documented in the *Austrian Birth Register*. Panel A of Figure 2 presents the CD rates across weekdays. We make two striking observations. First, there are clearly less CDs on Saturdays and Sundays. On weekends, we observe a CD rate of 16.7 %, whereas the average rate on weekdays is 23.0 %. The most obvious explanation for this difference of more than 6 percentage points is that planned CDs are not scheduled on weekends. In addition, obstetricians are less likely to recommend (non-medically indicated or clinical gray area) CDs given that hospitals face a certain capacity constraint on weekends. We attempt to isolate the latter effect. The second striking feature is the higher CD rate on Fridays (24.1 %) compared to the other working days with an average rate of 22.8 %. This pattern was first noted by Brown III (1996), who analyzed data from US military hospitals in the early 1990s. Military obstetricians did not earn extra income for performing a CD. Brown III (1996) attributed the sharp increase in unplanned CDs on Fridays (between 3 pm and 9 pm) to obstetricians' demand for leisure on weekends.[14] This means that non-medically indicated (or clinical gray area) CDs are more likely to occur on a Friday than on another working day, since obstetricians prefer to end their shift on time.[15]

Planned CDs Our estimation strategy relies on (non-medically indicated or clinical gray area) supply-determined CDs. Therefore, we exclude planned CDs and skip observations with a gestation length shorter than 39 weeks. This sample restriction is rationalized by the medical advice of Austrian scholars to perform a planned CD in week 38 of gestation (Husslein and Langer, 2000). A similar recommendation was made by the Swiss Society for Gynaecology and Obstetrics (SGGG, 2015). In their guidelines, they advise doctors to perform a planned CD in gestation weeks 38 or 39. The corresponding recommendation of the German Society for Gynaecology and Obstetrics issued in 2006 (DGGG, 2010) also stipulates the termination of gestation week 37 (i. e., weeks 38 and 39) for a CD on request. Deviating from the guidelines of professional societies in the German-speaking countries, the American College of Obstetricians and Gynecologists stresses potential complications with respect to elective CDs and therefore states that CD on maternal request should not be performed before a gestational age of 39 weeks (Committee on Obstetric Practice, 2007).

In the Web Appendix A, we summarize a supplementary analysis to assess the compliance of obstetricians to perform planned CDs in gestation week 38. While we do not have information on whether a CD was planned or not, we exploit the fact that planned CDs are only done on weekdays and estimate the changing likelihood of a CD on weekdays (as compared to leisure days) over time and by gestational age. Our estimation results suggest that planned CDs are predominantly performed in weeks 37 and 38.[16] Moreover, as a robustness analysis, we provide fertility regressions for a sample that includes only observations with a gestation length of 40 weeks or more (see Section 3.2). The results are very similar to those based on the main sample.

[14] We cannot observe the time at which expectant mothers enter a hospital.

[15] A similar pattern is also found in data for California (Spetz et al., 2001), Spain (Costa-Ramón et al., 2018) and Finland(Costa-Ramón et al., 2019).

[16] Keeping births with a gestation of 39 weeks in the sample saves 59, 451 observations (more than 21 %). The sample restriction further eliminates all preterm births.

Panel B of Figure 2 shows the CD rate across weekdays for the reduced sample. Two findings are noticeable. First, compared to the full sample, CD rates are lower across all weekdays. Second, the relative difference in CD rates between weekends and weekdays decreases after excluding pregnancies shorter than 39 weeks. Both facts support our interpretation that the reduced sample comprises only unplanned CDs. We interpret the remaining differential between weekdays and weekends as a result of a reduced likelihood of supply-determined CDs, given the limited medical capacity during weekends. The difference between Fridays and other working days remains approximately the same in the reduced sample and may result from the obstetricians' demand for leisure on weekends. Panel C of Figure 2 depicts the CD rates across weekdays for the sample including only observations with a gestation length shorter than 39 weeks. The vast majority of planned CDs should be included in this sample. This assumption is supported by the fact that the CD rates in this sample are substantially higher and the weekend-weekday gradient increases in comparison to the full sample in Panel A.

Public holidays Given our explanations for an increased CD rate on Fridays (obstetricians' demand for leisure on weekends) and reduced CD rates on weekends (capacity constraints), we expect to find identical patterns among working days, other than Friday, that precede a public holiday and public holidays occurring during the week.[17] In Panel D of Figure 2, we distinguish between (i) working days preceding another working day (ii) Fridays and working days preceding a public holiday (pre-leisure day, PLday), and (iii) Saturdays, Sundays, and public holidays (leisure day, LDay). Panels E and F of Figure 2 show the distribution of CD rates on the basis of this new division of days for the samples with a gestation length of at least 39 weeks and less than 39 weeks. The figures support our interpretation that extra CDs before non-working days are caused by the obstetricians' demand for leisure and as such, are entirely supply determined.

Next, we present descriptive statistics. Table 3 compares mother and child characteristics for different birth days (leisure day, pre-leisure day, and other day). We find no difference across columns in the sex ratio, gestation length, or birth weight. Only the CD rate varies by group. The comparison of mother's age, occupation, educational attainment, family status, religious denomination, and province of residence does not show important differences across columns either. Finally, a mother's pre-treatment employment rates are identical across groups. According to these statistics, mothers' average pre-treatment socioeconomic characteristics are observationally identical, irrespective of whether the birth took place on a leisure day, pre-leisure day, or any other working day.

Health prior to first birth: It would be further instructive to observe maternal health prior to first birth. For the sub-sample of women from one Austrian federal state, we have access to detailed information on healthcare spending. The database of the Upper Austrian health insurance fund ("*Oberösterreichische Gebietskrankenkasse*") comprises information on healthcare expenditure and hospitalization for all private employees and their dependents in Upper Austria (see Section 2.1). We observe 9,802 first time mothers, who gave birth in an Upper Austrian hospital between 2005 and 2007. The upper panel of Table 4 depicts average quarterly outpatient expenditure for medical attendance and medical drugs for different birth days for this group of mothers. The differences in the expenditure categories across birthdays are minor. The only notable pattern is across time

[17] Austria has 13 nation-wide public holidays per year. Our data covers a period of 13 years or 4,748 days, of which 156 days (3.3 percent) were public holidays that fell on a weekday and 86 days (1.8 percent) preceded a public holiday that did not occur on a Friday, Saturday, Sunday, or public holiday.

(and not across groups). Outpatient expenditure increase prior to birth (compare quarters $1-4$ with quarters $5-8$), which results from (preventive) healthcare services during pregnancy. The lower panel of Table 4 shows information on the inpatient sector. These figures refer to births in 2006 and 2007 only, since inpatient information is not available before 2005. Again, we do not find any important differences across birth days. Thus, based on the subsample of Upper Austrian mothers, we hypothesize that the mothers who give birth on different days of the week are observationally identical, even in terms of their pre-treatment health.[18]

Fertility preferences: Ultimately, we are concerned that women giving birth on different week days have different characteristics that are correlated with fertility, such as fertility preferences. In principle, it is possible that women with different fertility preferences are observationally identical in terms of their socio-economic characteristics, health prior to first birth, and birth outcomes at parity 0. In this case, our balancing checks provided in Tables 3 and 4 would not be informative. While we have no information on fertility preferences in our estimation sample, we refer to related data. The first wave of the Austrian *Generations and Gender Survey* comprises detailed information on female respondents' socio-economic characteristics and several questions on their fertility preferences. In particular, women are asked whether they '*[...] intend to have a/another child during the next three years'*. Using the sub-sample of female respondents with zero or one biological child, we compare the average characteristics of women with and without plans for a/another child in Table 5. To the extent possible, we have selected the characteristics in line with those presented in Tables 1 and 3 for our estimation sample. It turns out that women with a stated preference for higher fertility look different as compared to women with a stated preference for lower fertility. The former have less likely compulsory schooling only, and they have more often a university degree. Women who want more kids are more often married and catholic. We even see a different distribution regarding their province of residence. Finally, we find that women who plan to have more children are substantially healthier. All these findings suggest that women with different stated fertility preferences differ in observable characteristics. The absence of such differences in our balance tables supports the idea that our IV is not correlated with unobserved fertility preferences in our estimation sample. Although the available data does not allow for a more direct test of differences in fertility preferences, this suggestive evidence is reassuring.

2.5 Estimation strategy

We analyze the effects of mode of delivery for the first birth on subsequent maternal behavior. In particular, we aim to determine whether mother i's CD (compared to a vaginal birth) has a causal impact on her fertility and labor supply. Accordingly, we define a binary variable $CD_{i(t=0)}$ equal to one if the mode of delivery for the first birth (in year $t=0$) is a CD and zero otherwise. Fertility is captured by the binary variable $F_{i(t+p)}$, which is equal to one if the mother gives birth

[18] We also estimate models in which we regress each observable characteristic on whether the birth took place on a leisure day or a pre-leisure day with and without controls (see Appendix Tables C.1 and C.2 for details). Overall, we find very few statistically significant coefficients. The significant coefficients for birth weight and gestation length indicate somewhat later births on leisure days. These variables are potentially affected by a CD and thus represent bad controls. Therefore, we exclude them from our main regressions. Appendix Tables C.23 and C.24 show that the results are robust to including birth weight and gestation length as additional control variables.

to a another child in or before year $t+p$ with $p \subseteq \{1,2,3,4,5,6\}$. Our outcome equation is defined as

$$F_{i(t+p)} = \beta \cdot CD_{i(t=0)} + x'_{i(t=0)} \cdot \gamma + \epsilon_{i(t+p)}, \tag{1}$$

where $x_{i(t=0)}$ denotes a vector of covariates comprising information on the mother's age, marital status, educational attainment, occupation, religious denomination, child's sex, month of birth, year of birth, and province of residence.[19] Finally, $\epsilon_{i(t+p)}$ denotes the error term. To study a mother's post-birth labor supply, we define a binary indicator $E_{i(t+p)}$, which is equal to one if mother i is employed in year $t+p$, and estimate an equation analogous to (1).

A probit estimation of equation (1) is likely to provide a biased estimate for β because opting for a CD cannot be expected to be exogenous. The cause for a medically indicated CD — for instance, mother's health condition or complications during birth or pregnancy — may also influence subsequent maternal behavior regarding fertility and labor supply. For demand-determined CDs, it is *a priori* unclear whether (and if yes, how) maternal preferences for a CD are related to subsequent fertility behavior. To overcome the potential endogeneity problem, we suggest two IVs. Both exploit variation in the likelihood of a (non-medically indicated or clinical gray area) supply-determined CD across weekdays and public holidays, as discussed above.

2.5.1 Leisure day IV

First, we define a binary indicator $leisure_{i(t=0)}$ equal to one if the first birth took place on a Saturday, Sunday, or public holiday. In estimating

$$CD_{i(t=0)} = \tau^l \cdot leisure_{i(t=0)} + x'_{i(t=0)} \cdot \gamma^l + \varepsilon^l_{i(t=0)}, \tag{2}$$

we expect to obtain a negative estimate of τ^l since obstetricians are less likely to initiate supply-determined CDs on these days owing to capacity constraints. The identifying assumption of this approach is that $leisure_{i(t=0)}$ is not correlated with $\epsilon_{i(t+p)}$.

It is reassuring that mothers who give birth on different weekdays are observationally identical in terms of birth outcomes and other maternal characteristics. We see one potential concern regarding the exclusion restriction. Ideally, the quality of medical care on weekends and public holidays is comparable to that provided on weekdays.[20] However, a reduced number of physicians on duty may result in lower quality of care. There are several arguments to be noted here. First, we believe that the organizational structure of Austrian hospitals — they must always provide sufficient resources to perform a surgery such as an emergency CD — guarantees constant quality across weekdays. However, we are unable to substantiate this argument with empirical evidence. Second, a quality difference *per se* does not impose a threat on our identification, provided it does

[19] We checked the robustness of our estimation results to the inclusion of the variables gestation length and birth weight (see Section 3.4). We do not include these two variables in our baseline specification, since these two variables are potentially affected by mode of delivery.

[20] Several medical science studies highlight that patients admitted to hospitals on weekdays have better health outcomes than those admitted on weekends (see, for instance, Bell and Redelmeier, 2001). The source of this so-called "weekend effect", however, is unclear. The literature discusses healthcare system factors (such as lower weekend staffing and availability of clinical services) and patient factors (e.g., patients admitted on weekends could be more complex). For instance, Dobkin (2003) shows that after correcting for patients' observable characteristics, differences in outcomes vanish. Notably, there is little evidence for a weekend effect on births. For instance, Gould et al. (2003) report that after adjusting for birth weight, no weekend effect was found on infant mortality.

not affect subsequent fertility. Third, since mothers typically stay several days in the hospital after birth, they receive the same postnatal care quality on average, irrespective of the weekday of birth.[21] Finally, any bias would provide us with a more conservative estimate. If the quality of care was indeed lower on weekends, then one could assume that women giving birth on weekends have reduced fertility. In contrast, we propose that CDs (which are *less* likely to occur on a weekend) lead to lower future fertility.

2.5.2 Pre-leisure day IV

Our second IV allows us to abstract from the latter complications. We define a binary indicator $preleisure_{i(t=0)}$ equal to one if the first birth took place on a Friday or any other workday preceding a public holiday. In the resulting equation

$$CD_{i(t=0)} = \tau^p \cdot preleisure_{i(t=0)} + x'_{i(t=0)} \cdot \gamma^p + \varepsilon^p_{i(t=0)}, \tag{3}$$

we expect a positive estimate of τ^p since obstetricians have a preference for leisure on weekends and public holidays. Put differently, we expect physicians to be more inclined to initiate a supply-determined CD the day before a weekend or public holiday so that they can leave the hospital on time. An alternative interpretation is that obstetricians carry out more supply-determined CDs on these days to avoid a potential CD on capacity-constrained weekends or public holidays. In either case, the obstetricians might try to initiate a vaginal birth by labor induction, but in many cases this procedure does not take and it leads to a CD. In our estimation model, in equations (1) and (3), we include $leisure_{i(t=0)}$ in both equations as a covariate. This means, identification comes only from $preleisure_{i(t=0)}$.

The credibility of this IV strategy is supported by three facts: first, women who give birth on a Friday or any other workday preceding a public holiday are observationally identical to all other women. This can be seen in Tables 3 and 4 by comparing the respective column (2) with columns (1) and (3). Second, we can safely assume that a consistent quality of medical care is provided on pre-leisure days, since staffing is identical to workdays preceding a workday. Third, our assumption that the focus on births with a gestation age of 39 weeks or higher excludes planned CDs is less crucial in the case of this alternative IV. Even if planned CDs remain in our estimation sample, these cases should be equally distributed across pre-leisure days and workdays preceding a workday. In sum, there is no obvious reason why $preleisure_{i(t=0)}$ should be correlated with $\epsilon_{i(t+p)}$.

We provide additional descriptive evidence based on a supplementary data source from the biggest maternity ward in Austria located at the Kepler University Hospital in Linz (Maternity Ward Data). These data cover duty rosters of obstetricians and comprehensive information about births including time of delivery. Appendix B includes a short description of the data, figures on admission times and the distribution of planned CDs over weekdays, and estimation results on the availability of medical doctors across different days of the week.

Appendix Figure B.1 shows that the pattern of admission times is similar for pre-leisure day and Monday-Thursday births, whereas the distribution of admissions for weekend births clearly de-

[21] For instance, let us assume an average hospital stay of five days. Depending on the weekday of admission, a hospital stay is 3/5 (We, Th, Fr, Sa), 4/5 (Tu, Su), or 5/5 (Mo) working days. This share is on average 3.5/5 for women admitted on the weekend and 3.6/5 for women admitted during the week. Thus, the only source of concern is the quality on the exact day of birth.

viates. This is another indication that pre-leisure and Monday-Thursday births are pre-treatment similar.[22] Appendix Table B.1 reports regression results for the number of available physicians per birth. In line with expectations, the results indicate a significantly lower availability of physicians on weekends (Leisure day). The medical infrastructure on weekends is approximately 44 % lower than during normal workdays. This reflects the postulated mechanism behind our leisure day instrument. Equally important, estimation results do not reveal any difference in the availability of physicians between a pre-leisure day and a normal workday. The coefficient for Pre-leisure Day remains insignificant and quantitatively small. The identical staffing on pre-leisure days and normal working days validates our pre-leisure instrument the mechanism of which rests on doctors' leisure preferences.

Finally, Appendix Figure B.2 depicts the distribution of planned CDs across weekdays. It can be seen that (almost) all scheduled CDs are performed during weekdays. The average number of planned CDs is approximately 0.7 per working day with slightly higher figures on Thursdays and pre-leisure days. The supplementary evidence based on the Maternity Ward Data strengthens the credibility of our instruments.

2.5.3 Estimation method

Since our treatment variable and all our outcome variables are binary, we use a *bivariate probit model* (BPM) to estimate equations (1), (2), and (3). In contrast to the linear IV model (implemented via 2SLS), the BPM explicitly accounts for the binary structure of the dependent variables (Chiburis et al., 2012). It assumes that the outcome and treatment variable are each determined by latent linear index models with jointly normal error terms with zero means and unit variances.

$$\begin{pmatrix} \epsilon_{i(t+p)} \\ \varepsilon_{i(t=0)} \end{pmatrix} \sim BVN(0, 0, 1, 1, \rho) \tag{4}$$

The additional parameter ρ measures the covariance of the error terms. An insignificant estimate of ρ would suggest that a simple probit model of equation (1) suffices to estimate causal effects. The functional form assumption identifies the model even in absence of an exclusion restriction. However, we include IVs in equations (2) and (3) for identification.

We are interested in *average marginal treatment effects* (ATEs). The estimator for the ATE is

$$\widehat{ATE} = \frac{1}{n} \sum_{i=1}^{n} \left[\Phi(x'_{i(t=0)} \cdot \hat{\gamma} + \hat{\beta}) - \Phi(x'_{i(t=0)} \cdot \hat{\gamma}) \right] \tag{5}$$

where Φ denotes the standard normal cumulative distribution function. Standard errors are clustered by district of residence and derived using the delta method.

A 2SLS model would be the natural alternative to the BPM. Chiburis et al. (2012) compare these two methods analytically and by means of simulation. They find that the BPM estimator is '*more efficient than [a 2SLS], especially when the model specification includes additional covari-*

[22] The data further allow the calculation of waiting times as the difference between time of birth and time of hospital admission at the individual level. The differences in waiting times across non-planned pre-leisure day, leisure day, and Monday-Thursday births of nulliparous mothers are minor. The median waiting time for the three groups is 11.8, 11.7, and 11.2 hours, respectively. Neither the pre-leisure day nor the leisure day mothers seem to be differently treated in terms of labor induction or delay of delivery.

ates.' Furthermore, they argue that '*the [2SLS] standard errors are often too large for meaningful hypothesis testing, especially when [the treatment probability] is close to 0 or 1.*'

The 2SLS model does not rely on the joint normality assumption (4) and the resulting estimates are *local average treatment effects*. Nevertheless, our 2SLS estimates (discussed in subsection 3.4) are remarkably similar to the BPM estimates.

3 Estimation results

This section presents our main empirical findings. We begin by summarizing the estimated effect of the weekday of birth on the likelihood of a CD (Section 3.1). Then, we present the estimated effect of a CD on subsequent fertility. Here, we try to distinguish between a quantum and tempo effect (Section 3.2). The effects of a CD at parity 0 on maternal labor market outcomes are presented in Section 3.3. Finally, in the robustness section (3.4), we report 2SLS estimation results, include potentially endogenous covariates, and control for supply-side characteristics.

3.1 Effects of weekday of birth on likelihood of CDs

Table 6 presents the results of our BPM on the basis of the two IV approaches. Panel A focuses on the leisure IV and summarizes output from the joint estimation of equations (1) and (2). Panel B covers the pre-leisure IV and lists output from the joint estimation of equations (1) and (3). The different columns refer to a series of separate estimations and show the estimated effect of a CD at parity 0 on the likelihood of a second birth 1–6 years after the first birth. This and all other tables report estimated average marginal effects with standard errors in parentheses.

First, we focus on the estimated effects of the weekday (i.e., leisure and pre-leisure days) on the likelihood of a CD. The second row of Panel A shows that the likelihood of a CD is 5 percentage points lower on the weekend or a public holiday than a working day. This effect is equivalent to a reduction of about 22 percent. We attribute the reduction to the obstetricians' reduced possibilities to induce non-medically indicated and/or clinical gray area CDs on these days when hospitals face tighter capacity constraints. The second row of Panel B shows that the likelihood of a CD is higher on Fridays or working days preceding a public holiday. The estimated effect amounts to 0.81 percentage points or about 3.6 percent. We attribute this increase in obstetricians' demand for leisure to the onset of a weekend or public holiday. This preference of doctors increases their incentive to induce CDs on these days, which allows them to end their shift on time. The Maternity Ward Data (see Appendix B) support this argument. Appendix Figure B.3 compares the time of birth between acute CDs performed on a pre-leisure day or workday (Monday-Thursday). As compared to a normal workday, we observe a pronounced clustering of CD births in the early afternoon between 12 pm and 4 pm on a pre-leisure day. This descriptive evidence demonstrates the mechanism of our pre-leisure instrument: doctors perform more CDs during early afternoon on Fridays and days before a holiday.

The leisure IV has comparably more power than the pre-leisure IV. The former IV has a F-statistic of 964.88 and the latter amounts to 16.4. For BPM, no study provides critical values regarding weak identification. To assess the strength of our IVs, we estimate eqs. (2) and (3) using a linear probability model. Thus, we estimate the first stages of a 2SLS estimation, for which critical values are provided by (Stock and Yogo, 2005, Table 5.2, p. 101). In our case (with one endogenous variable and one IV), the critical Cragg-Donald Wald F-statistic is 16.38. We obtain a Cragg-Donald Wald F-statistic of 933.4 and 17.54 for the leisure and pre-leisure IV. However, this statistic requires an assumption of i.i.d. errors. In the presence of clustering and heteroskedasticity, the Kleibergen-Paap rk statistic is, therefore, additionally considered in practice. No study appears to exist that provides threshold values that the rk statistic should exceed for weak identification not to be considered a problem. Researchers usually use a value of 10 as an indication of a strong instrument, following the general proposal of Staiger and Stock (1997). We obtain a Kleibergen-Paap rk statistic of 621.0 and 11.7 for the leisure and pre-leisure IV. In summary, both IVs can be considered sufficiently strong according to either approach.

Covariates The estimated direct marginal effects of the covariates are in line with those in previous studies and consistent across specifications. Appendix Table C.3 provides the full estimation output for the model of column (VI) in Panel A of Table 6. The estimated birth year fixed effects reflect the increasing trend in CDs over time. A CD is *ceteris paribus* almost 10 percentage points more likely to occur in 2007 than in 1995. Notably, the estimated effect of giving birth on the weekend or a public holiday (leisure IV) is almost half the size of this secular time trend. The most important maternal characteristic is maternal age at birth. Mothers 36 years of age or older are almost 22 percentage points more likely to have a CD as compared to mothers below 19 years of age. A mother's educational attainment, occupation, family status, and religious denomination also have predictive power. For instance, the relationship between educational attainment and likelihood of a CD follows a U-shaped pattern. Up to the level of "post-secondary college", an increase in educational attainment is associated with the reduced likelihood of a CD (−4.4 percentage points). Beyond this level, we find a slightly positive impact of further education. Finally, CDs are less likely in the case of a girl child (−2.4 percentage points).

3.2 Effects of CDs on subsequent fertility behavior

For both IVs, we observe that a CD at parity 0 negatively affects the likelihood of giving birth to another child beginning three years after the first birth (see the first rows of Panels A and B in Table 6). The estimated effects are remarkably similar for the two alternative IVs. The quantitative effects can easily be compared across IVs and years in the graphical representation of the estimated marginal effects provided in Figure 3. Compared to a vaginal delivery, a CD reduces the likelihood of a second birth three years after the first birth by 7.4 and 6.4 percentage points. The estimated effects are statistically significant and persist over the whole period under consideration. Six years after the first birth, the likelihood of having a second child is reduced by 12.3 and 13.0 percentage points. This effect is equivalent to a 24 percent decrease.[23]

While we are confident that our sample restriction to births with a gestation age of 39 weeks or more excludes planned CDs (see the Web Appendix A), we are unable to provide a watertight statistical test. To further substantiate our claim, we provide two additional analyses. First, we replicate our estimation based on a sample that even excludes births with a gestation age of 39 weeks. Thus, we include only births with a gestation age of 40 weeks or more. It seems very unlikely that Austrian obstetricians schedule a CD that late. Based on this restricted sample, we obtain

[23]The outcome equation for the pre-leisure IV provides a falsification check of the leisure IV since it is included as a covariate. We do not find a significant direct effect of the leisure IV on fertility. The estimated marginal effect of −0.001 is economically and statistically not significant (standard error: 0.003). This supports the exclusion restriction for the leisure IV.

very similar estimation results (see Appendix Table C.5). Second, we refer to the very similar results provided by the pre-leisure IV. These estimates should be unbiased, even if potentially endogenous planned CDs remain in the estimation sample, since we exploit only variations across working days. Here, we only need the much weaker assumption that planned CDs are equally likely across working days. To make this even more explicit, we present a BPM model that uses the observations of births on working days only. Put differently, we exclude all births on leisure days. The estimated treatment effects of this specification (see Appendix Table C.6) do not substantially differ from those presented in Table 6.

3.2.1 Naïve estimation model

We estimate a naïve probit model for two reasons. First, for comparison with the existing correlational literature. Second, as a benchmark to assess endogeneity concerns against. The results suggest that a CD at parity 0 is associated with a decrease in the likelihood of a second birth by only 5.3 percentage points three years after the first birth (see Appendix Table C.4 for full estimation output). The effect increases only slightly (in absolute terms) over the period of 4–6 years after the first birth. Thus, the naïve probit model provides qualitatively equivalent results (compared to the BPM); however, the estimates are substantially downward biased in absolute terms. This illustrates the importance of accounting for endogeneity in CDs. A more formal way to assess the endogeneity of CDs is using the parameter ρ (see the last rows of panels A and B of Tables 6 and 7). The estimated correlations between the error terms of the two equations are statistically significant 4–6 years after the first birth. Their sign suggests that unobservable factors that are positively related to a CD are also positively related to the likelihood of a second birth. One possible explanation is that mother's with a taste for non-medically indicated CDs tend to have more children.

3.2.2 Treatment effect heterogeneity

A CD can influence future fertility through a biological mechanisms (i.e. women are biologically less likely to get pregnant and/or have a successful pregnancy) or choice (i.e. women decide having less additional children as a result of a CD). We would expect the magnitude of the first mechanism to be similar among women with low and high education. However, one would expect choices to be affected differently. First, there are baseline differences in the share of women that intent to have an additional child after the first born. For example, based on data from the first wave of the *Austrian Generations and Gender Survey*, we observe that 51% of the women with an university degree intend to have a child in the next three years, while the share halves to 26% among those with at most compulsory education. The fertility preferences of highly-educated mothers might not only be stronger, but also slightly more persistent. Second, to the extent that medical complications associated to a CD can result in a longer absence from work, we would expect those more attached to the labor market to be less likely to choose to have another child. Therefore, the overall difference between mothers with differential educational background in the magnitude of the fertility response to a CD is an empirical issue. To explore potential treatment effect heterogeneity, we estimate our model including the interaction between CD and an indicator for high education (defined as at least lower secondary schooling) in equation (1). Because this interaction is endogenous, we include the interaction between our instrument and the indicator for high education as a second IV in equations (2) and (3). The results in Appendix Table C.7 suggest slightly smaller, albeit statistically significant, fertility effects for highly-educated mothers.

3.2.3 Tempo versus quantum effect

Since the vast majority of mothers in our estimation sample is still in child-bearing age even six years after their first birth (see Appendix Figure C.2), we cannot conclude whether the estimated impact is a so-called *quantum* or only a *tempo effect*. Put differently, it is possible that treated women only delay childbearing. If so, we can expect shrinking treatment effects beyond the sixth year after the first birth, reflecting additional births by treated mothers, which compensate for reductions in earlier years (tempo effect). Alternatively, a CD at parity 0 could reduce lifecycle fertility (quantum effect). In this case, the treatment effects are expected to stay constant (or at least significantly negative) beyond the sixth year until the end of the treated women's reproductive years.

To determine whether the estimated effects capture a quantum or tempo effect, we conduct two additional estimation analyses. First, we estimate our model for the subgroup of mothers who we can observe at least up to 10 years after their first birth. In doing so, we explore the idea that the likelihood of a second child decreases as the age of the first child increases. Second, we impose an age restriction and focus on a sub-group of mothers who were close to the end of their reproductive years. If the estimated effects hold in this sample, it would support the interpretation of a quantum effect.

Longer horizon The blue (dark) bars in Figure 4 summarize the estimated marginal effects for the subsample of mothers who gave birth before 2004. An advantage of this subgroup is that it allows us to observe their subsequent fertility up to 10 years after their first birth. The blue (light) bars replicate the baseline estimates for the full sample of births until 2007. This comparison reveals two important findings: first, the two sets of estimates are comparable in the first six years after birth and second, the effects from the seventh to tenth year after birth are similar to the estimate for the sixth year. The stability of the estimated treatment effects over a 10-year period after the first birth supports the interpretation of a quantum effect.

Longer horizon for older mothers The green (light) bars in Figure 5 summarize the estimated marginal effects of a CD on the likelihood of a second birth for the subsample comprising mothers who were 30 years or older at the first birth. These women will generally be between 40 and 55 years old after a period of ten years. The blue (dark) bars replicate the estimates for all mothers who gave birth before 2004, irrespective of their age at first birth. We observe only minor differences in the estimated effects for average-aged and older mothers. Ten years after the first birth, the fertility effect is −19 percentage points for average-aged mothers and about −21 percentage points for older mothers.[24] We interpret this as strong evidence in favor of a quantum effect of a CD at first birth and reject the hypothesis that the estimated fertility effects of a CD only represent a postponement of subsequent births (tempo effect).

[24]Notably, the first stage of the leisure IV is somewhat stronger for older women than average-aged women. A birth on a weekend or public holiday is less likely to be a CD by 5.8 percentage points in the former and 4.4 percentage points in the latter (see Appendix Tables C.8 and C.9).

3.2.4 Third birth

Our final fertility analysis focuses on the potential impact of a CD at parity 0 on the probability of having a third child. Here as well, we restrict the sample to mothers who we observe at least for ten years after their first birth. In this analysis, we disregard the timing and mode of any second birth since these are "bad controls." We simply estimate the effects of a CD at first birth on the likelihood of having a third child in subsequent years using the above-described BPM. The estimated marginal effects are summarized by the green (dark) bars in Figure 6. We find a significant and negative impact beginning seven years after the first birth. It is plausible that the onset of the effect on a third child is delayed compared to that on a second child (see blue (light) bars). The estimated effect in the seventh year amounts to −4.6 percentage points and increases (in absolute terms) over time up to −6.2 percentage points in the tenth year after the first birth. The latter effect is equivalent to a reduction of 54 percent. Thus, the estimated treatment effect is about twice that on the probability of a second child. This result provides additional evidence for a permanent and negative impact of CD at parity 0 on subsequent fertility.[25]

3.3 Effects of CDs on subsequent labor market behavior

Given the permanent negative impact of a CD at parity 0 on subsequent fertility, it seems plausible that treated mothers also adjust their labor market behavior. Since a lower number of children correspond with lower childcare responsibilities, one would expect an increase in maternal labor supply.

To test this hypothesis, we estimate a BPM for maternal employment. Table 7 summarizes the estimation results on the basis of the two alternative IV approaches (panels A and B). The dependent variable $E_{i(t+p)}$ is equal to one if the mother is employed p years after the first birth. The different columns refer to a series of separate estimations and show the estimated marginal effect of a CD at parity 0 on the likelihood of maternal employment 1–6 years after the first birth. To identify the mother's employment status, we verify whether she has any employment spell in the calendar months 12, 24, 36, ... after birth. We find a statistically significant positive effect on the likelihood of maternal employment starting 3 years after the first birth. The timing of this effect is plausible since the onset of the fertility and labor supply effect coincide. In this particular year, after the first birth, treated mothers are 13.5 percentage points more likely to be employed. Given an average employment rate of about 51 percent, the estimated treatment effect corresponds to an increase of 26.4 percent. Thus, the quantitative effect of a CD at parity 0 on fertility and labor market behavior are comparable. Four years after birth, the effect on labor market behavior peaks at about 14.5 percentage points (see Panel A) and declines thereafter. Six years after the first birth, treated mothers are still about 9 percentage points (or about 14 percent) more likely to be employed. This pattern, which is comparable across both IV approaches, makes perfect sense since part of the untreated mothers will return to the labor market as their "additional" child grows older.

The statistically significant negative impact of CD on employment in the first year after birth may point to direct negative health effects after the surgery. Mothers may suffer from serious physical or mental health problems, for example, as a CD impedes a mother from returning to

[25]For the probability of a fourth child, we find no statistically significant effects. However, only a small share of women have more than three children (after 6 years: 0.39 percent; after 10 years: 1.56 percent).

the labor market quickly. Note, however, only a rather small share of women (21 percent) are employed in the first year after birth. This results from the Austrian parental leave system that provides paid parental leave up to two years after birth. The estimated effect is less pronounced in the leisure IV specification. Figure 7 provides a graphical representation of the employment effects.

In a next step, we aim to determine how much additional labor income is generated from the positive effect of a CD at parity 0 on maternal labor supply. To do so, we consider the first six years after birth. One way to approximate the change in labor income is by simply combining our estimated effect on the employment probability (Table 7) using the average income of employed women. This simple calculation leads to an effect of about Euro $8,763$ or 17 percent. Clearly, there are more sophisticated procedures.[26] A tractable solution is to impute the earnings of zero to all non-employed women (Neal and Johnson, 1996) and estimate a Tobit type 1 model. In our specific case, we face the challenge of a Tobit type 1 model with a binary endogenous regressor. We implement our IV estimation applying a control function approach using a Tobit specification for daily earnings. This procedure provides us with a comparable earnings gain of up to 22 percent three years after delivery, as shown in Appendix Table C.11. The coefficient on CD is statistically significant only in year three of the leisure IV specification. The Tobit approach does not work in the case of the pre-leisure IV (see Section 3.4).

The further analysis of maternal labor supply is also based on the leisure IV. The Web Appendix C includes full estimation output for the leisure and pre-leisure IVs. We obtain equivalent results for both IVs.

3.3.1 Treatment effect heterogeneity

Appendix Table C.13 summarizes the estimated effects for mothers with low and high educational attainment. The results suggest that both groups increase their labor supply in the third to the sixth year after birth. However, the interaction effects suggest stronger effects among highly-educated mothers. The fact that employment of highly-educated women is more responsive is consistent with previous literature. For example, Lundborg et al. (2017) find that women with higher earnings potential are more responsive to having children than women with low earnings potential. Hence, we restrict the following analyses (part-time versus full-time employment and a longer time horizon) to mothers with high educational attainment.

3.3.2 Intensive margin

The set of estimations discussed above informs us about labor supply responses at the extensive margin. It would be interesting to further analyze adjustments at the intensive margin. While our data do not include information on (contracted) working hours, we can use the pre-treatment earnings as a natural reference point to define a proxy variable for the adjustment along the intensive margin. For conceptual simplicity, we assume that all employed women were full-time employees before their first birth, which seems to be a plausible assumption for young and well-educated

[26] For instance, one could also employ a self-contained estimation analysis with the wage rate as an outcome variable. Such an analysis, however, gives rise to the classical problem in applied labor economics that the wage rate is censored for non-employed women. The econometric literature provides methods for the joint estimation of a structural model of the participation decision and wage determination. A credible implementation of such a model, however, requires an additional exclusion restriction and is beyond the scope of this add-on exercise.

women without children. Mothers are considered full-time employees in the post-treatment period if they earn at least 75 percent of their pre-treatment salary.[27] If their post-treatment salary is less than 75 percent of their pre-treatment salary, we call them part-time employed. According to this definition, the share of highly-educated part-time employed mothers (among highly-educated and employed mothers) is about 41 percent one year after the first birth and steadily declines over time to about 34 percent six years after the first birth. Figure 8 summarizes the estimates for the part-time and full-time employment of highly-educated mothers. The estimates reveal that our employment effects are driven by an increase in full-time employment. We find positive effects of CD on full-time employment from 2–4 years after the first birth. The quantitative impact reaches a statistically significant maximum of 13.3 percentage points in year 3 (leisure IV specification). The effects on part-time employment are small and in all cases, statistically non-significant. Given that the main fertility effect is observable in years 3 and 4 — i.e., for a majority of the second births among untreated women — it is plausible to see the most pronounced labor market responses in this period.

3.3.3 Longer horizon

To investigate the employment effects beyond the sixth year after birth, we confine the analysis to highly-educated women who gave birth before 2004. This allows us to follow their labor market behavior up to at least ten years after the first birth. Figure 9 summarizes the estimated treatment effects of the respective BPMs and contrasts these with the baseline estimates for highly-educated mothers. This new set of results confirms a trend that was already adumbrated by our baseline specification: the positive employment effect of a CD at parity 0 peaks after three years, decreases thereafter, and becomes statistically non-significant in the subsequent years. This pattern is in line with the fertility effects and availability of childcare institutions in Austria. On average, an untreated mother's "additional second child" turns three between the sixth and seventh year after the first birth. Public childcare facilities are widely available for children from the age of three. Austria has a well-structured *Kindergarten* network that provides heavily subsidized daycare for children aged 3–6 years. However, there is a perpetual shortage of day nurseries (*Kinderkrippe*) for children below the age of three. Thus, we interpret the fading out of the employment effect as untreated women return to the labor market when their second child starts *Kindergarten*.

3.4 Sensitivity analysis

We check the sensitivity of our estimation results with respect to the inference method, included covariates, and controlling for hospital competition. We briefly report on these sensitivity checks. The detailed estimation output based on the leisure and pre-leisure IVs is provided in the Web Appendix C.

Estimation methods The advantage of the BPM is that it explicitly accounts for the binary nature of the treatment and outcome variable. However, it imposes functional form assumptions on the error terms (i. e., bivariate normal distribution with zero means, unit variances, and covariance ρ). For comparison, we re-estimated our models with a linear IV model that ignores the binary structure of outcome and treatment variables. This 2SLS estimation works well for the leisure IV.

[27] We exclude all mothers who were not employed in the year before birth from this analysis (22,313 women or 16 percent).

The estimated local average treatment effects on fertility and maternal labor supply (see Appendix Tables C.17 and C.18) are comparable to the average treatment effects obtained using the BPM. For instance, the 2SLS suggests that a CD (due to delivery on a leisure day) reduces the probability of having another child six years after birth by 10.7 percentage points. The 2SLS estimates also confirm the positive effect on maternal employment. A CD at parity 0 significantly increases maternal employment 3–4 years after the first birth. The quantitative effect peaks in the third year (10.1 percentage points) and decreases thereafter. Earnings results from a 2SLS model (shown in Appendix Table C.12) give a similar picture as compared to the Tobit approach.[28] We also use 2SLS to investigate robustness with respect to geographical fixed effects and find remarkably stable results. Appendix Tables C.19 and C.20 report 2SLS results with district fixed effects. We estimate and report a Two Stage Residual Inclusion (2SRI) model to further investigate the methodological robustness of our results. 2SRI provides another estimation method that accommodates the binary nature of the dependent variables. It includes residuals from the first stage regression in the second stage, with probit models on both stages. The approach is similar to 2SLS in spirit and does not rely on the joint normality assumption, while sharing its binary structure with the BPM. The 2SRI results in Appendix Tables C.21 and C.22 are remarkably similar to our BPM estimates.

Covariates In our baseline specification, we do not control for gestation length and birth weight, since these two variables are potentially affected by mode of delivery. This is particularly true for planned CDs. Our results are not sensitive to the inclusion of these two covariates. The estimated treatment effects basically do not change (see Appendix Tables C.23 and C.24). The only difference is that treatment effects on maternal employment are slightly lower 3–6 years after the first birth. The result corroborates our presupposition that the estimation sample does not comprise a significant number of planned CDs.

Hospital competition Most of our covariates capture the demand side. We construct a simple proxy for hospital competition to verify whether supply-side characteristics affect the impact of the weekday of birth on the likelihood of a CD. In particular, we use an annual data set on the number of physicians employed in obstetric and gynecological departments in all Austrian hospitals. For each observation in our sample, we define an area of 45 minutes car drive. Then we divide the number of hospital-based physicians within this area by the total population measured in $1,000$. The resulting variable varies between 0 and 3.02 and has a mean of 0.25. Based on this measure, we define ten quantile variables and include these as additional covariates in our regression analysis. The first stage is basically unchanged after controlling for hospital competition. The same holds true for the estimated effect of a CD on subsequent fertility behavior and maternal employment (see Appendix Tables C.25 and C.26). Finally, we do not find important variation in the estimated effects across sub-samples below and above the sample median. We conclude that our estimated effects are not related to hospital competition.

[28]Based on the pre-leisure IV, the 2SLS model does not replicate the significant effects on fertility and maternal labor market outcomes. It seems that the ignorance of the binary structure of the outcome and treatment variables only works if the IV is very strong (as in the case of the leisure IV). The same applies for the Tobit approach used for maternal earnings.

4 Conclusions

About 28 percent of all births in OECD member countries were cesarean delivered in 2013. However, Molina et al. (2015) recommend that this share should not exceed 19 percent to optimize maternal and neonatal outcomes.[29] Thus, at least 32 percent of all CDs could be avoided. Our results suggest that reducing CD rates to the recommended level would substantially decrease healthcare costs and increase fertility.

Our IV estimation results imply that mothers who have a non-medically indicated CD at parity 0 are 18.7 and 6.2 percentage points less likely to give birth to a second and third child. This corresponds to an estimated reduction in lifetime fertility of 0.25 children (13.6 percent). Assuming the estimated effect size applies more broadly, a reduction in the average CD rate from 28 percent to a recommended level of 19 percent would increase the fertility rate by 0.018 children per woman. This would increase the average total fertility rate in the OECD area by 1.1 percent (see Web Appendix D). Our analyses suggest that such fertility gains would be accompanied by moderate reductions in female labor supply.

We find that the effect on maternal employment starts 3 years after the first birth, it is largest 4 years after birth, and declines thereafter. This pattern is not only consistent with our fertility effects, but also with the existing evidence that investigates the effects of fertility on maternal employment at the intensive margin (see, e.g., discussion in Lundborg et al. (2017)). This suggests that the overall effects of CDs on employment are driven by their impact on fertility rather than other channels like maternal or children health.

Although further research on the causal effects of CDs on the short and long-term health outcomes of children and their mothers is needed, the existing evidence on the effects of non-medically indicated CDs driven by physicians' preferences for leisure suggests that they also worsen some health outcomes (Costa-Ramón et al., 2018, 2019). This suggests that policies aimed at decreasing (at least these) unnecessary CDs could potentially not only increase fertility, but also decrease healthcare costs and improve population health.

Profound public efforts are being made in all high-income countries to foster fertility. In 2014, OECD member countries spent an average of 2.6 percent of GDP on families (OECD Public Social Expenditure Database, category: public social spending on the family). Many of these countries implement specific policies to increase fertility such as fully subsidized assisted reproduction. Our analysis shows that one sensible policy to increase fertility is to reduce CD rates to the recommended level. To achieve this goal, supply and demand need to be addressed. Healthcare market regulations need to reduce monetary supply-side incentives for CDs. The reimbursement of obstetric care in diagnoses-related groups based on hospital funding systems could be changed in favor of vaginal births. To reduce the demand for CDs to a healthy level, public health campaigns should be implemented.

A limitation of our analysis is that we cannot directly reveal the causal mechanism. It remains unclear why CD mothers have reduced fertility. However, given the rather low associations between CDs and physiological reproductive problems in observational studies, we suspect the mother's choice to be the more important causal channel. Clearly, future research is needed to explore the role of a physiological or choice mechanism. Moreover, more design-based approaches are needed

[29] In fact, World Health Organization (2015) recommends an even lower level of 10–15 percent.

to study the effects of a CD on children. Our results highlight that any analysis of medium- or long-term child outcomes must account for the fact that treated children generally belong to smaller families.

References

Almgren, M., Schlinzig, T., Gomez-Cabrero, D., Gunnar, A., Sundin, M., Johansson, S., Norman, M., and Ekstróm, T. J. (2014). Cesarean Delivery and Hematopoietic Stem Cell Epigenetics in the Newborn Infant: Implications for Future Health? *American Journal of Obstetrics and Gynecology*, 211(5):502.e1–502.e8.

Angrist, J. D. and Evans, W. N. (1998). Children and Their Parents' Labor Supply: Evidence from Exogenous Variation in Family Size. *American Economic Review*, 88(3):450–477.

Bailey, M. J. (2006). More Power to the Pill: The Impact of Contraceptive Freedom on Women's Life Cycle Labor Supply. *Quarterly Journal of Economics*, 121(1):289–320.

Belizán, J. M., Althabe, F., Barros, F. C., and Alexander, S. (1999). Rates and Implications of Caesarean Sections in Latin America: Ecological Study. *British Medical Journal*, 319:1397–1402.

Bell, C. M. and Redelmeier, D. A. (2001). Mortality Among Patients Admitted to Hospitals on Weekends as Compared With Weekdays. *New England Journal of Medicine*, 345(9):663–668.

Bragg, F., Cromwell, D., Edozien, L., and et al. (2010). Variation in Rates of Caesarian Section Among English NHS Trusts After Accounting for Maternal and Clinical Risk: Cross Sectional Study. *British Medical Journal*, 341:c5065.

Brown III, H. S. (1996). Physician Demand for Leisure: Implications for Cesarean Section Rates. *Journal of Health Economics*, 15:233–242.

Card, D., Fenizia, A., and Silver, D. (2018). The Health Effects of Cesarean Delivery for Low-Risk First Births". Unpublished manuscript (updated april 2018), University of California, Berkeley.

Cardwell, C. R., Stene, L. C., Joner, G., Cinek, O., Svensson, J., Goldacre, M. J., Parslow, R. C., Pozzilli, P., Brigis, G., Stoyanow, D., Urbonaitė, B., Šipetić, S., Schober, E., Ionescu-Tirgoviste, C., Devoti, G., de Beaufort, C. E., Buschard, K., and Patterson, C. C. (2008). Caesarean Section is Associated With an Increased Risk of Childhood-Onset Type 1 Diabetes Mellitus: a Meta-Analysis of Observational Studies. *Diabetologia*, 51(5):726–735.

Chiburis, R. C., Das, J., and Lokshin, M. (2012). A Practical Comparison of the Bivariate Probit and Linear IV Estimators. *Economics Letters*, 117(3):762–766.

Committee on Obstetric Practice (2007). Cesarean Delivery on Maternal Request. *ACOG Committee Opinion*, 394:1–4.

Costa-Ramón, A. M., Kortelainen, M., Rodríguez-González, A., and Sääksvuori, L. (2019). The Long-Run Effects of Cesarean Sections. VATT Working Papers 125, VATT Institute for Economic Research, Helsinki, Finland.

Costa-Ramón, A. M., Rodríguez-González, A., Serra-Burriel, M., and Campillo-Artero, C. (2018). It's about Time: Cesarean Sections and Neonatal Health. *Journal of Health Economics*, 59:46–59.

Currie, J. and MacLeod, W. (2008). First Do No Harm? Tort Reform and Birth Outcomes. *Quarterly Journal of Economics*, 123(2):795–830.

Daltveit, A. K., Tollånes, M. C., Philstrøm, M., and Irgens, L. M. (2008). Cesarean Delivery and Subsequent Pregnancies. *Obstetrics and Gynecology*, 111(6):1327–1334.

Declercq, E. R., Sakala, C., Corry, M. P., and Applebaum, S. (2006). Listening to Mothers II: Report of the Second National U.S. Survey of Women's Childbearing Experiences. New York: Childbirth Connection.

DGGG (2010). Absolute und relative Indikationen zur Sectio caesarea. Leitlinien, Empfehlungen, Stellungnahmen, Deutsche Gesellschaft für Gynäkologie und Geburtshilfe.

Dobkin, C. (2003). Hospital Staffing and Inpatient Mortality. Unpublished Manuscript, University of California, Berkeley.

Dubay, L., Kaestner, R., and Waidmann, T. (1999). The Impact of Malpractice Fears on Cesarean Section Rates. *Journal of Health Economics*, 18(4):491–522.

Epstein, A. J. and Nicholson, S. (2009). The Formation and Evolution of Physician Treatment Styles: an Application to Cesarean Sections. *Journal of Health Economics*, 28(6):1126–1140.

FHF (2008). Étude sur les Césariennes. Fédération Hospitalière de France, Paris.

Gibbons, L., Belián, J. M., Lauer, J. A., Betrán, A. P., Merialdi, M., and Althabe, F. (2010). The Global Numbers and Costs of Additionally Needed and Unnecessary Caesarean Sections Performed per Year: Overuse as a Barrier to Universal Coverage. *World Health Report*, Background Paper No 30.

Goldani, H. A. S., Bettiol, H., Barbieri, M. A., Silva, A. A. M., Agranonik, M., Morais, M. B., and Goldani, M. Z. (2011). Cesarean Delivery is Associated With an Increased Risk of Obesity in Adulthood in a Brazilian Birth Cohort Study. *The American Journal of Clinical Nutrition*, 93(6):1344–1347.

Gould, J. B., Qin, C., Marks, A. R., and Chavez, G. (2003). Neonatal Mortality in Weekend vs Weekday Births. *Journal of the American Medical Association*, 289(22):2958–2962.

Grant, D. (2009). Physician Financial Incentives and Cesarean Delivery: New Conclusions from the Healthcare Cost and Utilization Project. *Journal of Health Economics*, 28(1):244–250.

Gruber, J. and Owings, M. (1996). Physician Financial Incentives and Cesarean Section Delivery. *The Rand Journal of Economics*, 27:99–123.

Gurol-Urganci, I., Bou-Antoun, S., Lim, C., Cromwell, D., Mahmood, T., Templeton, A., and van der Meulen, J. (2013). Impact of Caesarean Section on Subsequent Fertility: A Systematic Review and Meta-analysis. *Human Reproduction*, 28(7):1943–1952.

Hanley, G., Janssen, P., and Greyson, D. (2010). Regional Rariation in the Cesarian Delivery and Assisted Vaginal Delivery Rates. *Obstetrics and Gynecology*, 115(6):1201–1208.

Husslein, P. and Langer, M. (2000). Elektive Sektio vs. Vaginale Geburt - ein Paradigmenwechsel in der Geburtshilfe? *Der Gynaekologe*, 33(12):849–856.

Jensen, V. and Wüst, M. (2015). Can Caesarean Section Improve Child and Maternal Health? The Case of Breech Babies. *Journal of Health Economics*, 39:289–302.

Koechlin, F., Lorenzoni, L., and Schreyer, P. (2010). Comparing Price Levels of Hospital Services across Countries: Results of a Pilot Study. OECD Health Working Paper, No. 53, OECD Publishing.

Koplin, J., Allen, K., Gurrin, L., Osborne, N., Tang, M. L. K., and Dharmage, S. (2008). Is Caesarean Delivery Associated with Sensitization to Food Allergens and IgE-Mediated Food Allergy: A Systematic Review. *Pediatric Allergy and Immunology*, 19(8):682–687.

Kozhimannil, K. B., Law, M. R., and Virnig, B. A. (2013). Cesarean Delivery Rates Vary Tenfold Among US Hospitals; Reducing Variation May Address Quality and Cost Issues. *Health Affairs*, 32(3):527–535.

Lundborg, P., Plug, E., and Würtz Rasmussen, A. (2017). Can Women Have Children and a Career? IV Evidence from IVF Treatments. *American Economic Review*, 107(6):1611–1637.

Martins, C. and Gaffan, E. A. (2000). Effects of Early Maternal Depression on Patterns of Infant-Mother Attachment: A Meta-Analytic Investigation. *Journal of Child Psychology and Psychiatry*, 41(6):737–746.

Maurin, E. and Moschion, J. (2009). The Social Multiplier and Labor Market Participation of Mothers. *American Economic Journal: Applied Economics*, 1(1):251–272.

Menacker, F. and Hamilton, B. E. (2010). Recent Trends in Cesarean Delivery in the United States. NCHS data brief, no 35, Hyattsville, MD: National Center for Health Statistics.

Molina, G., Weiser, T. G., Lipsitz, S. R., Esquivel, M. M., Uribe-Leitz, T., Azad, T., Shah, N., Semrau, K., Berry, W. R., Gawande, A. A., and Haynes, A. B. (2015). Relationship Between Cesarean Delivery Rate and Maternal and Neonatal Mortality. *Journal of the American Medical Association*, 314(21):2263–2270.

Murphy, D. J., Stirrat, G. M., Heron, J., and the ALSPAC Study Team (2002). The Relationship Between Cesarean Section and Subfertility in a Population Based Sample of 14541 Pregnancies. *Human Reproduction*, 17(7):1914–1917.

Neal, D. A. and Johnson, W. R. (1996). The Role of Pre-Market Factors in Black-White Wage Differences. *Journal of Political Economy,*, 104(5):869–895.

NICE (2011). Caesarean Section, NICE Guidelines. https://www.nice.org.uk/guidance/cg132/resources. Accessed: 2014-02-02.

Norberg, K. and Pantano, J. (2016). Cesarean Sections and Subsequent Fertility. *Journal of Population Economics*, 29(1):5–37.

OECD (2013). *Health at a Glance 2013: OECD Indicators*. OECD Publishing, Paris.

OECD (2015). *Health at a Glance 2015: OECD Indicators*. OECD Publishing, Paris.

OECD (2016). OECD Family Database. Accessed: 2016-07-05.

OFSP (2013). Accouchements par Césarienne en Suisse [Births by Caesarians in Switzerland]. Office féderal de la Santé Publique, Bern.

O'Neill, S. M., Agerbo, E., Kenny, L. C., Henriksen, T. B., Kearney, P. M., Greene, R. A., and Khashan, P. M. A. S. (2014). Cesarean Section and Rate of Subsequent Stillbirth, Miscarriage, and Ectopic Pregnancy: A Danish Register-Based Cohort Study. *PLoS Medicine*, 11(7).

O'Neill, S. M., Kearney, P. M., Kenny, L. C., Henriksen, T. B., Lutomski, J. E., Greene, R. A., and Khashan, A. S. (2013). Caesarean Delivery and Subsequent Pregnancy Interval: A Systematic Review and Meta-analysis. *BMC Pregnancy and Childbirth*, 13(1).

O'Neill, S. M., Kearney, P. M., Kenny, L. C., Khashan, A. S., Henriksen, T. B., Lutomski, J. E., and Greene, R. A. (2013). Caesarean Delivery and Subsequent Stillbirth or Miscarriage: Systematic Review and Meta-Analysis. *PLOS ONE*, 8(1).

Sachs, B., Kobelin, C., Castro, M., and Frigoletto, F. (1999). The Risks of Lowering the Cesarean-Delivery Rate. *New England Journal of Medicine*, 340:54–57.

SGGG (2015). Guideline Sectio Caesarea. Kommission Qualititätssicherung, Schweizerische Gesellschaft für Gynäkologie und Geburtshilfe, Universitäts-Frauenklinik Bern.

Silver, R. M., Landon, M. B., Rouse, D. J., Leveno, K. J., Spong, C. Y., Thom, E. A., Moawad, A. H., Caritis, S. N., Harper, M., Wapner, R. J., Sorokin, Y., Miodovnik, M., Carpenter, M., Peaceman, A. M., O'Sullivan, M. J., Sibai, B., Langer, O., Thorp, J. M., Ramin, S. M., and Mercer, B. M. (2006). Maternal Morbidity Associated With Multiple Repeat Cesarean Deliveries. *Obstetrics and Gynecology*, 107(6):1226–1232.

Spetz, J., Smith, M. W., and Ennis, S. F. (2001). Physician Incentives and the Timing of Cesarean Sections: Evidence from California. *Medical Care*, 39(6):536–550.

Staiger, D. and Stock, J. H. (1997). Instrumental variables regression with weak instruments. *Econometrica*, 65(3):557–586.

Stock, J. and Yogo, M. (2005). Testing for Weak Instruments in Linear IV Regression. In Andrews, D. W., editor, *Identification and Inference for Econometric Models*, pages 80–108. Cambridge University Press, New York.

Thavagnanam, S., Fleming, J., Bromley, A., Shields, M. D., and Cardwell, C. R. (2008). A Meta-Analysis of the Association Between Caesarean Section and Childhood Asthma. *Clinical and Experimental Allergy*, 38(4):629–633.

Triunfo, P. and Rossi, M. (2009). The Effect of Physicians' Remuneration System on the Cesarean Section Rate: The Uruguayan Case. *International Journal of Health Care Finance and Economics*, 9(4):333–345.

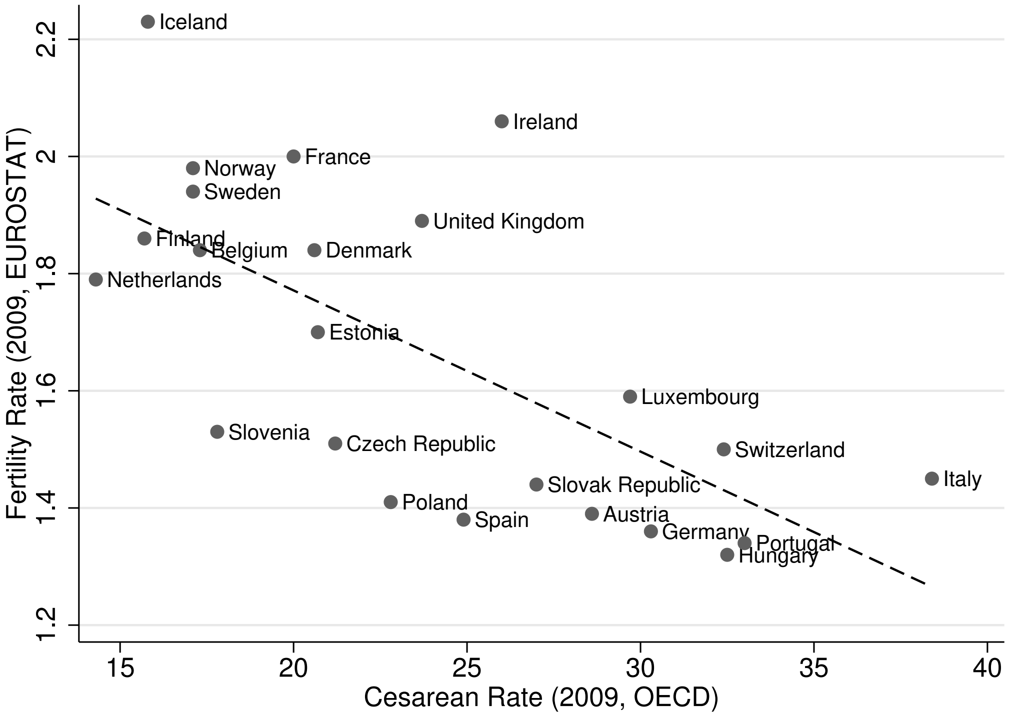
Weisman, O., Granat, A., Gilboa-Schechtman, E., Singer, M., Gordon, I., Azulay, H., Kuint, J., and Feldman, R. (2010). The Experience of Labor, Maternal Perception of the Infant, and the Mother's Postpartum Mood in a Low-Risk Community Cohort. *Archives of Women's Mental Health*, 13:505–513.

World Health Organization (2015). WHO Statement on Caesarean Section Rates. Executive Summary, Department of Reproductive Health and Research, Geneva, Switzerland.

Zweimüller, J., Winter-Ebmer, R., Lalive, R., Kuhn, A., Wuellrich, J., Ruf, O., and Büchi, S. (2009). The Austrian Social Security Database (ASSD). Working Paper 0903, The Austrian Center for Labor Economics and the Analysis of the Welfare State, University of Linz.

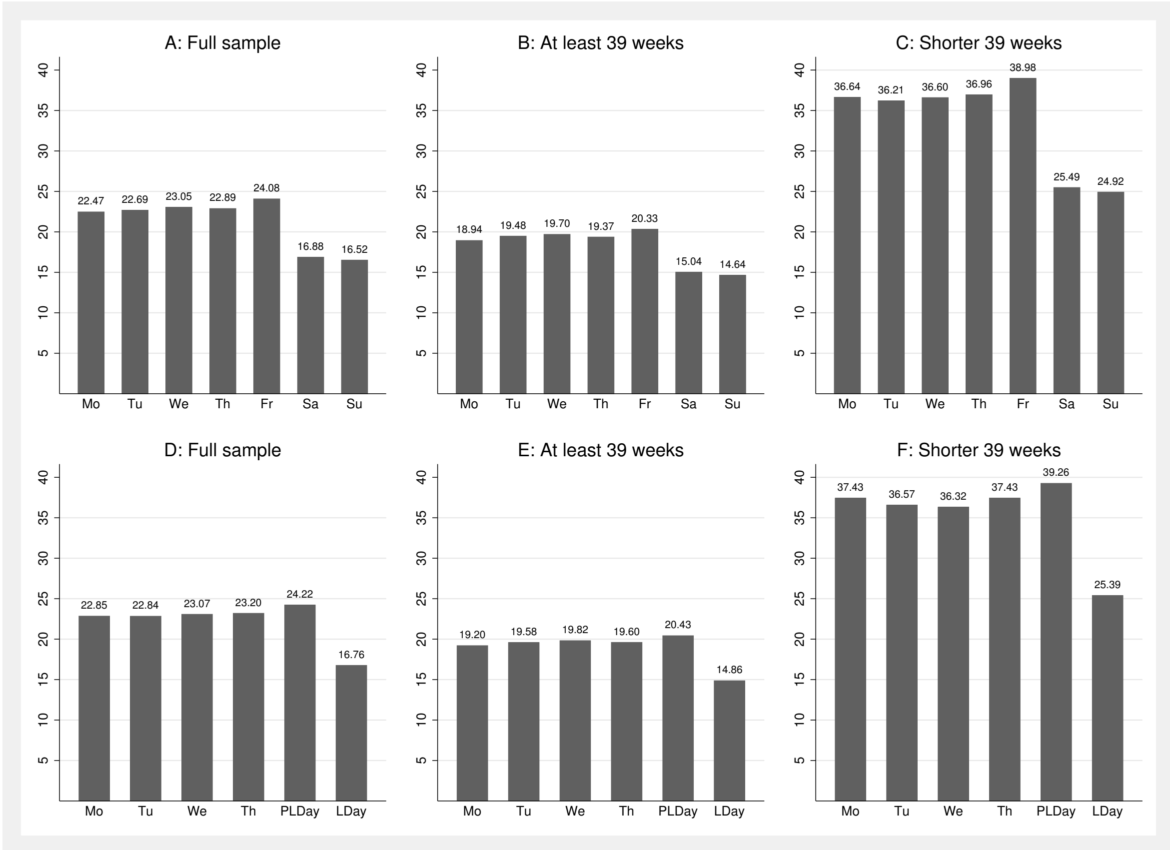
5 Tables and figures (to be placed in article)

Figure 1 – CD rate and fertility rates in selected European countries



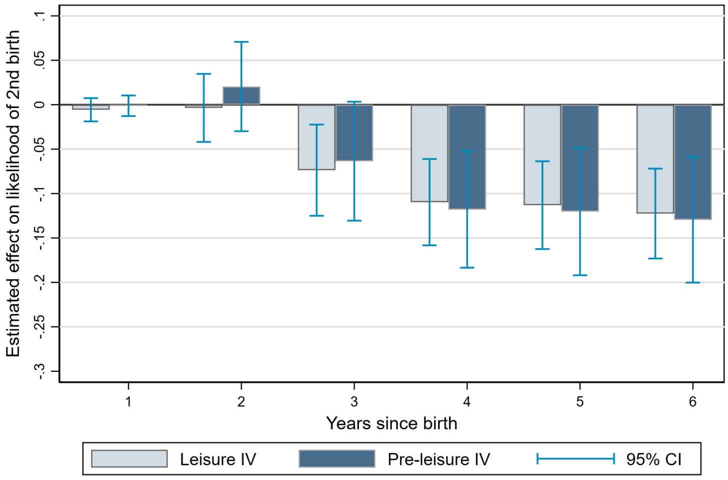
Notes: This figure depicts the cross-country correlation between CD rates and fertility rates for the sample of European countries for which OECD data on CDs is available. The correlation coefficient of −0.68 is significant at the 99.9 percent level.

Figure 2 – CD rates per weekday



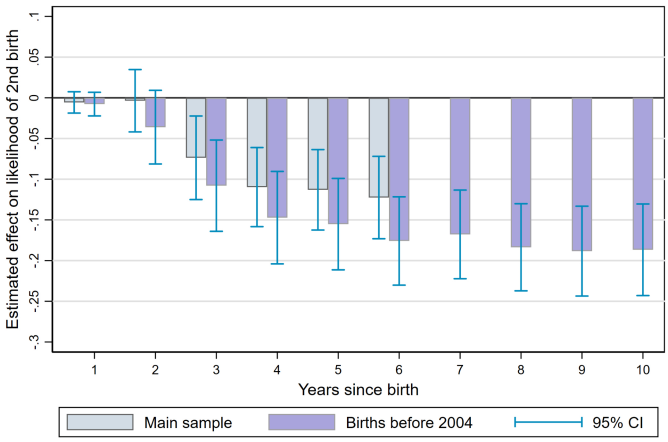
Notes: Data include births of nulliparous Austrian mothers for 1995–2007. The figure depicts CD rates in percent over weekdays (A, B, C) and type of days (D, E, F) for the full sample (A, D), the sample of deliveries at week 39 or later (B, E), and deliveries before week 39 (C, F), respectively. PLDay denotes Fridays and working days preceding a public holiday. LDay denotes Saturdays, Sundays and public holidays.

Figure 3 – Effect of a CD at parity 0 on the likelihood of a 2nd birth



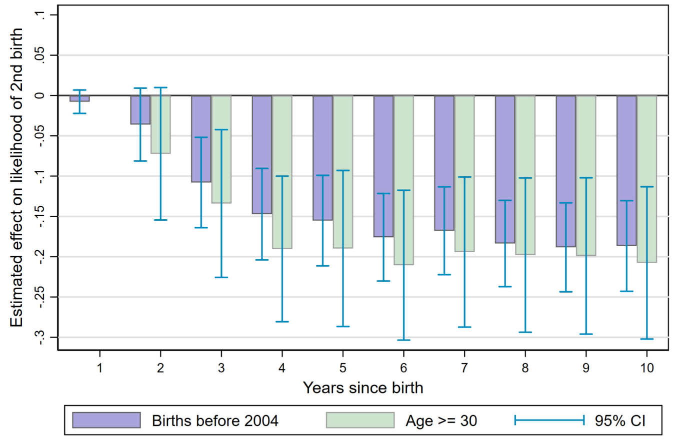
Notes: This graph depicts the estimated marginal effects of a CD at parity 0 on the likelihood of a second birth based on a series of bivariate probit models that use two alternative IVs. The estimates are equivalent to those summarized in Table 6. See notes to this table for further information.

Figure 4 – Effect of a CD on fertility up to 10 years after first birth



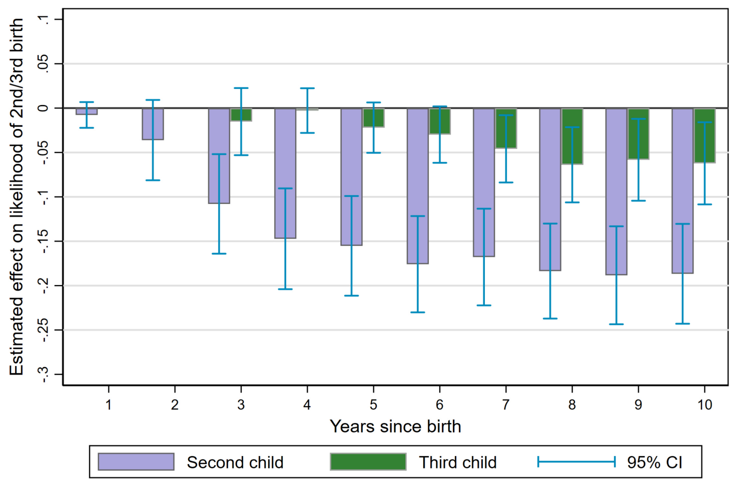
Notes: This graph depicts the estimated marginal effects of a CD at parity 0 on the likelihood of a second birth for the full sample of births from 1995-2007 (Main sample) and a subsample of births from 1995-2004 that we can observe up to at least 10 years (Births before 2004). The estimates are based on a series of bivariate probit models that use the leisure IV. Full estimation output for the 1995-2003 subsample can be found in Appendix Table C.8. See notes to this table for further information.

Figure 5 – Effect of a CD on fertility by average-aged and older mothers



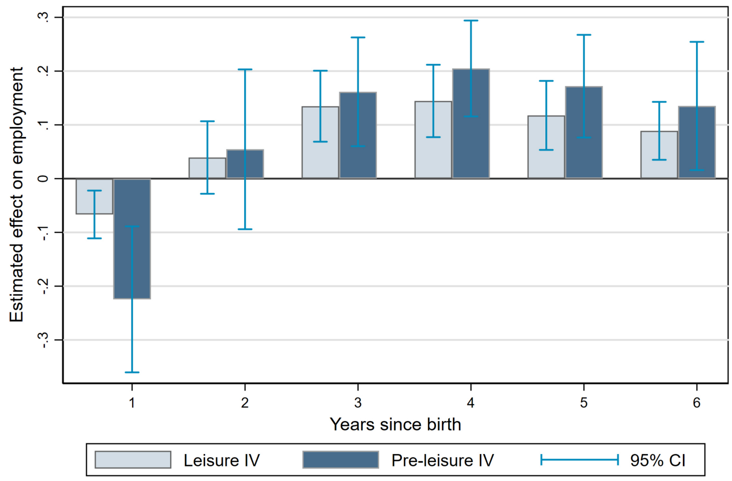
Notes: This graph depicts the estimated marginal effects of a CD at parity 0 on the likelihood of a second birth for the sample of births from 1995-2003 (Births before 2004) and, for the same time period, a subsample of older mothers who are at least 30 years old at first birth (Age>= 30). The estimates are based on a series of bivariate probit models that use the leisure IV. Full estimation output for the subsample of older mothers can be found in Appendix Table C.9. See notes to this table for further information.

Figure 6 – Effect of a CD at parity 0 on the likelihood of a 3rd birth



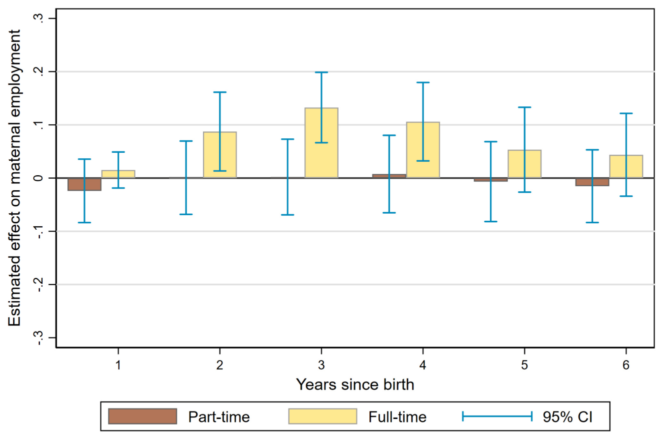
Notes: This graph depicts the estimated marginal effects of a CD at parity 0 on the likelihood of a second and third child including births from 1995-2003. The estimates are based on a series of bivariate probit models that use the leisure IV. Full estimation output for the likelihood of a third child can be found in Appendix Table C.10. See notes to this table for further information.

Figure 7 – Effect of a CD at parity 0 on mother's employment



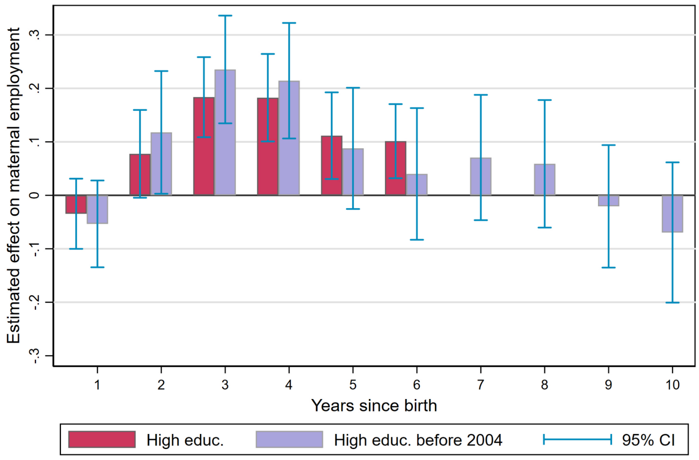
Notes: This graph depicts the estimated marginal effects of a CD at parity 0 on the likelihood of maternal employment based on a series of bivariate probit models that use two alternative IVs. The estimates are equivalent to those summarized in Table 7. See notes to this table for further information.

Figure 8 – Effect of a CD on full-time and part-time employment of highly-educated mothers



Notes: This graph depicts the estimated marginal effects of a CD at parity 0 on the likelihood of full-time and part-time maternal employment of highly-educated mothers. The estimates are based on a series of bivariate probit models that use the leisure IV. A mother is employed full-time (part-time) after treatment if she earns more (less) than 75 % of her pre-treatment salary. Full estimation output for this set of estimations can be found in Appendix Tables C.14 and C.15. See notes to these tables for further information.

Figure 9 – Effect of a CD on employment of highly-educated mothers up to 10 years after first birth



Notes: This graph depicts the estimated marginal effects of a CD at parity 0 on the likelihood of employment of highly-educated mothers including all births from 1995-2007 (High educ.) and a subsample of births from 1995-2003 that we can observe up to at least 10 years (High educ. before 2004). The estimates are based on a series of bivariate probit models that use the leisure IV. Full estimation output for the 1995-2003 subsample can be found in Appendix Table C.16. See notes to this table for further information.

Table 1 – Mother and child characteristics by treatment status

	(1) Vaginal Birth		(2) CD		(3) Difference
	Mean	SD	Mean	SD	
Girl	50.2		46.2		* * *
Birth weight	335.6	(40.2)	337.1	(47.6)	* * *
Gestation length	40.3	(0.9)	40.3	(1.0)	* * *
Mother's age	26.4	(4.9)	28.1	(5.4)	* * *
Mother's occupation					* * *
Self-employed	1.7		2.5		
Farming	0.6		0.6		
White collar	71.5		72.4		
Blue collar	15.6		14.6		
Missing	8.0		7.0		
Unknown	2.7		2.9		
Mother's education					* * *
Compulsory school	10.3		9.8		
Apprenticeship	38.9		36.7		
Interm. tech. and voc. school	19.0		18.6		
Academic secondary school	17.1		17.1		
Post-secondary college	5.0		5.1		
University	6.7		9.5		
Unknown	3.0		3.2		
Mother's family status					* * *
Single	51.6		48.3		
Married	46.4		48.6		
Widowed	0.1		0.1		
Divorced	1.9		3.0		
Mother's religion					* * *
Roman Catholic	85.1		83.2		
Protestant	4.2		4.4		
Old Catholic	0.1		0.1		
Jewish	0.0		0.0		
Muslim	1.3		1.0		
Other	0.7		0.7		
Unknown	8.6		10.5		
Province of residence					* * *
Burgenland	2.7		3.5		
Carinthia	7.4		6.8		
Lower Austria	18.3		19.3		
Upper Austria	18.6		16.0		
Salzburg	7.1		5.4		
Styria	15.0		18.7		
Tyrol	9.6		9.3		
Vorarlberg	5.0		3.8		
Vienna	16.3		17.1		
Number of observations	231,025		51,806		

Notes: This table depicts descriptive statistics measured at the time of birth for mothers with vaginal delivery (column (1)) and CD mothers (column (2)). Birth weight is measured in dekagrams, gestation length in weeks, mother's age in years. All other figures indicate percentages. Standard deviations in parentheses. Differences between groups are tested by means of two-independent-sample t-tests for birth weight, gestation length, and mother's age. Chi-square tests are employed for the categorical variables girl, mother's occupation, mother's education, mother's family status, mother's religion, and province of residence. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table 2 – Outcome variables by treatment status

	(1) Vaginal Birth Mean	(2) CD Mean	(3) Difference
Second child			
Year 1	0.6	0.4	* * *
Year 2	15.9	11.6	* * *
Year 3	34.4	28.0	* * *
Year 4	44.2	36.8	* * *
Year 5	50.3	42.4	* * *
Year 6	54.4	46.1	* * *
Third child			
Year 1	0.0	0.0	
Year 2	0.1	0.1	
Year 3	0.6	0.5	* * *
Year 4	2.2	1.4	* * *
Year 5	4.4	3.1	* * *
Year 6	6.4	4.5	* * *
Employment			
Year -4	67.5	73.5	* * *
Year -3	75.9	79.8	* * *
Year -2	82.5	84.7	* * *
Year -1	86.6	87.6	* * *
Year 0	7.1	8.4	* * *
Year 1	20.9	23.7	* * *
Year 2	44.1	47.7	* * *
Year 3	50.2	55.2	* * *
Year 4	54.0	58.7	* * *
Year 5	59.5	63.6	* * *
Year 6	64.5	68.2	* * *
Number of observations	231,025	51,806	

Notes: This table depicts outcome variables before and after birth for mothers with vaginal delivery (column (1)) and CD mothers (column (2)). The figures indicate percentages. Differences between groups are tested by means of chi-square tests. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table 3 – Mother and child characteristics for different birth days

	(1) Leisure Day		(2) Pre-leisure Day		(3) Other Day	
	Mean	SD	Mean	SD	Mean	SD
Cesarean	14.9		20.4		19.6	
Girl	49.3		49.4		49.5	
Birth weight	336.1	(41.3)	335.5	(42.0)	335.9	(41.8)
Gestation length	40.3	(0.9)	40.3	(0.9)	40.3	(0.9)
Mother's age	26.7	(5.0)	26.7	(5.1)	26.7	(5.1)
Mother's occupation						
Self-employed	1.8		1.9		1.9	
Farming	0.5		0.5		0.6	
White collar	71.8		71.8		71.5	
Blue collar	15.3		15.2		15.5	
Missing	7.9		7.7		7.8	
Unknown	2.8		2.8		2.8	
Mother's education						
Compulsory School	10.2		10.1		10.2	
Apprenticeship	38.3		38.2		38.7	
Interm. tech. and voc. School	19.0		19.1		18.8	
Academic secondary school	17.2		17.4		17.0	
Post-secondary college	5.0		5.1		5.0	
University	7.3		7.0		7.3	
Unknown	3.0		3.1		3.0	
Mother's family status						
Single	51.1		51.2		50.9	
Married	46.8		46.7		46.9	
Widowed	0.1		0.1		0.1	
Divorced	2.0		2.1		2.1	
Mother's religion						
Roman Catholic	84.6		84.8		84.9	
Protestant	4.4		4.3		4.1	
Old Catholic	0.1		0.1		0.1	
Jewish	0.0		0.0		0.0	
Muslim	1.2		1.2		1.2	
Other	0.7		0.8		0.7	
Unknown	9.0		8.8		8.9	
Province of residence						
Burgenland	2.8		2.8		2.9	
Carinthia	7.2		7.2		7.3	
Lower Austria	18.6		18.5		18.4	
Upper Austria	18.2		18.0		18.1	
Salzburg	6.9		6.8		6.8	
Styria	15.6		15.7		15.8	
Tyrol	9.2		9.8		9.6	
Vorarlberg	4.8		4.9		4.8	
Vienna	16.6		16.3		16.4	
Employment						
Year -4	68.3		68.7		68.6	
Year -3	76.6		76.9		76.6	
Year -2	83.0		83.1		82.8	
Year -1	86.8		86.9		86.7	
Number of observations	83,281		46,571		152,979	

Notes: This table depicts mother and child characteristics across different types of birth days. Birth weight is measured in dekagrams, gestation length in weeks. All other figures indicate percentages. Employment rates refer to years before birth. Standard deviations in parentheses.

Table 4 – Maternal health outcomes prior birth for different birth days

	(1) Leisure Day		(2) Pre-leisure Day		(3) Other Day	
	Mean	SD	Mean	SD	Mean	SD
Outpatient: 1-4 quarters prior birth						
Medical attendance	113.1	(97.4)	115.6	(104.5)	116.5	(101.3)
Medication	16.7	(102.0)	17.5	(70.4)	15.7	(82.5)
Outpatient: 5-8 quarters prior birth						
Medical attendance	55.6	(92.0)	56.3	(82.6)	56.9	(86.0)
Medication	16.7	(134.3)	16.6	(84.3)	13.4	(76.6)
Number of observations	2,960		1,621		5,221	
Inpatient: 1-4 quarters prior birth						
Days in general hospitals	0.38	(1.69)	0.39	(2.04)	0.40	(1.88)
Days in special clinics	0.02	(0.42)	0.02	(0.35)	0.02	0.60)
DRG revenues	124.8	(506.8)	130.2	(643.9)	132.3	(571.7)
Number of observations	2,251		1,166		3,903	

Notes: This table depicts healthcare expenditure and days of hospitalization across different types of birth days for a subsample of mothers at parity 0 who gave birth to a child in an Upper Austrian hospital between 2005 and 2007 in the upper panel, and 2006 and 2007 in the lower panel. Outpatient expenditure, inpatient expenditure, and DRG-revenues are measured in Euro per quarter over the period 1–4 and 5–8 quarters prior birth. Hospitalization indicates the number of days spent in hospital per quarter over one year prior birth. Standard deviations in parentheses. Data provided by the Upper Austrian Sickness Fund (Oberösterreichische Gebietskrankenkasse).

Table 5 – Mother characteristics by stated fertility preferences using data from the Generations and Gender Survey

	Do you want more kids in the next 3 years[a]			
	(1)		(2)	
	No more kids		More kids	
	Mean	SD	Mean	SD
Age	30.90	(8.83)	30.11	(5.63)
Education				
Compulsory School	0.119		0.058	
Apprenticeship/ Interm. tech. and voc. school	0.508		0.490	
Academic secondary school	0.192		0.206	
University	0.181		0.246	
Family status				
Single	0.749		0.709	
Married	0.193		0.246	
Widowed	0.003		0.002	
Divorced	0.054		0.043	
Religion				
Roman catholic	0.741		0.790	
Protestant	0.044		0.022	
Muslim	0.012		0.025	
Other	0.022		0.031	
Unknown	0.181		0.132	
Province of residence				
Burgenland	0.028		0.040	
Carinthia	0.047		0.078	
Lower Ausria	0.174		0.197	
Upper Ausria	0.166		0.193	
Salzburg	0.066		0.047	
Styria	0.151		0.139	
Tyrol	0.090		0.076	
Vorarlberg	0.050		0.040	
Vienna	0.227		0.190	
Self-assessed health (5-point scale)				
Very healthy	0.604		0.676	
Very healthy or healthy	0.904		0.944	
Number of observations	928		554	

Notes: The figures presented in this table are based on data from the first wave of the Austrian *Generations and Gender Survey* (GGS). These data provide micro-level information for about 5,000 men and women aged 18 to 45 on July 1st 2008 living in private households in Austria, who sufficiently spoke German. The figures are based on the sub-sample of female respondents with zero or one biological child. [a] The variable which is used to split the sample into two groups of women according to their stated fertility preferences is based on the question '*Do you intend to have a/another child during the next three years?*', which provided the following answer categories '*Definitely yes*', '*Probably yes*', '*Probably not*', and '*Definitely not*'. Column (1) comprises respondents who answered either '*Probably not*' or '*Definitely not*'. Column (2) comprises respondents who answered either '*Probably yes*' or '*Definitely yes.*

Table 6 – Bivariate probit results: the effect of a CD at parity 0 on the likelihood of a second birth

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0057	−0.0036	−0.0737***	−0.1097***	−0.1131***	−0.1225***
	(0.0067)	(0.0195)	(0.0262)	(0.0248)	(0.0252)	(0.0258)
Eq. (2): $leisure_{i(t=0)}$	−0.0497***	−0.0497***	−0.0498***	−0.0497***	−0.0497***	−0.0497***
	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1315	−0.0752	0.0349	0.0786**	0.0835**	0.0999**
	(0.2049)	(0.0474)	(0.0426)	(0.0380)	(0.0386)	(0.0399)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0012	0.0205	−0.0636*	−0.1181***	−0.1202***	−0.1295***
	(0.0059)	(0.0257)	(0.0342)	(0.0333)	(0.0366)	(0.0361)
Eq. (3): $preleisure_{i(t=0)}$	0.0081***	0.0081***	0.0081***	0.0081***	0.0081***	0.0080***
	(0.0024)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	−0.0160	−0.1327**	0.0184	0.0916*	0.0945*	0.1107**
	(0.1981)	(0.0610)	(0.0551)	(0.0510)	(0.0557)	(0.0554)
Mean of dependent variable	0.006	0.151	0.332	0.429	0.489	0.529
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes bivariate probit estimation results for the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table 7 – Bivariate Probit results: the effect of a CD at parity 0 on the likelihood of maternal employment

	(I) $E_{i(t+1)}$	(II) $E_{i(t+2)}$	(III) $E_{i(t+3)}$	(IV) $E_{i(t+4)}$	(V) $E_{i(t+5)}$	(VI) $E_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0667***	0.0393	0.1347***	0.1445***	0.1176***	0.0889***
	(0.0227)	(0.0344)	(0.0337)	(0.0344)	(0.0328)	(0.0275)
Eq. (2): $leisure_{i(t=0)}$	−0.0495***	−0.0497***	−0.0496***	−0.0495***	−0.0495***	−0.0496***
	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1554***	−0.0267	−0.1509***	−0.1728***	−0.1491***	−0.1176***
	(0.0464)	(0.0513)	(0.0507)	(0.0525)	(0.0504)	(0.0447)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.2246***	0.0545	0.1615***	0.2049***	0.1720***	0.1351**
	(0.0693)	(0.0758)	(0.0516)	(0.0455)	(0.0487)	(0.0609)
Eq. (3): $preleisure_{i(t=0)}$	0.0087***	0.0081***	0.0079***	0.0080***	0.0079***	0.0079***
	(0.0022)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.4870***	−0.0494	−0.1913**	−0.2659***	−0.2341***	−0.1924*
	(0.1441)	(0.1131)	(0.0786)	(0.0716)	(0.0771)	(0.0998)
Mean of dependent variable	0.214	0.448	0.511	0.549	0.602	0.652
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes bivariate probit estimation results for the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of maternal employment 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Web Appendix

This Web Appendix (not for publication) provides additional material discussed in the unpublished manuscript '*Cutting Fertility? The Effect of Cesarean Deliveries on Subsequent Fertility and Maternal Labor Supply*' by Martin Halla, Harald Mayr, Gerald J. Pruckner, and Pilar García-Gómez.

Part

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A The Changing Likelihood of a CD over Time by Gestational Age and Weekday

In this section, we explore the changes in the likelihood of a CD over time. In particular, we compare the year 1995 (henceforth $p1$) with the year 2007 (henceforth $p2$). From $p1$ to $p2$ the CD rate had increased from about 13 to 27 percent (see also Appendix Figure C.1). In a first step, we explore whether the increase in the CD rate is equally distributed across gestational ages $gest_g$. Therefore, we estimate the following model via a logit estimation:

$$CD_{i(p=1,2)} = \alpha + \sum_{g=36}^{42} \beta_g \cdot gest_g + \sum_{g=36}^{42} \tau_g \cdot gest_g \cdot p2 + x'_{i(p=1,2)} \cdot \gamma + \epsilon_{i(p=1,2)}, \tag{6}$$

Table A.1 summarizes the respective estimation output. Column (I) shows that the increased CD rate is due to higher likelihoods of a CD birth at all gestational ages; however, the strongest increases can be observed for week 37 (plus 12 percentage points), week 38 (plus 16 percentage points) and week 39 (plus 14 percentage points). In a second step, we amend the estimation model in (7) with a three-way interaction between gestational age, $p2$, and weekday (and a two-way interaction between age and weekday):

$$\begin{aligned} CD_{i(p=1,2)} = \alpha + \sum_{g=36}^{42} \beta_g \cdot gest_g + \sum_{g=36}^{42} \tau_g \cdot gest_g \cdot p2 + \sum_{g=36}^{42} \delta_g \cdot gest_g \cdot \text{weekday} \\ + \sum_{g=36}^{42} \kappa_g \cdot gest_g \cdot p2 \cdot \text{weekday} + x'_{i(p=1,2)} \cdot \gamma + \epsilon_{i(p=1,2)}, \end{aligned} \tag{7}$$

This allows to explore whether the increase in the likelihood of a CD on weekdays versus leisure days is equally distributed across gestational ages.

The new estimates of τ_g tend to be lower, and the increases are more equally distributed across gestational ages. The estimates of κ_g show to which degree the likelihood of a CD increased over time on weekdays (as compared to leisure days) by gestational age. Given that planned CDs are exclusively done on weekdays, we can interpret the estimates of κ_g as indication, at which gestational ages planned CDs are done. Our estimation results suggest that planned CDs are predominantly performed in weeks 37 and 38. This is in line with the Austrian guidelines for gynecologists.

Table A.1 – The changing likelihood of a CD over time by gestational age (and weekday)

	(I)		(II)	
	Mfx	s.e.	Mfx	s.e.
Two-way interactions with $p2$				
$gest_{36} \cdot p2$	0.059***	(0.013)	0.037	(0.024)
$gest_{37} \cdot p2$	0.124***	(0.017)	0.053	(0.032)
$gest_{38} \cdot p2$	0.161***	(0.011)	0.090***	(0.024)
$gest_{39} \cdot p2$	0.135***	(0.009)	0.115***	(0.020)
$gest_{40} \cdot p2$	0.077***	(0.007)	0.084***	(0.014)
$gest_{41} \cdot p2$	0.105***	(0.009)	0.122***	(0.016)
$gest_{42} \cdot p2$	0.103***	(0.012)	0.113***	(0.022)
Three-way interactions $p2$				
$gest_{36} \cdot p2 \cdot$ weekday			0.031	(0.029)
$gest_{37} \cdot p2 \cdot$ weekday			0.095**	(0.038)
$gest_{38} \cdot p2 \cdot$ weekday			0.084***	(0.027)
$gest_{39} \cdot p2 \cdot$ weekday			0.022	(0.022)
$gest_{40} \cdot p2 \cdot$ weekday			−0.009	(0.016)
$gest_{41} \cdot p2 \cdot$ weekday			−0.025	(0.019)
$gest_{42} \cdot p2 \cdot$ weekday			−0.014	(0.026)
$\sum_g gest_g$	Yes		Yes	
$\sum_g gest_g \cdot$ weekday	No		Yes	
Socio-economic controls	Yes		Yes	
Number of observations	43,481		43,481	

Notes: This table summarizes logit estimations of the likelihood of a CD. The data include all first-time Austrian mothers who had a single inpatient birth in the year 1995 ($p1$) and the year 2007 ($p2$). The variables $gest_g$ are binary indicators for different gestational ages g, which are measured in weeks. The socio-economic control variables comprise information on mother's age, marital status, educational attainment, occupation, religious denomination, province of residence, and the child's birth weight. For details on these variables, see Table 1). Standard errors in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

B Maternity Ward Data

The University Hospital in Linz provides basic healthcare services for the Linz area and top level medical services for the whole province of Upper Austria. The clinic hosts the biggest Austrian maternity ward with more than 3,500 births per year. For this maternity ward, two datasets are available for the period from 2013 to 2019.

Duty rosters: Information on duty schedules include the physicians on duty, their type (qualified specialist or specialist in education) and medical specialty on a daily basis. More specifically, we observe which and how many doctors work in the operating rooms, the delivery rooms, or the obstetrics departments.

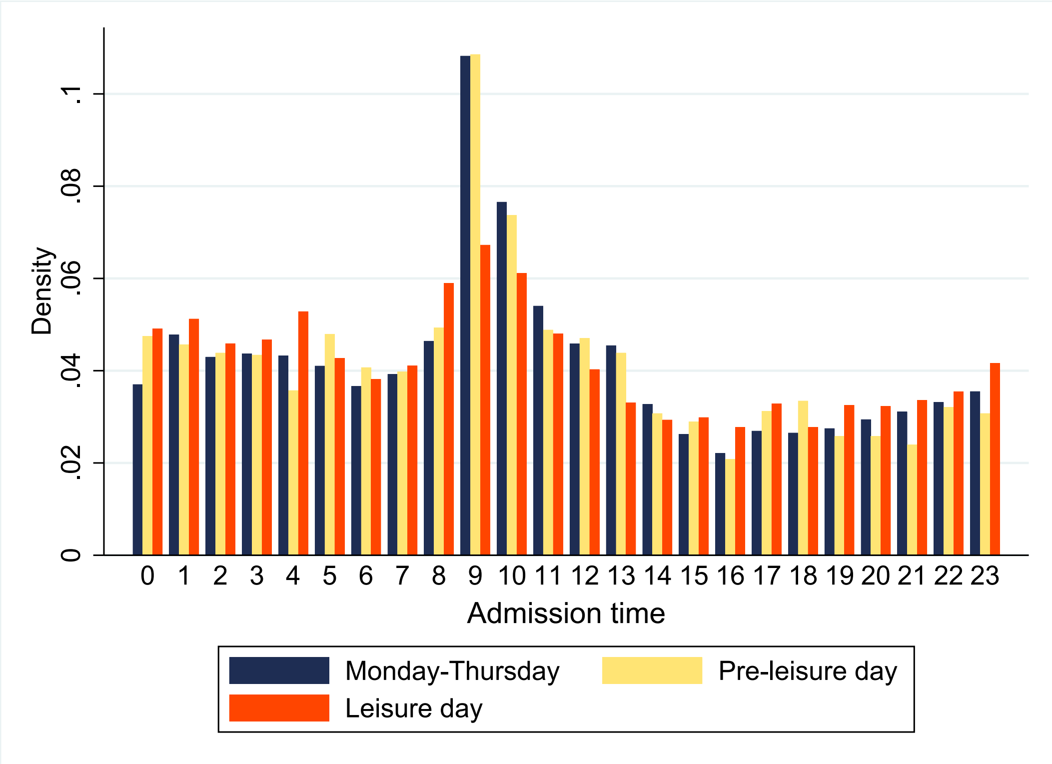
Birth data: The second dataset contains detailed individual-level information on $22,000$ births. The data cover the date and time of hospital admission and birth, the type of birth (spontaneous delivery or sectio), whether the hospital admission was acute or non-acute (planned), and the sex of the child. Finally, a unique mother ID enables us to identify nulliparous mothers.

Table B.1 – Doctor supply per day of week

	(1) # of physicians per birth	(2) # of physicians per birth
Pre-leisure Day	−0.046	−0.046
	(0.028)	(0.027)
Leisure Day	−0.436***	−0.435***
	(0.020)	(0.020)
Year fixed effects	no	yes
Mean of dep. variable	0.933	0.933
# of observations	2,129	2,129

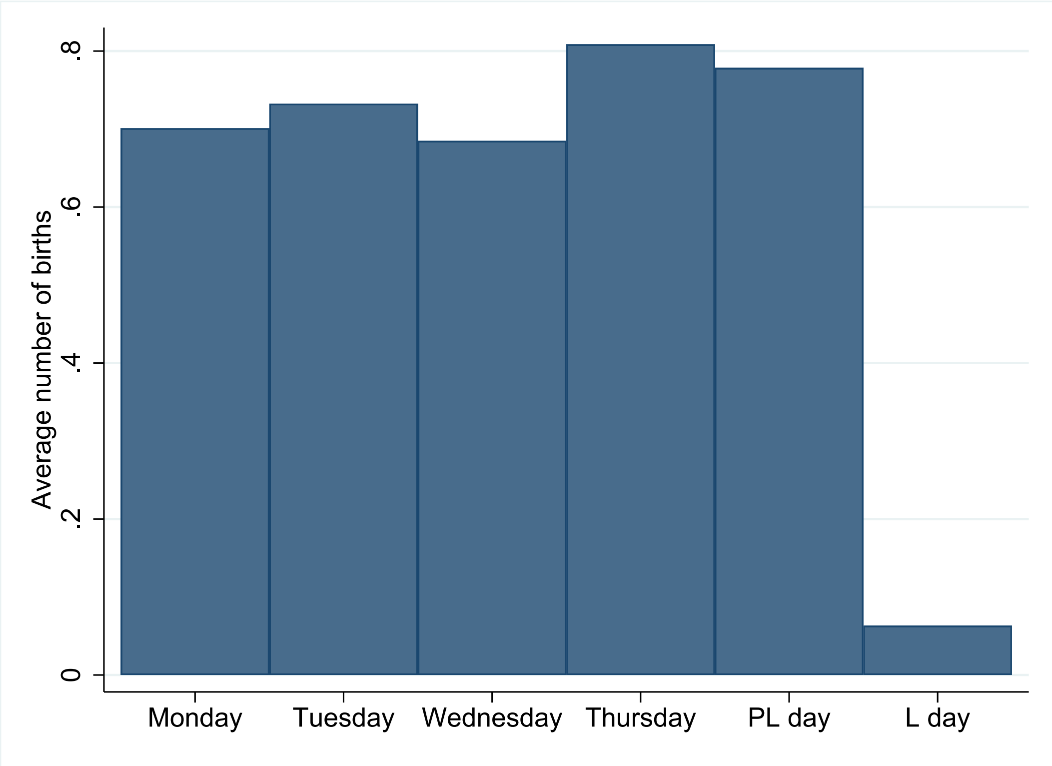
Notes: This table depicts estimation results of the number of available doctors per day and birth at the maternity ward of the University Hospital in Linz based on data from 2013 to 2019. Pre-leisure Day denotes Fridays and working days preceding a public holiday. Leisure Day denotes Saturdays, Sundays, and public holidays. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Figure B.1 – Distribution of admission times



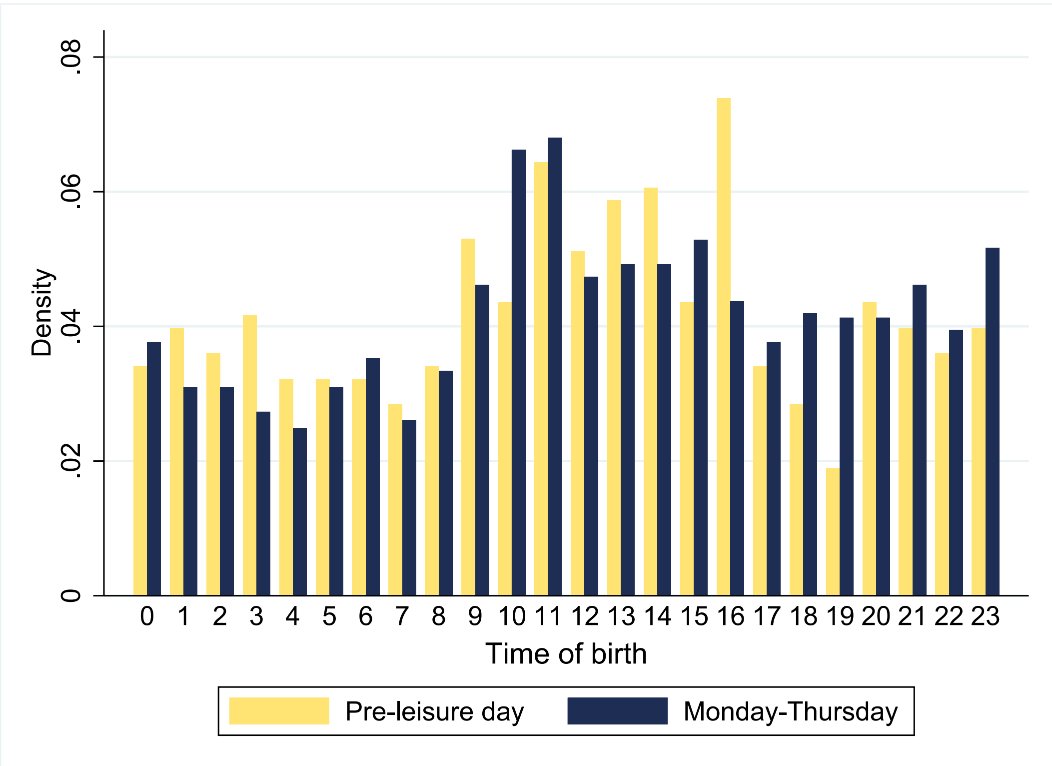
Notes: This figure depicts the distribution of admisssion times for all acute births in the maternity ward of the University Hospital in Linz from 2013 to 2019. Histograms depicted for pre-leisure day, leisure day, and Monday-Thursday births of nulliparous mothers.

Figure B.2 – Average number of planned (non-acute) sectios per day



Notes: This figure depicts the distribution of planned CD of nulliparous mothers over weekdays (Monday-Thursday).

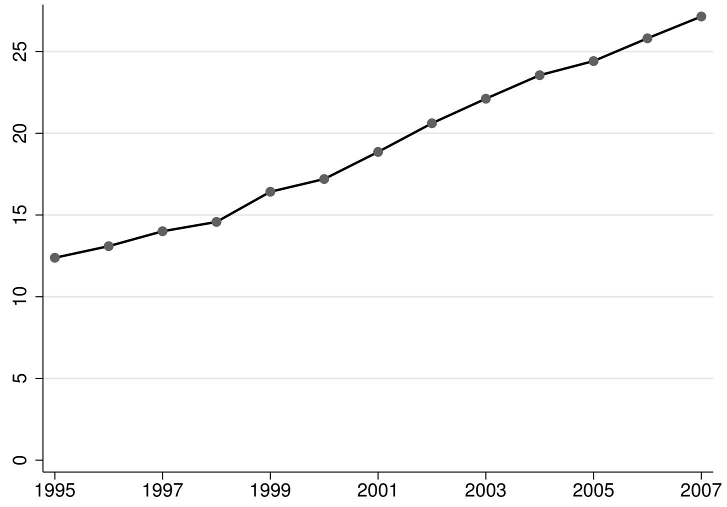
Figure B.3 – Distribution of times of births (acute sectios)



Notes: This figure depicts the distribution of birth times for acute sectios of nulliparous mothers. Histograms depicted for cesarean section births on pre-leisure days and workdays.

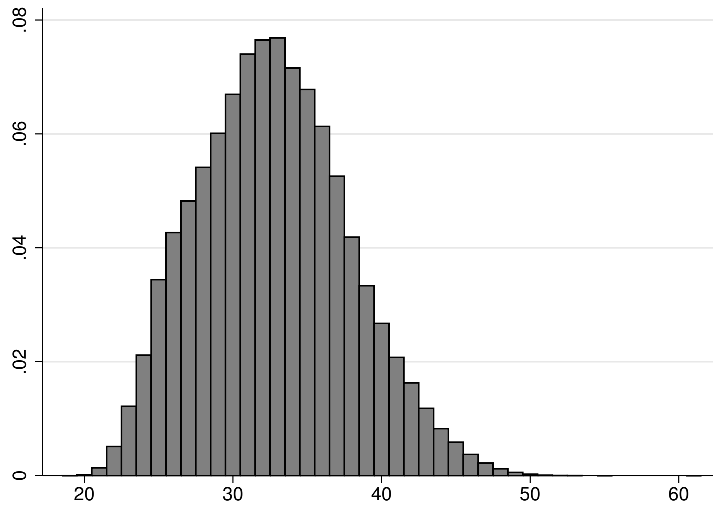
C Additional Tables and Figures

Figure C.1 – CD rate in Austria, 1995-2007



Notes: This figure depicts CD rates in Austria in percent from 1995-2007 (the time period under study).

Figure C.2 – Distribution of mother's age six years after first birth



Notes: This figure depicts the age distribution of mothers in the main sample six years after their first birth.

Table C.1 – Balance checks for birth outcomes and employment history

	(1)	(2)	(3)	(4)
	Leisure Day		Pre-leisure Day	
	No Controls	Controls	No Controls	Controls
Birth weight	0.2536	0.1751	−0.4487**	−0.4157*
	(0.1709)	(0.1683)	(0.2125)	(0.2167)
Gestation length	0.0216***	0.0208***	−0.0105**	−0.0049
	(0.0039)	(0.0039)	(0.0048)	(0.0050)
Mother's employment year −4	−0.0029	−0.0003	0.0021	0.0008
	(0.0019)	(0.0016)	(0.0024)	(0.0021)
Mother's employment year −3	−0.0008	0.0009	0.0028	0.0025
	(0.0017)	(0.0015)	(0.0021)	(0.0019)
Mother's employment year −2	0.0012	0.0020	0.0023	0.0024
	(0.0016)	(0.0014)	(0.0019)	(0.0018)
Mother's employment year −1	0.0013	0.0023*	0.0022	0.0027*
	(0.0014)	(0.0013)	(0.0017)	(0.0016)
Mother's wage year −4	0.0212	0.0879	−0.0011	−0.1209
	(0.1181)	(0.0880)	(0.1444)	(0.1115)
Mother's wage year −3	0.0166	0.0902	0.1158	0.0552
	(0.1123)	(0.0831)	(0.1373)	(0.1055)
Mother's wage year −2	−0.0401	0.0212	0.1252	0.0152
	(0.1093)	(0.0824)	(0.1357)	(0.1071)
Mother's wage year −1	−0.0970	−0.0136	0.1319	−0.0483
	(0.1083)	(0.0832)	(0.1338)	(0.1075)
Father recorded	0.0031	0.0028*	−0.0008	0.0021
	(0.0020)	(0.0015)	(0.0025)	(0.0019)

Notes: Column (1) shows coefficients β_1^y from individual regressions $y = \beta_1^y \cdot leisure + \epsilon_1^y$, where y indicates the row's observable characteristic. Column (2) shows coefficients β_2^y from individual regressions $y = \beta_2^y \cdot leisure + x' \cdot \gamma_y^2 + \epsilon_2^y$, where x includes mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Columns (3) and (4) show the results for regressions of observable characteristics on *preleisure*, where the leisure indicator is included as additional control variable in x. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.2 – Balance checks for health history

	(1)	(2)	(3)	(4)
	Leisure Day		Pre-leisure Day	
	No Controls	Controls	No Controls	Controls
Outpatient attendance year −1	−15.0722***	−11.8693**	−0.5747	−4.7439
	(5.7864)	(5.2087)	(7.3571)	(6.9190)
Outpatient attendance year −2	−7.1347	−6.2117	1.5054	0.6054
	(5.2253)	(5.1397)	(6.1426)	(6.1744)
Medication year −1	2.8848	3.3211	1.3252	3.1340
	(6.8040)	(7.0123)	(5.0156)	(4.8494)
Medication year −2	12.7830	13.7304	1.3281	6.6927
	(10.5734)	(10.9096)	(6.7026)	(6.4549)
Days in general hospitals year −1	−0.0204	−0.0175	−0.1051	−0.0911
	(0.0754)	(0.0747)	(0.0911)	(0.0957)
Days in special clinics year −1	−0.0002	−0.0018	0.0273	0.0278
	(0.0189)	(0.0176)	(0.0218)	(0.0225)
DRG revenue year −1	−14.8008	−16.6100	−18.7368	−14.5910
	(22.8649)	(22.4776)	(31.1097)	(32.2727)

Notes: Column (1) shows coefficients β_1^y from individual regressions $y = \beta_1^y \cdot leisure + \epsilon_1^y$, where y indicates the row's observable characteristic. Column (2) shows coefficients β_2^y from individual regressions $y = \beta_2^y \cdot leisure + x' \cdot \gamma_y^2 + \epsilon_2^y$, where x includes mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Columns (3) and (4) show the results for regressions of observable characteristics on *preleisure*, where the leisure indicator is included as additional control variable in x. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.3 – Bivariate probit results for fertility: all covariates, equations (1) and (2)

	Equation (2)		Equation (1)	
Eq. (2): $leisure_{i(t=0)}$	−0.0497***	(0.0020)		
Eq. (1): $CD_{i(t=0)}$			−0.1225***	(0.0258)
Girl	−0.0238***	(0.0013)	−0.0068***	(0.0018)
Birth year, base group 1995				
1996	0.0019	(0.0041)	0.0040	(0.0041)
1997	0.0134***	(0.0041)	0.0035	(0.0043)
1998	0.0161***	(0.0042)	0.0070*	(0.0041)
1999	0.0355***	(0.0048)	0.0069	(0.0048)
2000	0.0384***	(0.0051)	0.0238***	(0.0046)
2001	0.0454***	(0.0054)	0.0431***	(0.0044)
2002	0.0589***	(0.0051)	0.0434***	(0.0055)
2003	0.0639***	(0.0054)	0.0624***	(0.0057)
2004	0.0791***	(0.0053)	0.0671***	(0.0053)
2005	0.0874***	(0.0054)	0.0808***	(0.0057)
2006	0.0890***	(0.0060)	0.0855***	(0.0056)
2007	0.0973***	(0.0055)	0.0891***	(0.0058)
Birth month, base group January				
February	0.0000	(0.0032)	0.0070*	(0.0042)
March	−0.0034	(0.0036)	0.0136***	(0.0044)
April	0.0023	(0.0031)	0.0106**	(0.0042)
May	0.0033	(0.0032)	0.0150***	(0.0041)
June	0.0019	(0.0035)	0.0112**	(0.0044)
July	0.0021	(0.0034)	0.0109**	(0.0043)
August	0.0041	(0.0026)	0.0147***	(0.0040)
September	0.0030	(0.0032)	0.0114***	(0.0044)
October	0.0047	(0.0033)	0.0123***	(0.0047)
November	0.0070**	(0.0033)	0.0127***	(0.0033)
December	0.0110***	(0.0036)	0.0088*	(0.0049)
Mother's age decile, base group under 19				
19-21	0.0312***	(0.0049)	0.0511***	(0.0064)
22-24	0.0486***	(0.0046)	0.0767***	(0.0083)
25-27	0.0705***	(0.0048)	0.0604***	(0.0091)
28-30	0.0977***	(0.0051)	0.0060	(0.0094)
31-33	0.1293***	(0.0051)	−0.0768***	(0.0093)
34-36	0.1694***	(0.0051)	−0.1801***	(0.0106)
Above 36	0.2186***	(0.0054)	−0.3805***	(0.0140)
Mother's education, base group Compulsory school				
Apprenticeship	−0.0157***	(0.0034)	0.0091**	(0.0037)
Interm. tech. and voc. school	−0.0259***	(0.0037)	0.0333***	(0.0043)
Academic secondary school	−0.0358***	(0.0041)	0.0566***	(0.0052)
Post-secondary college	−0.0438***	(0.0047)	0.1053***	(0.0084)
University	−0.0272***	(0.0041)	0.1727***	(0.0083)
Unknown	−0.0248***	(0.0074)	0.0383***	(0.0096)
Family status, base group Single				
Married	0.0073***	(0.0021)	0.1145***	(0.0041)
Widowed	0.0400	(0.0270)	0.0362	(0.0380)
Divorced	0.0312***	(0.0052)	−0.0245***	(0.0066)
Mother's occupation, base group Self-employed				
Farming	−0.0109	(0.0102)	0.2268***	(0.0148)

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Table continued

White collar	−0.0169***	(0.0046)	0.2108***	(0.0099)
Blue collar	−0.0086	(0.0055)	0.2152***	(0.0105)
Missing	−0.0119**	(0.0054)	0.0777***	(0.0108)
Unknown	−0.0164**	(0.0082)	0.2019***	(0.0123)
Mother's religion, base group Roman Catholic				
Protestant	0.0068	(0.0047)	−0.0170**	(0.0074)
Old Catholic	0.0443**	(0.0191)	−0.1135***	(0.0371)
Jewish	−0.0095	(0.0255)	0.0637**	(0.0298)
Muslim	−0.0161*	(0.0083)	0.0291**	(0.0139)
Other	0.0130*	(0.0073)	0.0363**	(0.0168)
Unknown	0.0040	(0.0029)	−0.0787***	(0.0046)
Province of residence, base group Burgenland				
Carinthia	−0.0423***	(0.0089)	0.0080	(0.0135)
Lower Austria	−0.0253***	(0.0089)	0.0429***	(0.0148)
Upper Austria	−0.0530***	(0.0105)	0.0664***	(0.0145)
Salzburg	−0.0730***	(0.0106)	0.0810***	(0.0160)
Styria	0.0025	(0.0085)	0.0239*	(0.0125)
Tyrol	−0.0428***	(0.0103)	0.0762***	(0.0121)
Vorarlberg	−0.0802***	(0.0077)	0.1272***	(0.0090)
Vienna	−0.0386***	(0.0071)	−0.0381***	(0.0121)
Mean of dependent variable	0.183		0.529	
Number of observations	282,831		282,831	

Notes: This table summarizes bivariate probit estimation results for fertility six years after first birth based on the leisure IV and including all covariates. Estimations provide the effect of a CD at parity 0 on the likelihood of a second birth (Equation (1)) and the effect of the leisure instrument on the likelihood of a CD at parity 0 (Equation (2)). Reported estimates are average marginal effects. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.4 – Probit results for fertility

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
Eq. (1): $CD_{i(t=0)}$	−0.0017***	−0.0341***	−0.0524***	−0.0587***	−0.0583***	−0.0575***
	(0.0005)	(0.0021)	(0.0022)	(0.0024)	(0.0027)	(0.0027)
Girl	−0.0003	−0.0032**	−0.0062***	−0.0076***	−0.0067***	−0.0053***
	(0.0003)	(0.0013)	(0.0017)	(0.0017)	(0.0017)	(0.0017)
Birth year, base 1995	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
Birth month, base January	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
Mother's age, base under 19						
19-21	−0.0007	0.0259***	0.0529***	0.0556***	0.0537***	0.0498***
	(0.0006)	(0.0041)	(0.0062)	(0.0071)	(0.0068)	(0.0064)
22-24	−0.0029***	0.0319***	0.0856***	0.0953***	0.0864***	0.0745***
	(0.0006)	(0.0040)	(0.0071)	(0.0086)	(0.0084)	(0.0081)
25-27	−0.0044***	0.0208***	0.0803***	0.0868***	0.0752***	0.0568***
	(0.0007)	(0.0042)	(0.0072)	(0.0090)	(0.0090)	(0.0087)
28-30	−0.0061***	0.0036	0.0462***	0.0433***	0.0233***	0.0002
	(0.0008)	(0.0043)	(0.0069)	(0.0085)	(0.0087)	(0.0085)
31-33	−0.0063***	−0.0221***	−0.0019	−0.0212***	−0.0533***	−0.0856***
	(0.0009)	(0.0040)	(0.0066)	(0.0080)	(0.0078)	(0.0077)
34-36	−0.0054***	−0.0458***	−0.0706***	−0.1092***	−0.1532***	−0.1930***
	(0.0009)	(0.0049)	(0.0069)	(0.0077)	(0.0079)	(0.0080)
Above 36	−0.0078***	−0.1039***	−0.2185***	−0.2896***	−0.3528***	−0.3996***
	(0.0012)	(0.0061)	(0.0079)	(0.0095)	(0.0094)	(0.0092)
Mother's education, base Compulsory school						
Apprenticeship	−0.0016***	−0.0085***	0.0053	0.0060	0.0091**	0.0101***
	(0.0004)	(0.0028)	(0.0037)	(0.0039)	(0.0038)	(0.0038)
Interm. tech. and voc. school	−0.0018***	−0.0006	0.0160***	0.0225***	0.0306***	0.0352***

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Table continued

	(0.0006)	(0.0036)	(0.0044)	(0.0043)	(0.0041)	(0.0044)
Academic secondary school	−0.0027***	0.0076**	0.0380***	0.0467***	0.0552***	0.0593***
	(0.0007)	(0.0032)	(0.0050)	(0.0050)	(0.0049)	(0.0049)
Post-secondary college	−0.0039***	0.0141**	0.0672***	0.0878***	0.1041***	0.1089***
	(0.0010)	(0.0057)	(0.0077)	(0.0074)	(0.0078)	(0.0078)
University	−0.0031***	0.0490***	0.1184***	0.1470***	0.1674***	0.1757***
	(0.0010)	(0.0044)	(0.0073)	(0.0077)	(0.0080)	(0.0077)
Unknown	−0.0013	0.0049	0.0230**	0.0314***	0.0451***	0.0400***
	(0.0014)	(0.0065)	(0.0096)	(0.0095)	(0.0100)	(0.0096)
Family status, base Single	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
Mother's occupation, base Self-employed	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
Mother's religion, base Roman Catholic	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
Province of residence, base Burgenland	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*	*Yes*
Mean of dependent variable	0.006	0.151	0.332	0.429	0.489	0.529
Number of observations	282,521	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes probit estimation results for fertility. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 6 years after the first birth. Reported estimates are average marginal effects. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.5 – Bivariate probit results for fertility: subsample of births with gestation length of at least 40 weeks

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0034	0.0366	−0.0277	−0.0858**	−0.0858**	−0.1182***
	(0.0067)	(0.0265)	(0.0373)	(0.0346)	(0.0376)	(0.0374)
Eq. (2): $leisure_{i(t=0)}$	−0.0297***	−0.0297***	−0.0297***	−0.0297***	−0.0297***	−0.0297***
	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.0482	−0.1682***	−0.0355	0.0440	0.0445	0.0960*
	(0.2268)	(0.0611)	(0.0590)	(0.0520)	(0.0560)	(0.0565)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0025	0.0411	−0.0226	−0.0844**	−0.0871*	−0.1196***
	(0.0066)	(0.0320)	(0.0404)	(0.0397)	(0.0474)	(0.0453)
Eq. (3): $preleisure_{i(t=0)}$	0.0061***	0.0061***	0.0061***	0.0060***	0.0060***	0.0060***
	(0.0022)	(0.0021)	(0.0022)	(0.0022)	(0.0022)	(0.0022)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.0174	−0.1786**	−0.0436	0.0419	0.0463	0.0980
	(0.2246)	(0.0734)	(0.0638)	(0.0596)	(0.0705)	(0.0682)
Mean of dependent variable	0.006	0.152	0.332	0.429	0.490	0.530
Number of observations	223,380	223,380	223,380	223,380	223,380	223,380

Notes: This table summarizes bivariate probit estimation results for deliveries in week 40 and later, based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.6 – Bivariate probit results: the effect of a CD at parity 0 on the likelihood of a second birth (leisure days excluded)

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PRE-LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	0.0004	0.0248	−0.0567	−0.1272***	−0.1400***	−0.1390***
	(0.0065)	(0.0284)	(0.0454)	(0.0446)	(0.0453)	(0.0476)
Eq. (3): $preleisure_{i(t=0)}$	0.0084***	0.0084***	0.0084***	0.0084***	0.0084***	0.0083***
	(0.0025)	(0.0024)	(0.0025)	(0.0024)	(0.0024)	(0.0024)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	−0.0536	−0.1441**	0.0055	0.1051	0.1243*	0.1247*
	(0.2201)	(0.0685)	(0.0743)	(0.0698)	(0.0707)	(0.0748)
Mean of dependent variable	0.006	0.150	0.331	0.427	0.487	0.527
Number of observations	199,550	199,550	199,550	199,550	199,550	199,550

Notes: This table summarizes bivariate probit estimation results for the pre-leisure IV (joint estimation of Eq. (1) and (3)) when all births on leisure days are excluded. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 6 years after the first birth (Eq. (1)). The second row includes the effects of the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.7 – Bivariate Probit results (education interaction): the effect of a CD at parity 0 on the likelihood of a second birth

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0069	−0.0099	−0.0837***	−0.1176***	−0.1219***	−0.1311***
	(0.0070)	(0.0187)	(0.0248)	(0.0234)	(0.0241)	(0.0249)
Interaction high education	0.0011	0.0066*	0.0106**	0.0082*	0.0092*	0.0092**
	(0.0009)	(0.0038)	(0.0046)	(0.0049)	(0.0049)	(0.0046)
Eq. (2): $leisure_{i(t=0)}$	−0.0473***	−0.0473***	−0.0474***	−0.0473***	−0.0473***	−0.0473***
	(0.0030)	(0.0030)	(0.0029)	(0.0029)	(0.0029)	(0.0029)
Interaction high education	−0.0046	−0.0045	−0.0045	−0.0047	−0.0047	−0.0047
	(0.0035)	(0.0035)	(0.0034)	(0.0034)	(0.0034)	(0.0034)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1565	−0.0686	0.0421	0.0842**	0.0896**	0.1058***
	(0.2016)	(0.0457)	(0.0409)	(0.0366)	(0.0372)	(0.0387)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0025	0.0131	−0.0766**	−0.1282***	−0.1314***	−0.1402***
	(0.0066)	(0.0247)	(0.0318)	(0.0310)	(0.0336)	(0.0333)
Interaction high education	0.0010	0.0063*	0.0105**	0.0082*	0.0093*	0.0092**
	(0.0009)	(0.0038)	(0.0046)	(0.0049)	(0.0048)	(0.0045)
Eq. (3): $preleisure_{i(t=0)}$	0.0081***	0.0081***	0.0081***	0.0080***	0.0081***	0.0080***
	(0.0024)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
Interaction high education	−0.0045	−0.0045	−0.0045	−0.0047	−0.0046	−0.0047
	(0.0035)	(0.0035)	(0.0035)	(0.0034)	(0.0034)	(0.0034)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.0145	−0.1232**	0.0306	0.1005**	0.1042**	0.1197**
	(0.2218)	(0.0591)	(0.0517)	(0.0481)	(0.0519)	(0.0518)
Mean of dependent variable	0.006	0.151	0.332	0.429	0.489	0.529
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes bivariate probit estimation results for the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)), including interactions with a high education indicator in both equations. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 6 years after the first birth (Eq. (1)). The third row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.8 – Bivariate probit results: fertility for mothers observable at least 10 years

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$	(VII) $F_{i(t+7)}$	(VIII) $F_{i(t+8)}$	(IX) $F_{i(t+9)}$	(X) $F_{i(t+10)}$
PANEL A: LEISURE IV										
Eq. (1): $CD_{i(t=0)}$	−0.0077	−0.0361	−0.1080***	−0.1472***	−0.1552***	−0.1759***	−0.1678***	−0.1836***	−0.1884***	−0.1867***
	(0.0074)	(0.0231)	(0.0286)	(0.0290)	(0.0287)	(0.0277)	(0.0278)	(0.0273)	(0.0282)	(0.0287)
Eq. (2): $leisure_{i(t=0)}$	−0.0440***	−0.0440***	−0.0441***	−0.0439***	−0.0440***	−0.0440***	−0.0440***	−0.0440***	−0.0440***	−0.0440***
	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1682	−0.0067	0.0893*	0.1377***	0.1496***	0.1823***	0.1737***	0.1989***	0.2098***	0.2100***
	(0.2008)	(0.0542)	(0.0468)	(0.0446)	(0.0435)	(0.0425)	(0.0429)	(0.0431)	(0.0452)	(0.0463)
PANEL B: PRELEISURE IV										
Eq. (1): $CD_{i(t=0)}$	−0.0065	−0.0346	−0.1140***	−0.1674***	−0.1713***	−0.1861***	−0.1771***	−0.1909***	−0.2007***	−0.1930***
	(0.0095)	(0.0293)	(0.0328)	(0.0321)	(0.0344)	(0.0329)	(0.0323)	(0.0326)	(0.0331)	(0.0355)
Eq. (3): $preleisure_{i(t=0)}$	0.0099***	0.0098***	0.0099***	0.0095***	0.0095***	0.0095***	0.0095***	0.0095***	0.0095***	0.0095***
	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1328	−0.0101	0.0991*	0.1688***	0.1741***	0.1980***	0.1881***	0.2103***	0.2295***	0.2200***
	(0.2706)	(0.0687)	(0.0533)	(0.0493)	(0.0522)	(0.0503)	(0.0495)	(0.0511)	(0.0529)	(0.0570)
Mean of dependent variable	0.006	0.161	0.332	0.424	0.483	0.523	0.552	0.572	0.586	0.597
Number of observations	204,392	204,392	204,392	204,392	204,392	204,392	204,392	204,392	204,392	204,392

Notes: This table summarizes bivariate probit estimation results for mothers who we observe up to at least 10 years after first birth based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(X) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 10 years after the first birth (Eq. (1)). The second row includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.9 – Bivariate probit results: fertility for mothers at least 30 at first birth and observable at least 10 years

	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$	(VII) $F_{i(t+7)}$	(VIII) $F_{i(t+8)}$	(IX) $F_{i(t+9)}$	(X) $F_{i(t+10)}$
PANEL A: LEISURE IV									
Eq. (1): $CD_{i(t=0)}$	−0.0723*	−0.1340***	−0.1904***	−0.1898***	−0.2105***	−0.1942***	−0.1980***	−0.1991***	−0.2077***
	(0.0419)	(0.0468)	(0.0461)	(0.0494)	(0.0474)	(0.0475)	(0.0488)	(0.0495)	(0.0482)
Eq. (2): $leisure_{i(t=0)}$	−0.0578***	−0.0577***	−0.0576***	−0.0577***	−0.0577***	−0.0577***	−0.0577***	−0.0577***	−0.0577***
	(0.0042)	(0.0042)	(0.0042)	(0.0042)	(0.0042)	(0.0042)	(0.0042)	(0.0042)	(0.0042)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1251	0.1600*	0.2312***	0.2252***	0.2558***	0.2253***	0.2296***	0.2315***	0.2465***
	(0.1182)	(0.0873)	(0.0824)	(0.0864)	(0.0833)	(0.0819)	(0.0844)	(0.0859)	(0.0839)
PANEL B: PRELEISURE IV									
Eq. (1): $CD_{i(t=0)}$	−0.1823**	−0.2494***	−0.2830***	−0.2707***	−0.2706***	−0.2605***	−0.2566***	−0.2610***	−0.2630***
	(0.0919)	(0.0649)	(0.0514)	(0.0518)	(0.0476)	(0.0477)	(0.0509)	(0.0506)	(0.0503)
Eq. (3): $preleisure_{i(t=0)}$	0.0103*	0.0105**	0.0099**	0.0105**	0.0104**	0.0108**	0.0103**	0.0104**	0.0103**
	(0.0054)	(0.0050)	(0.0049)	(0.0050)	(0.0050)	(0.0051)	(0.0051)	(0.0051)	(0.0051)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.4207*	0.3836***	0.4027***	0.3708***	0.3639***	0.3427***	0.3332***	0.3412***	0.3449***
	(0.2288)	(0.1271)	(0.0976)	(0.0949)	(0.0864)	(0.0854)	(0.0910)	(0.0909)	(0.0902)
Mean of dependent variable	0.128	0.276	0.347	0.386	0.409	0.423	0.432	0.437	0.440
Number of observations	53,665	53,665	53,665	53,665	53,665	53,665	53,665	53,665	53,665

Notes: This table summarizes bivariate probit estimation results for older mothers based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(X) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 10 years after the first birth (Eq. (1)). The second row includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Mothers are included in the sample if they are at least 30 years old at first birth and can be observed up to at least 10 subsequent years. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.10 – Bivariate probit results: the effect of a CD at parity 0 on the likelihood of a third birth

	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$	(VII) $F_{i(t+7)}$	(VIII) $F_{i(t+8)}$	(IX) $F_{i(t+9)}$	(X) $F_{i(t+10)}$
Panel A: Leisure IV								
Eq. (1): $CD_{i(t=0)}$	−0.0152	−0.0028	−0.0220	−0.0298*	−0.0459**	−0.0639***	−0.0582**	−0.0622***
	(0.0193)	(0.0128)	(0.0145)	(0.0162)	(0.0193)	(0.0216)	(0.0235)	(0.0236)
Eq. (2): $leisure_{i(t=0)}$	−0.0441***	−0.0440***	−0.0440***	−0.0440***	−0.0440***	−0.0440***	−0.0440***	−0.0439***
	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)	(0.0021)
$\rho : Corr(\varepsilon_{it}^{p}, \epsilon_{it})$	0.3464	−0.0455	0.0723	0.0715	0.1135	0.1571**	0.1139	0.1130
	(0.3337)	(0.1399)	(0.0957)	(0.0797)	(0.0779)	(0.0759)	(0.0759)	(0.0715)
Panel B: Preleisure IV								
Eq. (1): $CD_{i(t=0)}$	0.0077	−0.0005	−0.0316*	−0.0424*	−0.0586*	−0.0819**	−0.0987**	−0.1213***
	(0.0122)	(0.0129)	(0.0182)	(0.0232)	(0.0325)	(0.0401)	(0.0445)	(0.0436)
Eq. (3): $preleisure_{i(t=0)}$	0.0097***	0.0098***	0.0099***	0.0099***	0.0098***	0.0099***	0.0099***	0.0099***
	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)
$\rho : Corr(\varepsilon_{it}^{p}, \epsilon_{it})$	−0.2470	−0.0697	0.1322	0.1307	0.1636	0.2188	0.2408*	0.2858**
	(0.2375)	(0.1374)	(0.1134)	(0.1074)	(0.1257)	(0.1352)	(0.1365)	(0.1238)
Mean of dependent variable	0.006	0.021	0.041	0.060	0.075	0.091	0.104	0.115
Number of observations	204,392	204,392	204,392	204,392	204,392	204,392	204,392	204,392

Notes: This table summarizes bivariate probit estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (III)-(X) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 3 to 10 years after the first birth (Eq. (1)). The second row includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Mothers are included in the sample if they can be observed up to at least 10 subsequent years. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.11 – Tobit results: the effect of a CD at parity 0 on mother's daily earnings

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	1.2700	2.3495	5.4209**	2.5748	2.4020	2.1751
	(1.6827)	(2.6117)	(2.6464)	(2.5173)	(2.6515)	(2.5658)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−5.4428	−7.9609	−4.5757	−16.1326	−22.9941	−17.0197
	(12.7331)	(18.3036)	(19.2993)	(18.5062)	(20.6145)	(21.7235)
Mean of dependent variable	7.440	20.712	24.387	26.808	29.785	32.606
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes estimation results from a Tobit type 1 model with a control function approach based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Residuals from estimating Eq. (2)/Eq. (3) with OLS are included as a control variable in the Tobit type 1 model. Columns (I)-(VI) provide effects of separate Tobit estimations for the effect of a CD at parity 0 on mother's daily earnings in EUR 1 to 6 years after the first birth (Eq. (1)). The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Bootstrapped standard errors in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.12 – 2SLS results: the effect of a CD at parity 0 on mother's daily earnings

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	2.0077	1.7084	4.8825*	2.2115	2.5816	2.5936
	(1.9518)	(3.2869)	(2.9575)	(2.7325)	(2.9247)	(2.9298)
Eq. (2): $leisure_{i(t=0)}$	−0.0480 ***	−0.0480 ***	−0.0480 ***	−0.0480 ***	−0.0480 ***	−0.0480 ***
	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	3.0450	−7.3464	−13.3815	−19.7991	−21.1140	−14.8010
	(14.3374)	(18.4462)	(20.7527)	(22.1051)	(21.1298)	(22.7971)
Eq. (3): $preleisure_{i(t=0)}$	0.0084 ***	0.0084 ***	0.0084 ***	0.0084 ***	0.0084 ***	0.0084 ***
	(0.0025)	(0.0025)	(0.0025)	(0.0025)	(0.0025)	(0.0025)
Mean of dependent variable	7.440	20.712	24.387	26.808	29.785	32.606
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes 2SLS estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the mother's daily earnings in EUR 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.13 – Bivariate Probit results (education interaction): the effect of a CD at parity 0 on the likelihood of maternal employment

	(I) $E_{i(t+1)}$	(II) $E_{i(t+2)}$	(III) $E_{i(t+3)}$	(IV) $E_{i(t+4)}$	(V) $E_{i(t+5)}$	(VI) $E_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0721***	0.0243	0.1307***	0.1320***	0.1014***	0.0727***
	(0.0214)	(0.0335)	(0.0346)	(0.0356)	(0.0340)	(0.0279)
Interaction high education	0.0094**	0.0124***	0.0041	0.0107**	0.0127**	0.0134***
	(0.0040)	(0.0048)	(0.0047)	(0.0049)	(0.0050)	(0.0047)
Eq. (2): $leisure_{i(t=0)}$	−0.0472***	−0.0474***	−0.0471***	−0.0471***	−0.0473***	−0.0472***
	(0.0030)	(0.0030)	(0.0029)	(0.0030)	(0.0029)	(0.0030)
Interaction high education	−0.0044	−0.0045	−0.0047	−0.0046	−0.0044	−0.0046
	(0.0034)	(0.0035)	(0.0034)	(0.0035)	(0.0034)	(0.0035)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1555***	−0.0143	−0.1481***	−0.1623***	−0.1342***	−0.1024**
	(0.0432)	(0.0491)	(0.0509)	(0.0529)	(0.0507)	(0.0441)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.2083***	0.0275	0.1564***	0.1874***	0.1451***	0.1024*
	(0.0641)	(0.0698)	(0.0524)	(0.0484)	(0.0519)	(0.0617)
Interaction high education	0.0079*	0.0124**	0.0036	0.0094*	0.0117**	0.0128**
	(0.0043)	(0.0048)	(0.0047)	(0.0050)	(0.0051)	(0.0050)
Eq. (3): $preleisure_{i(t=0)}$	0.0087***	0.0081***	0.0079***	0.0080***	0.0079***	0.0079***
	(0.0022)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
Interaction high education	−0.0041	−0.0044	−0.0048	−0.0047	−0.0043	−0.0046
	(0.0033)	(0.0035)	(0.0034)	(0.0035)	(0.0034)	(0.0035)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.4432***	−0.0189	−0.1864**	−0.2461***	−0.2011**	−0.1496
	(0.1360)	(0.1031)	(0.0787)	(0.0739)	(0.0795)	(0.0981)
Mean of dependent variable	0.214	0.448	0.511	0.549	0.602	0.652
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes bivariate probit estimation results for the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)), including interactions with a high education indicator in both equations. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of maternal employment 1 to 6 years after the first birth (Eq. (1)). The third row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.14 – Bivariate probit results: the effect of a CD at parity 0 on the likelihood of part-time employment of highly-educated mothers

	(I) $Epart_{i(t+1)}$	(II) $Epart_{i(t+2)}$	(III) $Epart_{i(t+3)}$	(IV) $Epart_{i(t+4)}$	(V) $Epart_{i(t+5)}$	(VI) $Epart_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0240	0.0007	0.0020	0.0076	−0.0067	−0.0152
	(0.0304)	(0.0352)	(0.0363)	(0.0371)	(0.0383)	(0.0349)
Eq. (2): $leisure_{i(t=0)}$	−0.0546***	−0.0546***	−0.0546***	−0.0546***	−0.0546***	−0.0546***
	(0.0028)	(0.0028)	(0.0028)	(0.0028)	(0.0028)	(0.0028)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1202	0.0125	0.0142	0.0004	0.0119	0.0170
	(0.1198)	(0.0765)	(0.0787)	(0.0813)	(0.0790)	(0.0696)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	0.0072	0.0160	−0.0379	0.0046	0.0470	−0.0292
	(0.0350)	(0.0454)	(0.0524)	(0.0484)	(0.0508)	(0.0732)
Eq. (3): $preleisure_{i(t=0)}$	0.0065**	0.0065**	0.0064*	0.0065**	0.0065**	0.0066**
	(0.0033)	(0.0033)	(0.0033)	(0.0033)	(0.0033)	(0.0033)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	−0.0029	−0.0206	0.1005	0.0067	−0.0973	0.0446
	(0.1351)	(0.0969)	(0.1141)	(0.1053)	(0.1017)	(0.1458)
Mean of dependent variable	0.081	0.188	0.190	0.187	0.208	0.226
Number of observations	114,210	114,210	114,210	114,210	114,210	114,210

Notes: This table summarizes bivariate probit estimation results for the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of part-time employment of highly-educated mothers 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. A mother is employed part-time after treatment if she earns less than 75 % of her pre-treatment salary. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.15 – Bivariate probit results: the effect of a CD at parity 0 on the likelihood of full-time employment of highly-educated mothers

	(I) $Efull_{i(t+1)}$	(II) $Efull_{i(t+2)}$	(III) $Efull_{i(t+3)}$	(IV) $Efull_{i(t+4)}$	(V) $Efull_{i(t+5)}$	(VI) $Efull_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	0.0151	0.0874**	0.1325***	0.1059***	0.0533	0.0437
	(0.0173)	(0.0377)	(0.0337)	(0.0376)	(0.0407)	(0.0397)
Eq. (2): $leisure_{i(t=0)}$	−0.0546***	−0.0545***	−0.0546***	−0.0545***	−0.0546***	−0.0546***
	(0.0028)	(0.0028)	(0.0028)	(0.0028)	(0.0028)	(0.0028)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	−0.0227	−0.1199*	−0.1710***	−0.1212**	−0.0420	−0.0234
	(0.0548)	(0.0675)	(0.0569)	(0.0609)	(0.0630)	(0.0603)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0437	0.1003**	0.1468***	0.1272***	0.0422	0.0429
	(0.0308)	(0.0410)	(0.0379)	(0.0440)	(0.0532)	(0.0539)
Eq. (3): $preleisure_{i(t=0)}$	0.0068**	0.0066**	0.0064**	0.0066**	0.0065**	0.0066**
	(0.0033)	(0.0033)	(0.0032)	(0.0032)	(0.0033)	(0.0033)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1706*	−0.1427*	−0.1951***	−0.1558**	−0.0248	−0.0221
	(0.0991)	(0.0736)	(0.0643)	(0.0713)	(0.0819)	(0.0816)
Mean of dependent variable	0.116	0.290	0.327	0.363	0.408	0.447
Number of observations	114,210	114,210	114,210	114,210	114,210	114,210

Notes: This table summarizes bivariate probit estimation results for the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of full-time employment of highly-educated mothers 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. A mother is employed full-time after treatment if she earns more than 75 % of her pre-treatment salary. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.16 – Bivariate probit results: employment for highly-educated mothers observable at least 10 years

	(I) $E_{i(t+1)}$	(II) $E_{i(t+2)}$	(III) $E_{i(t+3)}$	(IV) $E_{i(t+4)}$	(V) $E_{i(t+5)}$	(VI) $E_{i(t+6)}$	(VII) $E_{i(t+7)}$	(VIII) $E_{i(t+8)}$	(IX) $E_{i(t+9)}$	(X) $E_{i(t+10)}$
PANEL A: LEISURE IV										
Eq. (1): $CD_{i(t=0)}$	−0.0534	0.1177**	0.2353***	0.2143***	0.0878	0.0400	0.0707	0.0589	−0.0206	−0.0696
	(0.0414)	(0.0585)	(0.0514)	(0.0551)	(0.0578)	(0.0628)	(0.0598)	(0.0609)	(0.0584)	(0.0669)
Eq. (2): $leisure_{i(t=0)}$	−0.0470***	−0.0469***	−0.0469***	−0.0465***	−0.0469***	−0.0471***	−0.0472***	−0.0470***	−0.0471***	−0.0470***
	(0.0028)	(0.0027)	(0.0027)	(0.0028)	(0.0028)	(0.0027)	(0.0027)	(0.0028)	(0.0027)	(0.0027)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1171	−0.1387	−0.3073***	−0.2766***	−0.0991	−0.0323	−0.1036	−0.0944	0.0494	0.1537
	(0.0759)	(0.0863)	(0.0806)	(0.0886)	(0.0897)	(0.1019)	(0.1050)	(0.1129)	(0.1130)	(0.1329)
PANEL B: PRELEISURE IV										
Eq. (1): $CD_{i(t=0)}$	−0.1117	0.2142***	0.2571***	0.2904***	0.1893**	0.0653	0.0486	0.1516	0.0181	−0.0787
	(0.0841)	(0.0624)	(0.0654)	(0.0565)	(0.0840)	(0.1292)	(0.1098)	(0.1265)	(0.1357)	(0.1035)
Eq. (3): $preleisure_{i(t=0)}$	0.0094***	0.0087**	0.0084**	0.0082**	0.0081**	0.0086**	0.0088***	0.0092***	0.0089***	0.0087***
	(0.0032)	(0.0034)	(0.0034)	(0.0033)	(0.0034)	(0.0034)	(0.0033)	(0.0033)	(0.0034)	(0.0033)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.2266	−0.2850***	−0.3420***	−0.4002***	−0.2598*	−0.0736	−0.0652	−0.2685	−0.0259	0.1719
	(0.1580)	(0.0967)	(0.1043)	(0.0954)	(0.1356)	(0.2115)	(0.1903)	(0.2400)	(0.2661)	(0.2049)
Mean of dependent variable	0.256	0.508	0.543	0.581	0.639	0.692	0.728	0.759	0.784	0.804
Number of observations	94,920	94,920	94,920	94,920	94,920	94,920	94,920	94,920	94,920	94,920

Notes: This table summarizes bivariate probit estimation results for highly-educated mothers who we observe up to at least 10 years after first birth based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(X) provide separate estimations for the effect of a CD at parity 0 on the likelihood of maternal employment 1 to 10 years after the first birth (Eq. (1)). The second row includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.17 – 2SLS results: the effect of a CD at parity 0 on the likelihood of a second birth

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0094	−0.0458	−0.0890**	−0.0943**	−0.0989**	−0.1070**
	(0.0063)	(0.0300)	(0.0392)	(0.0423)	(0.0425)	(0.0433)
Eq. (2): $leisure_{i(t=0)}$	−0.0480***	−0.0480***	−0.0480***	−0.0480***	−0.0480***	−0.0480***
	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0412	−0.1266	−0.1329	0.1062	0.1357	0.1664
	(0.0416)	(0.2146)	(0.2980)	(0.2981)	(0.2979)	(0.2859)
Eq. (3): $preleisure_{i(t=0)}$	0.0084***	0.0084***	0.0084***	0.0084***	0.0084***	0.0084***
	(0.0025)	(0.0025)	(0.0025)	(0.0025)	(0.0025)	(0.0025)
Mean of dependent variable	0.006	0.151	0.332	0.429	0.489	0.529
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes 2SLS estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.18 – 2SLS results: the effect of a CD at parity 0 on the likelihood of maternal employment

	(I) $E_{i(t+1)}$	(II) $E_{i(t+2)}$	(III) $E_{i(t+3)}$	(IV) $E_{i(t+4)}$	(V) $E_{i(t+5)}$	(VI) $E_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	0.0112	0.0259	0.1011 ∗∗	0.0798∗	0.0660	0.0530
	(0.0344)	(0.0432)	(0.0417)	(0.0414)	(0.0418)	(0.0350)
Eq. (2): $leisure_{i(t=0)}$	−0.0480 ∗∗∗	−0.0480 ∗∗∗	−0.0480 ∗∗∗	−0.0480 ∗∗∗	−0.0480 ∗∗∗	−0.0480 ∗∗∗
	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.6326∗	−0.2677	−0.1678	−0.0779	−0.2947	−0.3269
	(0.3321)	(0.3250)	(0.3312)	(0.3087)	(0.3153)	(0.3203)
Eq. (3): $preleisure_{i(t=0)}$	0.0084 ∗∗∗	0.0084 ∗∗∗	0.0084 ∗∗∗	0.0084 ∗∗∗	0.0084 ∗∗∗	0.0084 ∗∗∗
	(0.0025)	(0.0025)	(0.0025)	(0.0025)	(0.0025)	(0.0025)
Mean of dependent variable	0.214	0.448	0.511	0.549	0.602	0.652
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes 2SLS estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of maternal employment 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.19 – 2SLS results with district FEs: the effect of a CD at parity 0 on the likelihood of a second birth

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0097	−0.0462	−0.0874**	−0.0913**	−0.0959**	−0.1040**
	(0.0063)	(0.0299)	(0.0389)	(0.0421)	(0.0422)	(0.0431)
Eq. (2): $leisure_{i(t=0)}$	−0.0479***	−0.0479***	−0.0479***	−0.0479***	−0.0479***	−0.0479***
	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0401	−0.1269	−0.1368	0.1005	0.1249	0.1549
	(0.0412)	(0.2115)	(0.2957)	(0.2944)	(0.2927)	(0.2809)
Eq. (3): $preleisure_{i(t=0)}$	0.0085***	0.0085***	0.0085***	0.0085***	0.0085***	0.0085***
	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)
Mean of dependent variable	0.006	0.151	0.332	0.429	0.489	0.529
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes 2SLS estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the district. Standard errors (in parentheses) are robust to heteroskedasticity of unknown form. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.20 – 2SLS results with district FEs: the effect of a CD at parity 0 on the likelihood of maternal employment

	(I) $E_{i(t+1)}$	(II) $E_{i(t+2)}$	(III) $E_{i(t+3)}$	(IV) $E_{i(t+4)}$	(V) $E_{i(t+5)}$	(VI) $E_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	0.0116	0.0244	0.0983 **	0.0768*	0.0652	0.0527
	(0.0342)	(0.0432)	(0.0418)	(0.0413)	(0.0417)	(0.0348)
Eq. (2): $leisure_{i(t=0)}$	−0.0479 ***	−0.0479 ***	−0.0479 ***	−0.0479 ***	−0.0479 ***	−0.0479 ***
	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)	(0.0019)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.6278*	−0.2714	−0.1771	−0.0778	−0.2869	−0.3210
	(0.3253)	(0.3207)	(0.3296)	(0.3036)	(0.3099)	(0.3153)
Eq. (3): $preleisure_{i(t=0)}$	0.0085 ***	0.0085 ***	0.0085 ***	0.0085 ***	0.0085 ***	0.0085 ***
	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)	(0.0024)
Mean of dependent variable	0.214	0.448	0.511	0.549	0.602	0.652
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes 2SLS estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)). Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of maternal employment 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the district. Standard errors (in parentheses) are robust to heteroskedasticity of unknown form. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.21 – 2SRI results: the effect of a CD at parity 0 on the likelihood of a second birth

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0022	−0.0181	−0.0775***	−0.1090***	−0.1125***	−0.1192***
	(0.0026)	(0.0123)	(0.0161)	(0.0162)	(0.0161)	(0.0164)
Eq. (2): $leisure_{i(t=0)}$	−0.0497***	−0.0497***	−0.0497***	−0.0497***	−0.0497***	−0.0497***
	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0017	−0.0121	−0.0753***	−0.1110***	−0.1172***	−0.1230***
	(0.0028)	(0.0142)	(0.0172)	(0.0178)	(0.0186)	(0.0186)
Eq. (3): $preleisure_{i(t=0)}$	0.0105***	0.0105***	0.0105***	0.0105***	0.0105***	0.0105***
	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
Mean of dependent variable	0.006	0.151	0.332	0.429	0.489	0.529
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes 2SRI estimation results for the leisure IV and the pre-leisure IV. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of a second birth 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.22 – 2SRI results: the effect of a CD at parity 0 on the likelihood of maternal employment

	(I) $E_{i(t+1)}$	(II) $E_{i(t+2)}$	(III) $E_{i(t+3)}$	(IV) $E_{i(t+4)}$	(V) $E_{i(t+5)}$	(VI) $E_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0247**	0.0282*	0.0699***	0.0751***	0.0564***	0.0351**
	(0.0121)	(0.0149)	(0.0181)	(0.0189)	(0.0174)	(0.0139)
Eq. (2): $leisure_{i(t=0)}$	−0.0497***	−0.0497***	−0.0497***	−0.0497***	−0.0497***	−0.0497***
	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0332**	0.0258	0.0625***	0.0764***	0.0533***	0.0309*
	(0.0144)	(0.0174)	(0.0191)	(0.0195)	(0.0186)	(0.0165)
Eq. (3): $preleisure_{i(t=0)}$	0.0105***	0.0105***	0.0105***	0.0105***	0.0105***	0.0105***
	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
Mean of dependent variable	0.214	0.448	0.511	0.549	0.602	0.652
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes 2SRI estimation results for the leisure IV and the pre-leisure IV. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of maternal employment 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. Reported estimates are average marginal effects. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.23 – Bivariate probit results: fertility (controlling for birth weight and gestation length)

	(I)	(II)	(III)	(IV)	(V)	(VI)
	$F_{i(t+1)}$	$F_{i(t+2)}$	$F_{i(t+3)}$	$F_{i(t+4)}$	$F_{i(t+5)}$	$F_{i(t+6)}$
Panel A: Leisure IV						
Eq. (1): $CD_{i(t=0)}$	−0.0059	−0.0104	−0.0895***	−0.1167***	−0.1168***	−0.1269***
	(0.0072)	(0.0183)	(0.0225)	(0.0212)	(0.0221)	(0.0232)
Eq. (2): $leisure_{i(t=0)}$	−0.0482***	−0.0481***	−0.0482***	−0.0482***	−0.0482***	−0.0482***
	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
$\rho : Corr(\varepsilon^{p}_{it}, \epsilon_{it})$	0.1362	−0.0585	0.0616*	0.0904***	0.0903***	0.1079***
	(0.2206)	(0.0447)	(0.0366)	(0.0324)	(0.0340)	(0.0359)
Panel B: Preleisure IV						
Eq. (1): $CD_{i(t=0)}$	−0.0016	0.0062	−0.0893***	−0.1256***	−0.1233***	−0.1335***
	(0.0056)	(0.0222)	(0.0280)	(0.0268)	(0.0293)	(0.0303)
Eq. (3): $preleisure_{i(t=0)}$	0.0078***	0.0078***	0.0078***	0.0077***	0.0078***	0.0077***
	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
$\rho : Corr(\varepsilon^{p}_{it}, \epsilon_{it})$	−0.0047	−0.0987*	0.0614	0.1042**	0.1002**	0.1179**
	(0.1901)	(0.0535)	(0.0453)	(0.0409)	(0.0446)	(0.0465)
Mean of dependent variable	0.006	0.151	0.332	0.429	0.489	0.529
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes bivariate probit estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)) if the covariates birth weight and gestation length are included. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of of a second birth 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are hospital competition, mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, gestational length, birth weight, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.24 – Bivariate probit results: employment (controlling for birth weight and gestation length)

	(I) $E_{i(t+1)}$	(II) $E_{i(t+2)}$	(III) $E_{i(t+3)}$	(IV) $E_{i(t+4)}$	(V) $E_{i(t+5)}$	(VI) $E_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0629***	0.0205	0.1108***	0.1234***	0.0881***	0.0545**
	(0.0232)	(0.0324)	(0.0299)	(0.0308)	(0.0320)	(0.0266)
Eq. (2): $leisure_{i(t=0)}$	−0.0480***	−0.0482***	−0.0481***	−0.0480***	−0.0481***	−0.0482***
	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
$\rho : Corr(\varepsilon_{it}^{p}, \epsilon_{it})$	0.1482***	0.0012	−0.1151***	−0.1411***	−0.1038**	−0.0625
	(0.0473)	(0.0481)	(0.0446)	(0.0464)	(0.0486)	(0.0426)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.1629***	0.0138	0.1135***	0.1546***	0.1042**	0.0507
	(0.0588)	(0.0584)	(0.0425)	(0.0405)	(0.0458)	(0.0474)
Eq. (3): $preleisure_{i(t=0)}$	0.0083***	0.0078***	0.0077***	0.0078***	0.0077***	0.0078***
	(0.0022)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
$\rho : Corr(\varepsilon_{it}^{p}, \epsilon_{it})$	0.3581***	0.0113	−0.1190*	−0.1882***	−0.1284*	−0.0564
	(0.1226)	(0.0867)	(0.0633)	(0.0617)	(0.0703)	(0.0753)
Mean of dependent variable	0.214	0.448	0.511	0.549	0.602	0.652
Number of observations	282,831	282,831	282,831	282,831	282,831	282,831

Notes: This table summarizes bivariate probit estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)) if the covariates birth weight and gestation length are included. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of maternal employment 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are hospital competition, mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, gestational length, birth weight, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.25 – Bivariate probit results: fertility (controlling for hospital competition)

	(I) $F_{i(t+1)}$	(II) $F_{i(t+2)}$	(III) $F_{i(t+3)}$	(IV) $F_{i(t+4)}$	(V) $F_{i(t+5)}$	(VI) $F_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0059	−0.0038	−0.0728***	−0.1099***	−0.1129***	−0.1222***
	(0.0071)	(0.0192)	(0.0261)	(0.0249)	(0.0255)	(0.0262)
Eq. (2): $leisure_{i(t=0)}$	−0.0497***	−0.0496***	−0.0497***	−0.0497***	−0.0497***	−0.0497***
	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1415	−0.0747	0.0337	0.0792**	0.0837**	0.0998**
	(0.2145)	(0.0466)	(0.0424)	(0.0381)	(0.0390)	(0.0404)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0010	0.0193	−0.0630*	−0.1191***	−0.1209***	−0.1294***
	(0.0059)	(0.0247)	(0.0338)	(0.0333)	(0.0371)	(0.0368)
Eq. (3): $preleisure_{i(t=0)}$	0.0081***	0.0081***	0.0081***	0.0080***	0.0081***	0.0080***
	(0.0024)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	−0.0224	−0.1298**	0.0177	0.0934*	0.0959*	0.1109**
	(0.1964)	(0.0590)	(0.0545)	(0.0510)	(0.0565)	(0.0565)
Mean of dependent variable	0.006	0.151	0.332	0.429	0.489	0.529
Number of observations	282,635	282,635	282,635	282,635	282,635	282,635

Notes: This table summarizes bivariate probit estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)) with supply density categories included as control variables. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of of a second birth 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are hospital competition, mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

Table C.26 – Bivariate probit results: employment (controlling for hospital competition)

	(I)	(II)	(III)	(IV)	(V)	(VI)
	$E_{i(t+1)}$	$E_{i(t+2)}$	$E_{i(t+3)}$	$E_{i(t+4)}$	$E_{i(t+5)}$	$E_{i(t+6)}$
PANEL A: LEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.0656***	0.0381	0.1333***	0.1446***	0.1191***	0.0917***
	(0.0225)	(0.0346)	(0.0336)	(0.0344)	(0.0324)	(0.0275)
Eq. (2): $leisure_{i(t=0)}$	−0.0495***	−0.0497***	−0.0496***	−0.0495***	−0.0495***	−0.0496***
	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)	(0.0020)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.1532***	−0.0250	−0.1491***	−0.1733***	−0.1516***	−0.1223***
	(0.0461)	(0.0517)	(0.0507)	(0.0525)	(0.0499)	(0.0448)
PANEL B: PRELEISURE IV						
Eq. (1): $CD_{i(t=0)}$	−0.2193***	0.0531	0.1604***	0.2056***	0.1745***	0.1414**
	(0.0701)	(0.0748)	(0.0512)	(0.0452)	(0.0482)	(0.0607)
Eq. (3): $preleisure_{i(t=0)}$	0.0087***	0.0081***	0.0079***	0.0080***	0.0078***	0.0078***
	(0.0022)	(0.0023)	(0.0023)	(0.0023)	(0.0023)	(0.0023)
$\rho : Corr(\varepsilon^p_{it}, \epsilon_{it})$	0.4759***	−0.0474	−0.1898**	−0.2674***	−0.2382***	−0.2029**
	(0.1459)	(0.1116)	(0.0781)	(0.0712)	(0.0765)	(0.0998)
Mean of dependent variable	0.214	0.448	0.511	0.549	0.602	0.652
Number of observations	282,635	282,635	282,635	282,635	282,635	282,635

Notes: This table summarizes bivariate probit estimation results based on the leisure IV (joint estimation of Eq. (1) and (2)) and the pre-leisure IV (joint estimation of Eq. (1) and (3)) with supply density categories included as control variables. Columns (I)-(VI) provide separate estimations for the effect of a CD at parity 0 on the likelihood of maternal employment 1 to 6 years after the first birth (Eq. (1)). The second row of each panel includes the effects of the leisure instrument (Eq. (2)) and the pre-leisure instrument (Eq. (3)) on the likelihood of a CD at parity 0. ρ is the estimated correlation between the error terms in both equations. The leisure indicator is included as a control variable in the pre-leisure IV specification. Additional control variables are hospital competition, mother's age, marital status, educational attainment, occupation, religious denomination, the child's sex, month of birth, year of birth, and the province. Standard errors clustered by district in parentheses. *, **, *** indicate statistical significance at the 10-percent, 5-percent, and 1-percent level.

D Estimated Impact on Fertility Rates: Back of the envelope Calculations

Our results imply that women who have a non-medically indicated CD at parity one are 16.51 percentage points less likely to have a second child (Table C.8), and 6.22 percentage points less likely to have a third child (Table C.10). We denote the number of children F, dummies indicating birth of child number n F^n, dummies indicating CD at parity n CD^n.

$$E(F) = \sum_{n=1}^{\infty} E(F^n)$$

$$E(F) = \sum_{n=1}^{\infty} P(F^n = 1)$$

$$E(F|CD^1 = 1) - E(F|CD^1 = 0) = \sum_{n=1}^{\infty} \left[P(F^n = 1|CD^1 = 1) - P(F^n = 1|CD^1 = 0)\right]$$

We know that $P(F^1 = 1)$ is not affected by CD^1. We will focus on mothers with at least on child from here. Moreover, we conservatively assume that a CD at first birth does not affect fertility beyond the third child. Using our estimates from Tables C.8 and C.10, we get:

$$E(F|F^1 = 1, CD^1 = 1) - E(F|F^1 = 1, CD^1 = 0) = \sum_{n=1}^{\infty} \left[P(F^n = 1|CD^1 = 1) - P(F^n = 1|CD^1 = 0)\right]$$

$$E(F|F^1 = 1, CD^1 = 1) - E(F|F^1 = 1, CD^1 = 0) = \sum_{n=2}^{3} \left[P(F^n = 1|CD^1 = 1) - P(F^n = 1|CD^1 = 0)\right]$$

$$E(F|F^1 = 1, CD^1 = 1) - E(F|F^1 = 1, CD^1 = 0) = -0.1867 - 0.0622$$

$$E(F|F^1 = 1, CD^1 = 1) - E(F|F^1 = 1, CD^1 = 0) = -0.2489$$

The Austrian total fertility rate in 2014 was 1.46 while roughly 20 percent of women in Austria remain childless (OECD, 2016). Consequently, the total fertility rate for mothers with at least one child is $1.46/0.80 = 1.825$. Hence, our estimated fertility reductions for first-time mothers correspond to $0.2489/1.8250 = 13.6\%$.

Reducing the CD rate from 28 percent to a recommended level of 19 percent would increase the total fertility rate by $0.2489 * 0.8 * 0.09 = 0.0164$. The OECD total fertility rate in 2014 was 1.68 (OECD, 2016). This implies a relative increase of $0.0179/1.68 = 1.07\%$.