Quaderni del Premio «Giorgio Rota»

N. 6, 2018

The Economics of Health and Medical Care

Centro di Ricerca e Documentazione Luigi Einaudi

Con il sostegno di Fondazione CRT

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Iniziativa realizzata con il sostegno di

Fondazione CRT

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Indice

Il Premio «Giorgio Rota»

L'intento del Premio *«Giorgio Rota» Best Paper Award* è di riprendere l'attività di ricerca annualmente condotta dal Comitato / Fondazione Giorgio Rota prima della sua inclusione nel Centro Einaudi, sulla relazione tra il pensiero e l'agire economico e un aspetto (ogni anno diverso) del vivere in società, mantenendo vivo il ricordo e l'insegnamento dell'economista Giorgio Rota, uno dei primi animatori del Centro, prematuramente scomparso.

Dal 2012 il Cento Einaudi ha dunque raccolto questa eredità rinnovando la formula della ricerca: è stato perciò istituito questo premio annuale dedicato a giovani ricercatori, con una qualificazione accademica nei campi dell'economia, sociologia, geografia, scienza politica o altre scienze sociali. I paper possono essere presentati sia in italiano che in inglese, e non devono essere stati pubblicati prima della data della Conferenza Rota, l'evento pubblico nel quale i vincitori hanno modo di presentare il loro lavoro.

La prima edizione aveva per tema *Contemporary Economics and the Ethical Imperative* e la Conferenza Giorgio Rota 2013 si è tenuta presso il Centro Einaudi il 25 marzo 2013 con keynote speech di Alberto Petrucci, LUISS Guido Carli, Roma.

La seconda edizione, nel 2013, è stata su *Creative Entrepreneurship and New Media* con Conferenza Giorgio Rota presso il Centro Einaudi, 14 aprile 2014 e keynote speech di Mario Deaglio, Università di Torino.

La terza edizione ha analizzato il tema *The Economics of Illegal Activities and Corruption*, con Conferenza Giorgio Rota presso il Centro Einaudi, 15 giugno 2015. Keynote speech di Friedrich Schneider, Johannes Kepler University (Linz, Austria).

La quarta edizione verteva su *The Economics of Migration*. Il 20 giugno 2016 si è tenuta la Conferenza Giorgio Rota presso il Campus Luigi Einaudi, in collaborazione con FIERI. Keynote speech di Alessandra Venturini, Università di Torino. Dal 2016 inoltre il Premio è sostenuto dalla Fondazione CRT.

La quinta edizione, del 2017, trattava di *Economic Consequences of Inequality*, e i saggi vincitori sono stati presentati alla Conferenza Giorgio Rota del 4 maggio 2017, tenutasi presso il Campus Einaudi in collaborazione con il Dipartimento di Economia e Statistica "Cognetti de Martiis". L'Introduzione è di Andrea Brandolini, Banca d'Italia.

La sesta edizione del Premio, tenutasi nel 2018, è incentrata sul tema *The Economics of Health and Medical Care*. I paper vincitori, presentati alla Conferenza Giorgio Rota il 1° giugno 2018, sono di Gabriel Facchini, Gianni Ghetti e Valentina Tonei e vengono riportati in questo volume. L'Introduzione è di Fabio Pammolli, Politecnico di Milano.

Chi era Giorgio Rota

Giorgio Rota (1943-1984) è stato professore di Economia politica presso l'Università di Torino e consulente economico. Per il Centro Einaudi, è stato coordinatore agli studi e membro del comitato di direzione di «Biblioteca della libertà».

Le sue pubblicazioni scientifiche abbracciano diversi temi: l'economia dei beni di consumo durevoli, l'economia del risparmio, il mercato monetario e finanziario, l'inflazione e la variazione dei prezzi relativi, il debito pubblico. Ricordiamo tra esse: *Struttura ed evoluzione dei flussi finanziari in Italia: 1964-73* (Torino, Editoriale Valentino, 1975); *L'inflazione in Italia 1952/1974* (Torino, Editoriale Valentino, 1975); nei «Quaderni di Biblioteca della libertà», *Passato e futuro dell'inflazione in Italia* (1976) e *Inflazione per chi?* (1978); *Che cosa si produce come e per chi. Manuale italiano di microeconomia*, con Onorato Castellino, Elsa Fornero, Mario Monti, Sergio Ricossa (Torino, Giappichelli, 1978; seconda edizione 1983); *Investimenti produttivi e risparmio delle famiglie* (Milano, Il Sole 24 Ore, 1983); *Obiettivi keynesiani e spesa pubblica non keynesiana* (Torino, 1983).

Tra le sue ricerche va particolarmente citato il primo *Rapporto sul risparmio e sui risparmiatori in Italia* (1982), risultato di un'indagine sul campo condotta da BNL-Doxa-Centro Einaudi, le cui conclusioni riscossero notevole attenzione da parte degli organi di stampa. Da allora il *Rapporto sul risparmio,* ora *Indagine sul risparmio*, continua a essere pubblicato ogni anno.

Fabio Pammolli

Introduzione[1]

Grazie al Centro Einaudi e al Dipartimento di Economia e Statistica 'Cognetti De Martiis' per avermi invitato; è un onore e un piacere anche e soprattutto per i riferimenti culturali a cui si richiamava Salvatore Carrubba nella sua introduzione e che rendono ancora più gradevole essere qui per me oggi.

Ho preparato alcuni brevi appunti e riflessioni da condividere con voi, partendo dal macro per arrivare al micro, che è stato trattato in modo assai interessante nei lavori dei tre vincitori del Bando Rota 2018 che saranno oggi presentati qui al Campus. La parte dell'analisi sanitaria in campo economico e dell'economia della Sanità è particolarmente rilevante in generale e anche di più in questa fase storica.

Farei alcune brevi considerazioni.

In tutti i paesi a economia e welfare sviluppati, la spesa sanitaria è, tra le voci di spesa sociale, quella che nelle prossime decadi farà registrare la crescita più intensa in termini di Pil, e soprattutto la più soggetta ad alea per la presenza di fattori – lato offerta e lato domanda – il cui impatto è difficilmente quantificabile. La quota percentuale di tale spesa, rispetto al prodotto interno, è prevista comunque crescere tra il 40% e il 60% nei prossimi quarant'anni. Inoltre, nei paesi industrializzati la spesa sanitaria totale tende a crescere a tassi superiori rispetto a quelli dell'economia nel suo complesso. I principali fattori determinanti la crescita della spesa sanitaria sono legati ai trend demografici ed epidemiologici in atto (che accentueranno la domanda di assistenza sociosanitaria, e soprattutto di *long-term care*) e soprattutto ai processi di innovazione tecnologica in campo biomedico.

[1] Questa prefazione riporta il discorso di presentazione in apertura alla Giorgio Rota Conference tenutasi il 1° giugno 2018 al Campus Einaudi di Torino.

Il caso italiano non è particolarmente rilevante ma è un caso interessante: mi riferisco al rapporto tra stato sanitario ed economia sanitaria, in quanto snodo tra variabili demografiche, variabili tecnologiche e variabili istituzionali, dato che l'Italia ha una peculiarità in relazione al modello di finanziamento della spesa sanitaria riferita all'invecchiamento della popolazione.

L'analisi econometrica sul caso italiano presenta un modello teorico empirico sull'impatto sanitario, la cui ipotesi di fondo è che l'evoluzione delle soluzioni terapeutiche, soprattutto in relazione al trattamento di patologie croniche riferite all'età avanzata e le intersezioni tra assistenza sanitaria e assistenza socio-assistenziale nel trattamento sanitario, determinano l'espandersi di un'area grigia tra spesa previdenziale e spesa sanitaria, che sarà sempre più grigia negli anni a venire.

Le conseguenze saranno importanti, perché la spesa sanitaria ha picchi sia in relazione all'età sia in relazione ai profili di spesa e in Italia abbiamo una situazione surreale di carico fiscale e di modalità degli interventi. Se noi consideriamo il Pil procapite e guardiamo a quale quota grava su ciascun lavoratore italiano per finanziare la spesa per la sanità, scopriamo che a ciascun occupato risponde un valore del 70% del Pil procapite, mentre il valore equivalente tedesco si aggira sul 40% (over 75). È impensabile pensare a forti guadagni di produttività del lavoro, ma in compenso l'invecchiamento progressivo della popolazione e i flussi migratori portano e porteranno conseguenze, e questo pur con l'ipotesi sottostante di una tenuta macroeconomica in relazione all'entità del debito pubblico.

Una ipotesi potrebbe essere quella di differenziare l'intervento iniziando già da coloro che non sono ancora sul mercato del lavoro, per riuscire ad accantonare i fondi appropriati che sono poi alla base del sistema del "pilastro". Di fronte alle proiezioni di spesa prima accennate, la stabilizzazione della spesa pubblica sul Pil ai livelli correnti implica riduzioni significative della copertura pubblica, con conseguente implicito affidamento della domanda al finanziamento privato: per l'Italia, il coverage del Ssn è proiettato in riduzione dall'attuale 75% a meno del 50% nel 2050. In questo scenario si rende indispensabile e necessario definire una governance in grado di combinare, sulla base di scelte positive, l'obiettivo della stabilità finanziaria con quello dell'adeguatezza/equità delle prestazioni.

Cedo il passo alle presentazioni dei giovani studiosi e mi auguro che i ragazzi possano trovare nuove opportunità, anche in relazione al fatto che stiamo sviluppando una collaborazione tra Milano e Torino. Grazie per l'attenzione.

Gabriel Facchini

Low staffing in the maternity ward: Keep calm and call the surgeon[1]

Abstract. This paper investigates the relationship between workload and choice of treatment in a large but understudied segment of the healthcare sector – maternity wards. Using detailed microdata on childbirth, I exploit quasi-random assignment of patients attempting to have a natural delivery to different ratios of patients-to-midwives and compare their likelihood of changing delivery method. I find that women who face a ratio higher than 1.33 are 34% more likely to give birth by cesarean section (C-sections). This effect is larger for patients who were already admitted with a higher risk of C-section, implying that provision of proper and timely care matters more for this type of patients. Because C-sections are faster than vaginal deliveries – in which the patient follows the course of labor –, the medical team may find it appealing to do more C-sections when time constrained. Using civil status as a proxy for bargaining power -assuming single women are on average more likely to be alone, I find that only single patients are subjected to unnecessary surgery. This provides evidence that high midwives' workload is yet another factor which triggers physician-induced-demand for C-sections.

Keywords. Cesarean section, capacity utilization, workload, midwives, physician-induced-demand

[1] Department of Economics, European University Institute, Florence, Italy. Email: gabriel.facchini@eui.eu. I would like to extend my gratitude to Andrea Ichino and Jérôme Adda for their invaluable support and advice throughout this research project. I also benefited greatly from discussions with Juan Dolado, Daniela Iorio, Matilde Machado and Matthew Neidell. I am extremely grateful to Carlo Dani, Simone Pratesi, Federico Mecacci, Lucia Pasquini, Franca Rusconi and Luigi Gagliardi for their clinical expertise, and to Tommaso Lanis and Francesca Superbi for their assistance in accessing the data. I also thank seminar participants at the EUI Micro-Econometrics Working Group. All errors remain my own.

1. INTRODUCTION

Over the last decades health care systems in developed countries have been under constant pressure to reduce costs, despite facing an increasing demand for health care services. In order to avoid a trade-off between cutting down on costs and a negative impact on patients' health outcomes, experts currently point towards the reduction of waste as the best way to go.[2] Among the several sources of waste, two widely cited ones are the lack of adoption of known best practices (e.g. effective preventive care) and overtreatment, that is, the carrying out of treatments that cannot possibly improve the patient's health (e.g. cases of physician induced demand). These two sources of waste are particularly salient in maternity ward settings.

The role of midwives -as opposed to physicians- in assisting birth speaks to the first point. Whereas relevant public health authorities have recently recognized that midwife-led care during labor is safer for low-risk pregnancies,[3] the media and midwifery colleges have long spoken of a "shortage of mid- wives",[4] which was also acknowledged by the World Health Organization (WHO) in 2009.[5] At the same time, cesarean sections (C-sections) rank high among greatly overused interventions[6], and governments and clinicians have expressed concern about its potential negative impact on patients' health.[7] Indeed, C-sections not only cost more than vaginal deliveries, but they also imply higher risks for both mother and infant[8] and, according to a growing medical literature, are associated to lower long-term outcomes of children's health.[9] In addition, because vaginal delivery after a C-section (VBAC) is very unlikely,[10] one C-section sets a path dependency for more C-sections in future births. There is also evidence that women who follow a C-section are

[2] See, for example, Berwick and Hackbarth (2012).
[3] For example, the National Institute for Health and Care Excellence (NICE) updated its guidelines in this direction in 2014.
[4] In a 2015 report, The Royal College of Midwives estimates that the UK "needs 2,600 more midwives to be able to cope with the number of births the country is experiencing". The Federal Association of Midwives of Spain (FAME) has as main objective to address the shortage of midwives in the health care system. The president of the Italian Midwifery Association recently stated that "there is a shortage of midwives. Too few to guaranty the proper level of care that other European Countries have".
[5] Büscher *et al.* 2009.
[6] While the international healthcare community considers an ideal rate of C-sections to be between 10-15%, country average rates in Europe vary from as low as 15.6% in The Netherlands to as high as 36.8% in Italy (OECD data 2012).
[7] WHO Statement on Caesarean Section Rates, WHO (2015).
[8] See Deneux-Tharaux *et al.* 2006; Gregory *et al.* 2012; Curtin *et al.* 2015.
[9] Infants born by C-section are not exposed to the maternal bacteria of the birth canal and as a consequence have different intestinal bacteria, which can affect their immune system and other important processes. For a meta-analysis of this literature see (Blustein and Liu 2015).
[10] VBAC rate is only 8.3% in the US, and 12% in Italy.

more likely to have less children,[11] something that is particularly alarming in developed countries with already low fertility rates.

In light of these concerns, a natural question is whether a situation of low staffing can result in more unnecessary C-sections being performed. This can happen either as a direct consequence of high workload -with midwives devoting less time to each patient, therefore rising the probability of complications that lead to surgery- or because physicians may find it optimal to induce some patients towards a C-section independently of their health status. Since a C-section takes less time than a vaginal birth – no need to wait for the appropriate dilation of the cervix –, midwives' workload can be reduced by shifting patients to surgeries.

This study causally tests whether patients follow a different delivery method depending on the effective staff level in the maternity ward at the moment of admission. It exploits a simple natural experiment: the majority of patients follow the natural course of birth and only go to the hospital once labor has already started and/or their water has broken (unlike, for example, scheduled cesarean sections). The effective staff level (e.g. the staff per patient ratio) observed by these patients at admission is orthogonal to their demographic and health characteristics (and to their ex-ante probability of delivering by C-section). The effective staff level at admission changes with the number of patients who arrived before and the number of midwives present in the delivery room, two variables that are unknown for the incoming patient.

The data for this project comes from a census of births from a large public hospital in Italy for the period 2011-2014. Three features of this dataset make it well suited for tackling the issue at hand. First, birth certificates have precise information on delivery method, allowing the identification of scheduled and unscheduled patients. Second, using patient's ID, each certificate was merged with hospital administrative data containing the exact time of admission and discharge. I use this information to compute the actual number of patients in the delivery room at each point in time. Finally, this is complemented with data on the number of midwives scheduled by month, day of the week and shift.

Results suggests that there is a non-linear relationship between effective midwifery staff and de- livery method: a newly admitted patient who faces a ratio of patients-to-midwives higher than 1.33 is 34% more likely to give birth by C-section. This means that, for first-time mothers, about 1.2 p.p. (or 5.7%) of all C-sections (both scheduled and unscheduled) are the consequence of low midwifery staffing.

[11] Norberg and Pantano 2016.

The second part of the analysis looks at possible mechanisms behind this change in delivery method. One possibility is that, in situations with a high ratio of patients-to-midwives, the time dedicated to each patient is lower and the quality of care inappropriate, eventually resulting in the need for C-section. If that is true, then one should see patients with marginally lower health being more affected. In order to test this hypothesis, two types of patients are compared: a low-health type, formed by those patients who had an emergency visit during their pregnancy or whose babies had an extreme weight at birth, and a high-health type, with all the remaining patients. Indeed, the gap between the probability of having a C-section between a low-health and a high-health patient widens with a higher workload.

Another factor that can explain the rise in C-sections alongside with workload is the presence of physician induced demand (PID). Because C-sections are faster than vaginal births, when faced with time constraints, physicians may decide to put some patients through surgery -without a medical necessity for it-, reducing the midwives' workload. Within the agency discrimination framework, physicians will choose to practice an unnecessary surgery on patients with lower bargaining power. This study tests for the presence of agency by comparing single women and non-single women, assuming that single patients are – on average – more likely to be alone in the delivery room. In those cases, the physician will need less effort in convincing the patient to have a C-section. Indeed, the data shows that the gap in the probability of delivering by C-section between these two groups is statistically significant only for high ratios of patients-to-midwives. On the other hand, I find that married and low-risk patients are between 24% and 35% more likely of not attaining skin-to-skin contact with their newborn when the number of patients per midwife is high. This provides more evidence that, by performing more C-sections, physicians are avoiding some bad outcomes.

This paper contributes to several strands of literature. First, it adds to existing work on the effect of staff ratios on health outcomes. Previous studies find none or very small effects when using census discharge data (Evans and Kim 2006; Cook *et al.* 2012), and a negative impact of crowding on health when focusing on patients in the Emergency Department (ED) (de Araujo *et al.*, 2013). This difference between areas makes sense given the particular time constrains of patients in the ED. The maternity wards lay somewhere in between these two. However, there is no study looking at the effect of staff ratio in maternity wards using a casual approach. The one that comes closest to this is Balakrishnan and Soderstrom (2000), using data from 225,473 maternity admissions at 30 hospitals in the state of Washington. They identify crowded days using a percentile cut-off from the distribution of patients' admissions for each hospital-year combination and the rate of C-sections as outcome. They find a positive and significant correlation between the two, only for those pregnancies that are classified as at-risk of C-section. A shortcoming of this paper is that they cannot differentiate between scheduled and unscheduled patients in their data, rising concerns about causal relationships. It could be the case that days with more patients are those with more planned C-sections, without necessarily having any effect on patients' health outcomes. I contribute to this literature by causally estimating the effect of low staffing ratios on delivery method.

Second, there is a vast number of empirical studies that look at different causes for the exceedingly high levels of C-sections. Starting from the paper by Gruber and Owings (1996) where they use physician's income drop as a trigger for more C-sections, to other incentives like relative prices between C-sections and vaginal deliveries (Gruber *et al.* 1999; Alexander *et al.* 2013; Allin *et al.* 2015), defensive medicine (Keeler and Brodie 1993; Lawthers *et al.* 1992; Currie and MacLeod 2008; Dranove and Watanabe 2009), and physician's scheduling convenience (Lefèvre 2014).[12] I provide of scheduled and unscheduled patients. Second, using patient's ID, each certificate was merged with hospital administrative data containing the exact time of admission and discharge. I use this information to compute the actual number of patients in the delivery room at each point in time. Finally, this is complemented with data on the number of midwives scheduled by month, day of the week and shift.

Third, this study also relates to the literature that empirically tests possible mechanisms behind PID. Two recent papers use information asymmetry variations in the maternity ward set up. Grytten *et al.* (2011) compare expert and non-expert patients and conclude that a model of statistical discrimination (expert patients are better at communicating with the physician) explain their results better than one of agency discrimination (physician influences the diagnosis and treatment for non-expert patients). On the contrary, Johnson and Rehavi (2016) find evidence that physicians are more likely to exploit the information asymmetry when it is profitable. They do so by comparing physician patients with non-physician patients, in settings with and without financial incentives to perform C-sections. I add to this body of work by using a different approach to test for bargaining power: whether the mother is alone in the delivery room.

The remainder of this paper is organized as follows: Section 2 describes the clinical and institutional setting. Section 3 discusses the identification strategy followed and describes the data. Section 4 reports the results, and Section 5 concludes.

[12] For an extensive review of this literature see Allin *et al.* 2015.

2. Clinical and Institutional Setting

Maternity wards receive two types of patients: scheduled and unscheduled. The former includes patients admitted for an elective C-section and those who will be induced.[13] For patients following an elective C-sections the date of delivery is set in advance, and there is no possibility for changing delivery method (unless the mother goes into labor before). These pregnancies typically present some health condition that constitute a risk for the mother and/or the baby if delivered vaginally. Similarly, induced patients already know in advance the date they will be induced but, although they will attempt a vaginal delivery, the physician may still decide to change delivery method on the way if seen necessary.

The remaining patients, those attempting to follow the natural course of labor and vaginal delivery, are the focus of this study. For these patients the process starts with frequent contractions and/or because they believe their water has broken (spontaneous onset of labor). Once the mother arrives to the hospital she is evaluated and if in active labor, she is admitted into a labor and delivery room and assigned a gynecologist and a midwife. If everything goes as plan and the patient can have a vaginal delivery, the midwife will be the one helping her throughout the whole process. Nevertheless, during labor there are several medical conditions that can emerge and complicate a vaginal birth, putting in danger the health of the infant and/or the mother. Under these circumstances, the midwife and gynecologist may decide to recommend having a C-section instead.

More importantly, the actual presence of some of these medical conditions depend heavily on the subjective opinion of the gynecologist.[14] The presence of this gray area -or asymmetry of information on when is a C-section necessary gives the gynecologist more room to suggest the patient to follow a surgery, even when not medically needed.

The maternity unit analyzed in this paper is part of one large teaching hospital in Italy. The staff working in the delivery room are paid a fixed salary, meaning they have no personal financial incentive to recommend any treatment. On the other hand, hospitals are reimbursed depending on a DRG (Diagnosed-related group) tariff system, which in general gives a higher reward for a C-section than a vaginal delivery.[15]

[13] Most inducements are performed on pregnancies that have past their due date and still haven't started labor.
[14] Two of these more 'subjective' conditions are dystocia (abnormally slow labor) and fetal distress.
[15] For a deeper discussion on the Italian Health System see Francese *et al.* 2014.

3. Empirical Methodology

3.1 A natural experiment

An ideal experiment to test for an effect of low-staffing in the maternity ward on patients' delivery method would imply assigning parturient women randomly between two different hospital types: a first one with already a large number of patients and a second type, identical to the first, but with few patients and hence ready to focus entirely on the coming patient. For obvious reasons this is not possible to implement in practice.

This paper focuses on patients who attempt vaginal delivery, and uses the exogenous variability in the number of patients and midwives present at admission to causally identify the impact of low staffing on delivery method. For the majority of births, the time of arrival is unknown to the hospital beforehand. In the same way, the level of capacity utilization of the maternity ward in a given point in time is unknown for future patients until they reach the hospital. For this sample of patients, their pre-admission probability of developing a complication and needing C-section is orthogonal to the level of crowding at the hospital.

The study sample includes all births that, up to the point of arriving to the hospital, followed the 'natural' course of pregnancy and labor. This means leaving out all scheduled deliveries where the physician decided, together with the patient, the date when the birth should take place. This type of patients are those who had an elective C-section or who were pharmaceutically induced to start labor.[16]

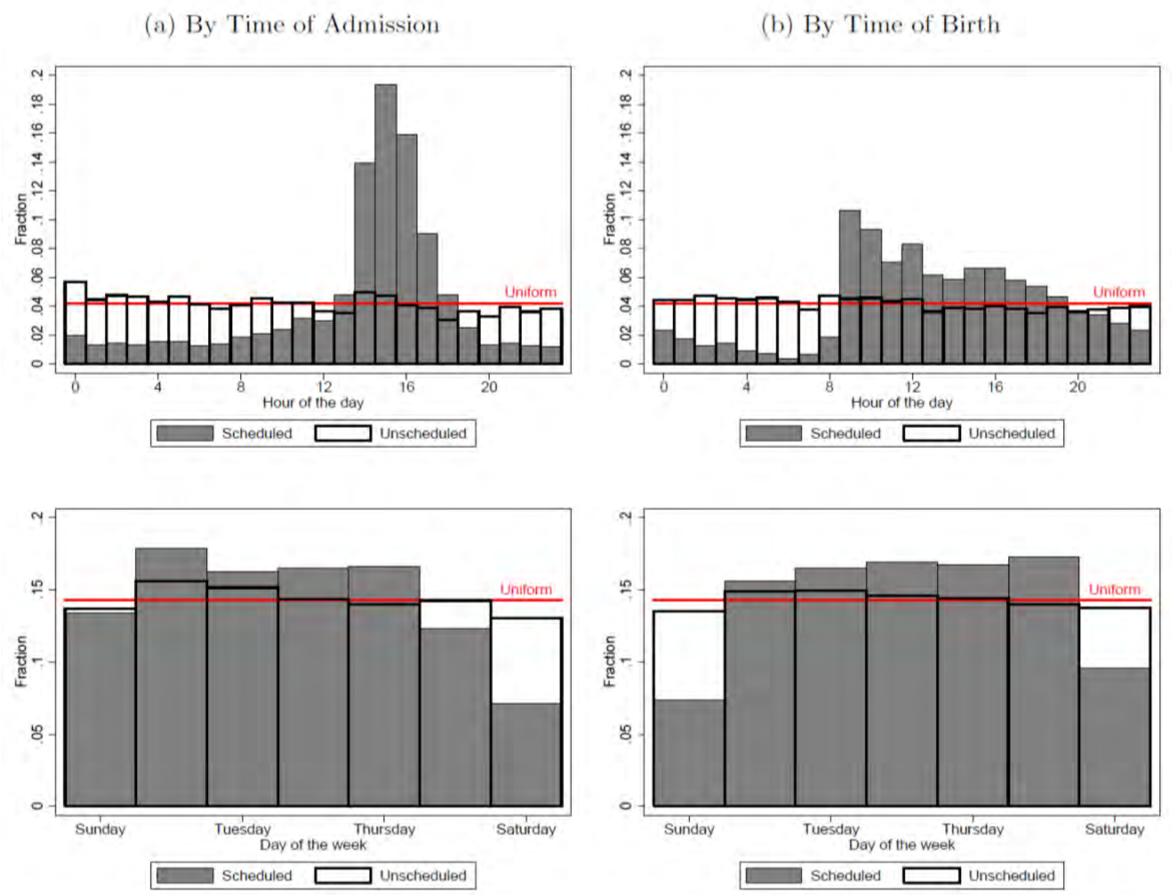
The left column of Figure 3.1 shows the distribution of admissions by hour of the day and day of the week. The right column does the same for births. Both are estimated for scheduled and unscheduled patients for comparison. We can immediately see that admissions of scheduled patients are concentrated in the afternoon, while births start at 9 a.m. and become less and less frequent as the day goes by. Instead, both admissions and births for unscheduled patients are very close to a uniform distribution across the day. When looking at the distribution by days of the week, again unscheduled patients are randomly distributed while scheduled patients are less common to be admitted on Saturdays, and less likely to have surgery on Sundays and Saturdays.

[16] For more evidence supporting the criteria for selecting the working sample see Appendix A.

3.2 Data

Previous studies looking at newborns' health tend to use anonymous birth certificates since they are publicly available for many countries and for long periods of time. However, these datasets commonly lack information on key variables needed for a rigorous study of staffing levels, namely, the exact date and time of admission of patients (demand side) and the number of staff available (supply side), for each hospital.

Figure 3.1. Distribution of admission and birthes



This study utilizes data from the Maternity Department of the Azienda Ospedaliero Universitaria Careggi (AOUC) for the years 2011 through 2014. This is the biggest hospital in the Province of Florence with more than 3,000 deliveries per year. The primary databases used are two: (i) birth certificates;[17] and (ii) hospital admissions.[18] Birth

[17] Certificato di assistenza al parto (CEDAP).
[18] Scheda di Dimissione Ospedaliera (SDO).

certificates constitute a census of all births that took place in the hospital in this period. It contains information on mother characteristics (e.g. com- munity of residence, education, civil status, age, previous deliveries, etc.), pregnancy characteristics (e.g. weeks of gestation, controls, assisted reproduction, etc.) and birth characteristics (e.g. time of birth, type of labor, attendant, place, weight of the baby etc.). The administrative hospital admission data provides information on the time of admission and time of discharge for each patient. Using unique mother-pregnancy identifiers, both databases can be merged together.

This data on patients is complemented with information on the level of staff scheduled to be present at each month, day of the week and shift of the day in the delivery room. Note that this is not the effective level of staff present at each point in time but the schedule that the personnel should follow. Anecdotal evidence suggests that deviations from planned levels are rare, even because the hospital calls in someone else when an employee misses her shift.

However, the richness of this dataset comes at a cost: because the information available corresponds only to one hospital in a four-year period the sample size is relatively small. Furthermore, due to the path dependence of treatment in second and higher order births, this study focuses on first-time mothers. There were approximately 5,240 singleton births at this hospital in the sample period. From this, about 870 observations are plural births and/or delivered by urgent C-section which will not be considered in the analysis because of their particular characteristics and handling within the hospital. Then further restricting the sample to non-induced planned-vaginal deliveries the number of observations goes down to around 2,685. Finally, after dropping observations with missing time of admission, maternal age, education, birth order, weight and prenatal visits, the number of observations in the working sample is about 2,600. The models described below are fitted to this sample.

Table 3.1 summarizes the variables used in the analysis. The first column corresponds to the whole sample. Most of the patients who attempt a vaginal delivery succeeded. Only about 12% had an in-labor C-section. Patients are on average 31 years old, only 36% has a university degree, and 44% are single. There are few cases with bad outcomes: only 4.6% have a 5-minute APGAR score below 9, and about 5% are born prematurely or weighting less than 2,500 grams. Columns 2 and 3 report statistics for patients with a low and high ex-ante risk of C-section respectively. Columns 4 and 5 do the same by civil status. By construction, patients from the high-risk are more likely to give birth by C-section, use of a neonatal intensive care unit, and have an APGAR score below 9. They are also more likely to be single and less likely to have a university degree. Finally, single patients are less likely to have a university degree and more likely to delivery by C-section, although other outcomes are similar to the married subsample.

Table 3.1. Descriptive statistics

	All	Low-risk	High-risk	Married	Single
Outcomes					
% vaginal birth	88.1	88.8	85.1	89.5	86.6
% in-labor C-section	11.9	11.2	14.9	10.5	13.4
Other interventions/health outcomes					
% operative birth	13.3	13.6	11.9	13.5	12.6
Average length-of-stay (hours)	76.0	75.3	79.0	76.7	75.9
% need of NICU	7.3	4.8	17.8	5.9	8.1
% lack of skin-to-skin contact	19.3	16.0	33.5	18.0	20.0
% non-exclusive breastfeeding	36.0	33.9	46.7	35.5	36.5
% APGAR score below 9	4.6	3.3	10.5	3.8	5.4
Mother's characteristics					
Average age	31.1	31.2	30.6	31.2	30.8
% with university degree	35.9	36.3	33.9	41.3	30.2
% single	44.2	43.5	47.0	0.0	100.0
Pregnancy's characteristics					
% born before 37 weeks of gestation	5.3	2.7	16.6	5.2	5.2
% with at least 1 ER visit	11.5	0.0	60.6	10.4	13.3
Newborn's characteristics					
% male	51.0	50.2	54.3	51.9	50.4
Average weight at birth	3,235	3,271	3,085	3,234	3,234
% low birthweight (<2,500 grams)	4.9	0.0	26.1	4.9	4.9
% high birthweight (>4,000 grams)	3.9	0.0	20.4	4.4	3.7
Observations	2,613	2,118	495	1,300	1,028
	All	**Low-risk**	**High-risk**	**Married**	**Single**

Note: Statistics for main sample of unscheduled first-time mothers, from 2011-2014. High-risk are patients who, at admission, have a higher probability of needing a C-section. Those are defined as patients with newborns with extreme birthweight and patients with an emergency department visit during pregnancy. Low-risk are those without any of those characteristics.

3.3 An exogenous measure of midwives' workload

A good measure of effective staff contains information on both number of patients and personnel. For the maternity wards setting of this paper I use the ratio between the number of patients and the number of midwives in the delivery room.[19] The richness of

[19] One drawback of this measure is that it constraints the coefficient of interest due to the simultaneous variations in numerator and denominator. The fact that my preferred model specification uses fixed effects by shift and day-of-the-week means that all the variation used for the estimation comes solely from fluctuations in the numerator, alleviating this issue. Furthermore, Appendix B repeats the main analysis using solely the number of patients as the covariate of interest, and results are qualitatively the same. In light of these results,

the data in hand allows to construct a very precise measure of the number of parturient women in the maternity ward at any point in time and to differentiate between those waiting to give birth and those in postpartum. But there are yet two decisions to be taken regarding the moment at which this ratio is calculated, and the type of patients to include in the numerator. On the former, because patients stay on average 7 hours in the delivery room between admission and birth, it is not obvious at what time to measure the level of staffing. The two most obvious options are at the time of admission and at the time of delivery. The last one has the advantage of measuring staff when needed the most, meaning, when the mother needs help to give birth. The problem with this option is that, given that physicians can rush a delivery (e.g. by doing a C-section), the level of staffing at time of birth can be endogenously determined. On the other hand, even though the level of staffing at time of admission can be relatively less relevant, it is indeed an exogenous shock. For these reasons I will use the ratio of patients to midwives calculated at the time of admission of each patient.[20]

On the second issue, it is important to clarify which patients are included in this measure of staffing. The first option would be to include all patients (regardless of whether they are scheduled or induced). One could think that, because the time of the admitted patient is random, there is no risk of endogeneity here. Nevertheless, since the outcome of interest is the probability of C-section, counting elective C-sections in the measure of staffing would make it biased. To see this, note that when there are more elective C-sections there are also more gynecologists ready to perform them. Incorporating elective C-sections in the numerator would not only include a demand side but also a change in the supply of physicians who can perform C-sections. Hence this study includes in the numerator all patients but those already scheduled to give birth by C-section.[21] Instead, the number of scheduled C-sections is included in the regression as control (see econometric specification below). More specifically, in this paper the workload observed by patient i at admission time t is define as

$$R_{it} = \frac{PVB_{it}}{MW_{it}} \tag{1}$$

in the main paper I will use the ratio of patients-to-midwives since it provides an advantage with regard to external validity (findings become less dependent on the size of the hospital studied).

[20] In the following section I perform several robustness check measuring staff levels at different points in time during a patient's stay, and discuss the results.

[21] Note that this is not the same sample as the study sample because it also includes induced deliveries. Those are not at risk of contaminating the measure because they will still attempt a vaginal delivery, and will need a midwife to help them.

where PVB is the number of patients waiting to attempt a vaginal birth, and MW is the number of midwives scheduled to be present in the delivery room.

Table 3.2 shows the mean number of midwives and patients (with its standard deviation) in the delivery room by day of the week and shift of admission. The number of midwives is higher during the morning shift (5), and lower at nights and Sundays (3). On the other hand, the average number of patients is virtually the same across days of the week and shifts, with a slightly lower level on Sundays.[22]

Table 3.2. Number of midwives and patients by day of the week and shift

Day	Shift	Midwives (mean)	Midwives (sd)	Patients§ (mean)	Patients§ (sd)
Weekdays	Morning (7am - 1pm)	5	0	7.31	2.81
	Afternoon (1pm - 7pm	4	0	7.48	2.89
	Night (7pm - 7am)	3	0	7.32	2.86
Saturdays	Morning (7am - 1pm)	4	0	7.53	2.63
	Afternoon (1pm - 7pm	4	0	7.41	2.70
	Night (7pm - 7am)	3	0	7.26	2.71
Sundays	Morning (7am - 1pm)	3	0	7.09	2.73
	Afternoon (1pm - 7pm)	3	0	7.08	2.76
	Night (7pm - 7am)	3	0	6.94	2.68

§ Number of patients waiting who attempt to have a vaginal birth.

Table 3.3 shows the distribution of the ratio of patients to midwives for the whole sample and then disaggregated by shift of admission. The ratio is unimodal and slightly skewed to the right.[23] At the median, there are 2 patients for every midwife in the delivery room. The 25th and 75th percentiles are 30% (below) and 34% (above) the median, respectively. Note that shifts later in the day have higher values of the ratio, meaning, more crowding. Remember that the distribution of patients is rather uniform across the day, hence this upward shift in the ratio comes exclusively from a lower supply (less midwives present).[24] The bottom rows of the Table 3.3 show the cutoff values for the lowest and highest quintiles (and by construction for the three middle quintiles altogether). The lowest

[22] The difference with Sunday is due to the fact that there are less induced births.
[23] See Figure D.1 for a graphic representation of the density distribution of the ratio by shift.
[24] In Figure D.2 one can see how the average ratio of patients to midwives by hour of admission shows a discrete jump up with each change in shift due to one less midwife being present.

quintile will be considered a case with no crowding, with a mean of 1 patient per midwife. The middle quintiles have a mean ratio of 1.9, somehow crowded. The highest ratio, with a mean of 3.2 patients per midwife, will be referred to as highly crowded or chaos.

Table 3.3. Descriptive statistics for ratio of patients to midwives by shift of admission

	All	Morning (7am - 1pm)	Afternoon (1pm - 7pm)	Night (7pm - 7am)
p1	0.60	0.40	0.50	0.67
p5	0.80	0.75	0.75	1.00
p25	1.40	1.20	1.50	1.67
p50	2.00	1.50	1.75	2.33
p75	2.67	2.00	2.33	3.00
p95	3.67	3.00	3.25	4.00
p99	4.50	4.00	4.50	4.67
mean	2.06	1.60	1.91	2.36
sd	0.86	0.70	0.74	0.88
<20th Percentile	1.33			
>80th Percentile	2.67			
Obs.	2,613	636	641	1,336

3.4 Econometric specification

The first part of the analysis estimates OLS regressions of a binary indicator for C-section on the treatment variable along with demographic and clinical controls. A simple reduced-form linear probability model of the following type is used:[25]

$$y_{it} = \alpha + \beta R_{it} + \theta X_{it} + \gamma_t cs_{it} + dow \times shift_t + year_t + month_t + \epsilon_{it} \quad (2)$$

where y_{it} is a dummy variable indicating whether birth *i* admitted at time *t* had an in-labor C-section, and R_{it} is the ratio of patients-to-midwives observed at admission as explained above. X_{it} contains individual-level control variables of mother and pregnancy characteristics.[26] To further control for supply side changes in physicians' availability I

[25] A probit model was also estimated assuming a normal distribution of the error term and results virtually the same (See Table D.3).

[26] These include: a dummy for whether the mother is above 34 years old, a dummy for whether the mother has a university degree, a dummy for whether this is her first pregnancy, a dummy for whether the infant is a male, a dummy for whether is a pre-term birth (below 37 weeks of gestation), a dummy for whether the baby

include the number of scheduled C-sections that took place while the indexed patient was in the delivery room (*cs*). Since most supply side changes in the maternity ward take place between shifts and days, in the most demanding specification I also add fixed effects for day-of-the-week (*dow*) times *shift*.[27] To control for seasonal and secular variation in outcomes, I also include monthly and yearly dummy variables. β is the coefficient of interest. As discussed above, if physicians are more likely to perform a C-section when the ratio of patients to midwives is high, then β should be positive.

Two models are estimated for the probability of delivering by C-section. First, I use the ratio of patients to midwives added linearly to the model. Because there can be nonlinear effects between staffing and delivery method, for the second model I split the sample in three categories based on the ratio of patients-to-midwives: low, medium, and high (or chaos). All those observations with a ratio below the 20th percentile are in the first group. These are cases of no crowding, or very low ratio of patients to midwives. The second group includes those observations between the 20th and 80th percentiles, and are categorized are cases with some crowding. Finally, the last group consists of all those above the 80th percentile. These are situations of very high ratios of patients to midwives. The cut offs for these groups are reported in the bottom of Table 3.3. In these models, the lowest quintile (low staffing is considered the reference group.[28] Table D.1 shows the coefficients of a regression of each of the pre-treatment controls on the ratio of patients-to-midwives. The lack of statistical significance for all cases provides support to the exogeneity assumption of my measure of staffing. Furthermore, for the non-linear specification, Table D.2 shows that the mean of the pre-treatment characteristics are not statistically different across the three groups of staffing (low, medium and high). Again, this emphasizes the strength of the quasi-natural experiment.

The last part of the analysis aims at understanding the mechanisms through which physicians decide to recommend some patients to change delivery method. Two hypotheses are tested. First, it could be the case that high values of the ratio of patients-to-midwives results in less midwifery time available for each patient. Under this scenario, patients who were admitted with an already higher risk of C-section (and that need more care) will be the most affected. At higher ratios, the probability of C-section should rise

is born with low weight (less than 2,500 grams), and a dummy for whether the mother had at least one emergency checkup during pregnancy.

27 This means that all the variation in this specification comes from within same day of the week and shift. For example, I would be comparing a mother who arrived on a Tuesday afternoon shift and finds many patients waiting with another woman arriving a different Tuesday afternoon but who observes few patients waiting.

28 See Appendix C for a more detailed discussion on model selection, where models of different polynomial degrees and categorical definitions of workload are tested.

faster for this group than for other patients -all else constant- due to their pre-treatment lower health. These patients with a higher risk are identified as those with extreme birthweight (below 2,500 grams or above 4,000 grams) or with at least one emergency visit to the hospital during pregnancy.

The second hypothesis has to do with agency discrimination. When resources are constrained, e.g. high ratio, physicians may see optimal to shift some patients to the operative theater and perform a C-section. This would reduce the workload on midwives by reducing the number of patients waiting in the delivery room. Because patients are heterogenous, physicians will find it easier to offer this treatment to some patients than others. This paper uses the patient's civil status as a proxy to whether the she is alone in the delivery room.[29] The assumption here is that, on average, single women are more likely to be alone in the delivery room.[30] In those cases, the physician only needs to convince one person about the change in procedure -not to mention the patient is in labor and in a lot of pain, which makes harder to analyze the pros and cons of each alternative-.

To test whether physicians' treatment covaries with the patients' characteristics above mentioned, estimate the following regression:

$$y_{it} = \alpha + \beta_1 R_{it} + \beta_2 R_{it} \times D_{it} + \beta_3 D_{it} + \theta X_{it} + \gamma_t cs_{it} + dow \times shift_t + year_t + month_t + \epsilon_{it} \tag{3}$$

where D_{it} is either one of two variables: an indicator for whether the patient has a high-risk of C-section, or whether she is single. The remaining variables are defined as in Eq. (2), adding civil status as a control. I expect high-risk and single patients to be more affected by a high ratio of patients, hence, a positive β_2 in both cases.

4. RESULTS

Table 4.1 presents the results of estimating Eq. (2). Starting from a regression with only the covariate of interest and fixed effects for year, month and day of the week in the first column, each remaining column sequentially adds more controls. The second column adds controls for mother and pregnancy characteristics, the third adds the number of scheduled

[29] This variable is constructed only with married and single women. For the sake of clarity, all women outside these two categories (divorced, separated and widows) are not considered.
[30] For a single woman in Tuscany, the odds of being alone in the delivery room are 1.25 times larger than the odds for a married woman being alone (ARS Toscana 2013).

C-sections taking place during patient's labor, the fourth column includes hour of admission fixed effects, while the last one instead uses shift-of- admission interacted with day-of-the-week fixed effects. This last model is the preferred one since it accounts for possible supply changes shift and day that may take place in the ward (apart from midwives). To save space, only the coefficients of treatment are included, but results for other co- variates are comparable to previous studies.[31] The numbers in parentheses in the table are standard errors. The average value of each dependent variable is included at the bottom of each panel to help understand whether coefficients are economically important. For all remaining estimations in this paper I will use the specification model in column (5).

Panel (A) of the table reports results for the Ratio of patients to midwives as a continuous variable, and Panel (B) reports results using a dummy variable for different levels of workload in order to test for non-linearities. First thing to notice is that coefficients across columns (models) only change in the third decimal. This is a good sign of exogeneity of the ratio of patients-to-midwives. Although the coefficient for the linear specification is not statistically significant, in the second panel the probability of having a C-section is about 4 p.p. (34%) higher for those who face a ratio of patients-to-midwives in the middle of the distribution compared to those in the reference group. For those patients arriving when the ratio of patients-to-midwives is very high (last quintile), there is not statistically significant effect on the probability of C-section. This may be due to some capacity constraints on the operative theater when workload is at its highest levels.

This effect would imply a 5.7% (or 1.2 p.p.) rise in total C-sections (scheduled and unscheduled), which is economically important and reasonable when compared with previous studies looking at all C-sections and changes in monetary compensation. *Allin et al.* (2015) find that doubling the compensation for a C-section relative to a vaginal delivery increases the likelihood that a physician opts for the former by just more than 5 p.p., all else equal. Gruber *et al.* (1999) suggests that cesarean delivery rates would rise by 3.9% in response to each $100 increase in the compensation received for a C-section, all else equal.

Table A.7 presents results of the effect by whether the patient arrived in a weekday or weekend, and by shift of admission. The estimations are very imprecise due to the few number of observations in each cell, and render all differences insignificant. Nevertheless, point estimates are slightly higher in weekends, as well as for admissions during the morning shift. Table A.8 shows results for a robustness check where I measure effective staff level at different points in time between a patient's admission and delivery. The effect

[31] See full regressions in Appendix D, Table D.7 and Table D.8.

of congestion disappears the further away from admission it is measured, which can be a result of the endogeneity issue mentioned before: physicians can adjust the timing of births. Finally, Table A.9 presents results for a placebo test where workload is measured 24 hours after admission (instead at admission as before). As expected, for all different specifications, the placebo is always statistically and clinically insignificant.

Table 4.1. Effect of effective staffing on the Probability of C-section

	(1)	(2)	(3)	(4)	(5)
Panel (A)					
Ratio patients	0.0019	0.0038	0.0046	0.0007	0.0004
to midwives	(0.0074)	(0.0074)	(0.0074)	(0.0083)	(0.0085)
Panel (B)					
20-80th Percentile	0.0432***	0.0441***	0.0452***	0.0410***	0.0408***
	(0.0149)	(0.0148)	(0.0148)	(0.0154)	(0.0156)
>80th Percentile	0.0193	0.0238	0.0261	0.0187	0.0197
	(0.0196)	(0.0195)	(0.0196)	(0.0210)	(0.0213)
Observations	2,613	2,613	2,613	2,613	2,613
Mean dep.	0.119	0.119	0.119	0.119	0.119
Time FE	✓	✓	✓	✓	✓
Controls		✓	✓	✓	✓
Other patients			✓	✓	✓
Hour FE				✓	
Shift*dayofweek FE					✓

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$
Reported coefficients are average marginal effects.

4.1 How do physicians choose which patients to send to the operative theater?

This part of the study digs deeper into the mechanisms behind the effect of staffing on the rate of C-sections. As mentioned before, two hypotheses are tested. First, low-staffing means there is less midwifery-time available for each patient, which may result in more patients needing C-section due to the lack of proper care. This effect should be higher for those patients who were admitted with an already higher risk of C-section. Secondly, physicians and midwives may actively decide to perform a C-section on some patients in moments of low-staffing to reduce the number of patients in the delivery room. In this case I expect patients with lower bargaining power -which I proxy by civil status- being more treated than others.

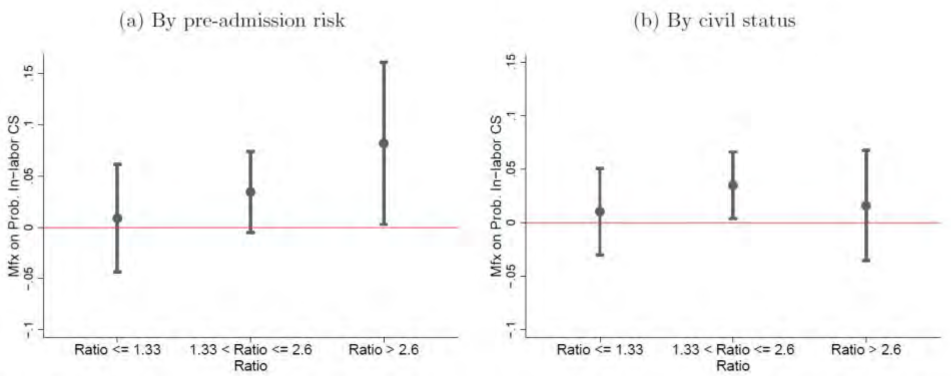
Table 4.2 reports the average marginal effects obtained for each group from estimating Eq. (3). As expected, a higher number of patients per midwife rises the probability of C-section more for single patients but not for married ones. Points estimates suggest that high-risk patients are more affected by workload than low-risk patients, although the only statistically significant coefficient is for this group. However, estimates are very imprecise.

Table 4.2. Effect of effective staffing on the Probability of C-section - by groups

	Low-risk	High-risk	Married	Single
20-80th Percentile	0.0330*	0.0587	0.0270	0.0516**
	(0.0175)	(0.0368)	(0.0202)	(0.0247)
>80th Percentile	0.0023	0.0755	0.0125	0.0182
	(0.0229)	(0.0557)	(0.0265)	(0.0332)
Observations	2,118	495	1,300	1,028
Mean	0.11	0.15	0.11	0.13

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$. Reported coefficients are average marginal effects from a regression of the probability of C-section on the interaction of treatment, a variable for being high risk and a variable for being single. The number of observations when using marital status is slightly smaller because the variable is missing for 11% of the working sample.

Figure 4.1. Difference in the effect of staffing on the probability of C-section by type of patient



Note: Dots are the average marginal effect of whether the patient is high-risk (a) or single (b). Bars are 90% confidence intervals.

Another way to look at it is by comparing the average marginal effects of being high-risk and single, across the different levels of the ratio of patients-to-midwives. This can be seen in Figures 4.1a and 4.1a respectively. Note that the effect of staffing, in both cases, is not statistically significant when the ratio is low. For the comparison based on ex-ante risk, the point estimate for the difference in the probability of C-section between the two groups gets higher with workload -albeit not statistically significant-. This is reasonable since ex-ante high-risk patients for whom the marginal benefit from midwives' attention is higher.

Instead, for the case of married vs. single mothers, the difference is statistically significant only for those in the middle of the distribution, but goes down again when workload is high. At high levels of workload, it is more likely that capacity constraints in the operative theater emerge as well. These "extra" C-sections only based on midwives' workload and not due to patients' health-status should go down during the busiest times.

4.2 The effect on other interventions and morbidity outcomes

The estimates above demonstrate that, when the ratio of patients-to-midwives is high, physicians send some patients to the operative theater to have a C-section. These patients are typically patients with a higher-risk of needing a C-section, or single women. However, are physicians using their high bargaining power to transfer some patients so midwives can provide give better care for the remaining patients? To test this, I estimate Eq. (3) again but now the outcome variable is one of the five indicators of morbidity and interventions mentioned before. If a high ratio lowers the quality of care, then those type of patients who are not likely to be sent to the operative theater would be the ones more affected by this.

In the economics literature the most commonly studied health outcomes for births are: weight, fetal mortality and maternal mortality. Nevertheless, both maternal and fetal deaths are extremely rare events (4 per 100,000 births and 2.7 per 1,000 births respectively for Italy). In the case of weight- at-birth, because treatment here is defined at the moment of admission to the hospital, it is considered a pre-defined outcome (not affected by treatment).[32]

The restricted-use version of the birth certificates in hand contains, however, some other measures of health and registers of medical interventions that are associated with

[32] In fact weight at birth is one of the variables used to assess the balancing of the sample between treatment and control groups.

health outcomes. The measures that occur in at least 1% of births are: having an operative birth[33], length-of-stay after birth (LOS), whether the newborn was transferred to a neonatal intensive care unit (NICU), no skin-to-skin contact, lack of exclusive breastfeeding, and whether the newborn had an APGAR score below 9.[34] A higher probability of needing NICU, having an operative birth[35] or a longer time in the hospital during crowded times can be signals of lower quality. Similarly, if human resources are scarce, physicians may decide to skip some steps of the service considered important but not essential. For example, they may decide that helping the newly mother achieve skin-to-skin contact with her newborn is not as important as helping another woman in labor to deliver. The same reasoning applies for not giving exclusive breast-feeding.

While it is clear why a higher probability of going to NICU having a low APGAR score, or staying longer in the hospital are not desirable, there are also compelling arguments regarding the importance of the remaining set of outcomes. In a systematic review, Ip *et al.* (2007) finds that breastfeeding is associated with both decreased risk for many early-life diseases and conditions as well as with health benefits to women.[36] At the same time, skin-to-skin contact has been shown to increase the probability and length of exclusive breastfeeding (Moore *et al.* 2007), as well as substantially reducing neonatal mortality amongst preterm babies in hospital (Lawn *et al.* 2010). In the case of operative births, even though it is still widely used, this delivery method is becoming less popular due to some evidence showing it increases maternal morbidity and can cause significant fetal morbidity (Ali and Norwitz 2009; Murphy *et al.* 2011; Towner *et al.* 1999).

Table 4.3 displays the average marginal effects for each of the four groups of women (high and low risk, married and single), and for the five outcomes above mentioned. Estimates are quite imprecise given the small sample size and the rarity of these morbidities. However, there is a statistically significant, large and positive effect of the high ratios of patients-to-midwives on the probability of not achieving skin-to-skin contact with the infant. Furthermore, this effect is only present for married patients, who are not more likely to get surgery when workload rises. These patients are between 24% and 35% more likely to not attain skin-to-skin contact with their newborn when the number of patients

[33] Operative vaginal delivery refers to a delivery in which the physician uses forceps or a vacuum device to assist the mother in transitioning the fetus to extra-uterine life.
[34] The Apgar score is a method used to quickly summarize the health of newborn children. The Apgar scale is determined by evaluating the newborn baby on five simple criteria on a scale from zero to two, then summing up the five values thus obtained. The resulting Apgar score ranges from zero to 10.
[35] A higher likelihood for operative birth has been linked to scarce or absent midwifery care and the presence of obstetrician or physicians instead (Hatem *et al.* 2008).
[36] "Breastfeeding and Maternal and Infant Health Outcomes in Developed Countries", AHRQ Publication No. 07-E007, April 2007.

per midwife is higher. This provides further evidence of the fact that, by shifting delivery method for some patients, physicians are avoiding some bad outcomes to occur.

Table 4.3. Effect of effective staffing on other health outcomes

	Low-risk						High-risk					
	Op. birth	LOS*	NICU§	No s2s‡	NEB†	Apgar<9	Op. birth	LOS*	NICU§	No s2s‡	NEB†	Apgar<9
20-80th Percentile	-0.0164	-0.0246	-0.0134	0.0298	-0.0325	-0.0096	-0.0616	-0.0085	-0.0013	0.0133	-0.0392	0.0070
	(0.0194)	(0.0184)	(0.0123)	(0.0216)	(0.0296)	(0.0104)	(0.0389)	(0.0414)	(0.0304)	(0.0497)	(0.0629)	(0.0298)
>80th Percentile	0.0257	-0.0380	-0.0064	0.0426	-0.0179	-0.0011	-0.0213	-0.0168	-0.0357*	0.0329	-0.0592	-0.0075
	(0.0264)	(0.0234)	(0.0160)	(0.0287)	(0.0384)	(0.0144)	(0.0270)	(0.0274)	(0.0195)	(0.0310)	(0.0422)	(0.0181)
Observations	2,613	2,521	2,609	2,297	2,044	2,613	2,613	2,521	2,609	2,297	2,044	2,613
Mean dep.	0.133	4.274	0.0728	0.193	0.360	0.0463	0.133	4.274	0.0728	0.193	0.360	0.0463

	Married						Single					
	Op. birth	LOS*	NICU§	No s2s‡	NEB†	Apgar<9	Op. birth	LOS*	NICU§	No s2s‡	NEB†	Apgar<9
20-80th Percentile	-0.0339	-0.0061	0.0124	0.0449*	-0.0183	-0.0033	-0.0213	-0.0168	-0.0357*	0.0329	-0.0592	-0.0075
	(0.0241)	(0.0222)	(0.0140)	(0.0265)	(0.0367)	(0.0124)	(0.0270)	(0.0274)	(0.0195)	(0.0310)	(0.0422)	(0.0181)
>80th Percentile	0.0096	-0.0194	0.0111	0.0699**	-0.0517	0.0117	-0.0119	-0.0159	-0.0086	0.0416	0.0210	-0.0142
	(0.0329)	(0.0291)	(0.0171)	(0.0345)	(0.0472)	(0.0178)	(0.0359)	(0.0342)	(0.0272)	(0.0432)	(0.0559)	(0.0235)
Observations	2,328	2,264	2,324	2,057	1,826	2,328	2,328	2,264	2,324	2,057	1,826	2,328
Mean dep.	0.133	4.274	0.0728	0.193	0.360	0.0463	0.133	4.274	0.0728	0.193	0.360	0.0463

Reported coefficients are average marginal effects. Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$
⋆LOS: Length-of-stay after birth (in log-hours); §NICU: Neonatal Intensive Care Unit; ‡No s2s: No skin-to-skin contact; †NEB: Non-exclusive breastfeeding.

4.3 Other possible channels?

Beyond the mechanisms mentioned in the previous section, there are -at least- two more channels that can explain the rise in C-sections along with the rise of the ratio of patients-to-midwives. The first and most obvious option is that patients who are admitted in low and high staffing times are different. Nevertheless, all tests performed in this study and previous research support the idea that, for those patients attempting a vaginal delivery, their time of arrival to the hospital is randomly distributed across the day and week.

The other possible explanation is that those type of patients who get these 'extra' C-sections actually prefer this delivery method. However, because the focus is exclusively on in-labor C-sections, the above estimates correspond to women who have already agreed on attempting labor in the process to attempt a vaginal delivery. Hence the effect is more likely to arise from decisions made in the delivery room regarding when to stop labor and change treatment, than from maternal preferences for C-sections. Nevertheless, because data comes from a public hospital, patients may be denied an elective C-section -even when preferred- if there is no medical reason for it. Hence it is not possible to totally rule out that some demographic groups may be more inclined towards having a C-section and physicians internalize this when deciding which patient is send to surgery.

4.4 Can these 'extra' C-sections be avoided?

Results above suggest that physicians do more surgeries when staffing is low. First-time mothers facing a ratio of patients-to-midwives between 1.33 and 2.66 are 4 p.p. (or 34%) more likely to have an in-labor C-section. A policy to eliminate overcrowding from maternity wards would have a very significant effect on the already high levels of C-sections seen in Italy. How to do that is not clear.

Considering only the hospital used in the analysis, in the absence of crowding, the "extra costs" for the public health system is of about €17,700 a year.[37] This is of course not enough to hire the necessary number of midwives to avoid low-staffing situations. Of course, this analysis is not complete since one should include other costs, like the drop in skin-to-skin contact when staffing is low, or the other non-financial costs of C-sections mentioned in the introduction of this study.

[37] Back of the envelope calculations suggest that there are about 86 'extra' C-sections in the 4 years in the sample due to crowding. According to the prices on acute interventions published by the Italian Ministry of Health, a vaginal delivery without complication is rated at €1,272, while a C-section costs €2,092. Hence the difference (€820 time the number of extra C-sections (107) divided by the number of years (4) gives €17,700.

Another possible policy is to concentrate maternity wards in fewer but bigger units and benefit from the economies of scale emerging. The larger the population a hospital serves, the lower the coefficient of variation of demand, and hence the higher the occupancy rate (Long and Feldstein 1967). For the hospital in case this may not really be a suitable alternative since it is already a large maternity ward and the only on its city.

5. CONCLUSIONS

In this paper I use a natural experiment set up -that patient characteristics is orthogonal to the level of staffing at the hospital at the moment of admission- and detailed data on births to estimate the impact of staffing on physician's treatment decisions. More specifically, I investigate whether different levels of midwifery effective staffing (patients-to-midwives) influence the probability that a patient will be sent to have a cesarean section. The contribution is threefold. First it proposes an innovative empirical approach that allows me to estimate physician's responses to exogenous shocks to effective staffing. Second, it provides suggestive evidence that physicians do not choose at random which patients to over-treat, but may instead use their bargaining power. Lastly, it brings to light yet another cause for the high C-section rates we see today: low effective staffing.

Focusing exclusively on patients attempting labor and vaginal delivery, this study finds that first- time mothers who -at admission- face a ratio of patients-to-midwives higher than 1.3 are about 34% (or 4 p.p.) more likely to change delivery method. There are two type of patients who are more affected by this. First, patients who upon admission have an already higher risk of C-section are more likely to develop complications due to limited care when few midwives are available. Secondly, single women, due to their lower bargaining power. I provide evidence that physicians may decide to induce some patients towards having a C-section to speed up the delivery and release the pressure on midwives in the delivery room. In summary, the evidence provided here suggests that physicians' way to deal with an exogenous shock in demand (patients) is to induce some patients towards an intervention that is faster, maximizing the aggregate health in the maternity ward.

My estimates imply that total number of C-sections for first-time mothers could be reduced by about 5.7% (1.2 p.p.) if situations of low-staffing are avoided. This would be a very important achievement given the already overly high rates of C-sections observed in developed countries. Nevertheless, it is not clear that public healthcare systems can quickly afford to tackle this issue.

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APPENDIX A: THE WORKING SAMPLE AND SCHEDULED PATIENTS

The working sample used in the main paper is restricted to only those unscheduled patients who at- tempt to have a vaginal delivery after going through labor and leaves out scheduled patients. Scheduled patients can be further divided in two groups: (i) elective C-sections, and (ii) pharmacologically- induced patients. This appendix shows evidence of how the latter group's transition through the maternity ward resembles more that of elective C-section rather than the one of unscheduled patients, and hence should not be included in the working sample.

One important caveat of the data is that one cannot disentangle scheduled from unscheduled patients among those who were pharmacologically induced. However, anecdotal evidence from the ward's staff suggest that most of them are scheduled (e.g. overdue pregnancy). Furthermore, a descriptive analysis of the data seems to corroborate that. Figures A.1 and A.2 present the distribution of patients across hours and days as performed in section 1.3.1 of the main paper except that now scheduled patients are further divided between elective C-sections and induced. Starting from Figure A.1, it shows that there is a pick in admissions for both elective C-sections and induced patients during the afternoon shift, and then again, a pick in time of birth (although the pick is later in the day for induced patients relative to the elective C-sections). Nevertheless, the picks are less pronounced for induced patients, suggesting that some of them may be arriving at random hours of the day like unscheduled patients do.

Even though the distribution by hours of induced patients seem to follow that of elective C-sections, their distribution by day of the week instead is closer to that of unscheduled patients. Even though admissions are slightly lower during weekends, births are evenly distributed across all days of the week. This is probably due to the fact that, as long as everything goes well, these patients are taken care of by midwives (not physicians).

The evidence provided in this appendix supports the idea of excluding both elective C-sections and pharmacologically induced patients from the working sample, but to include the latter group in the treatment variable given that they are primordially seen by midwives.

Figure A.1. Distribution of admissions and births by hour

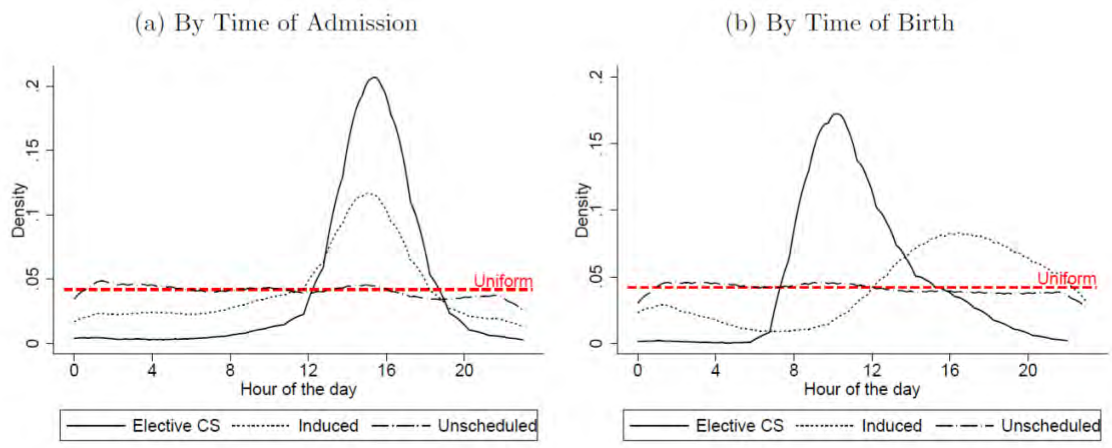
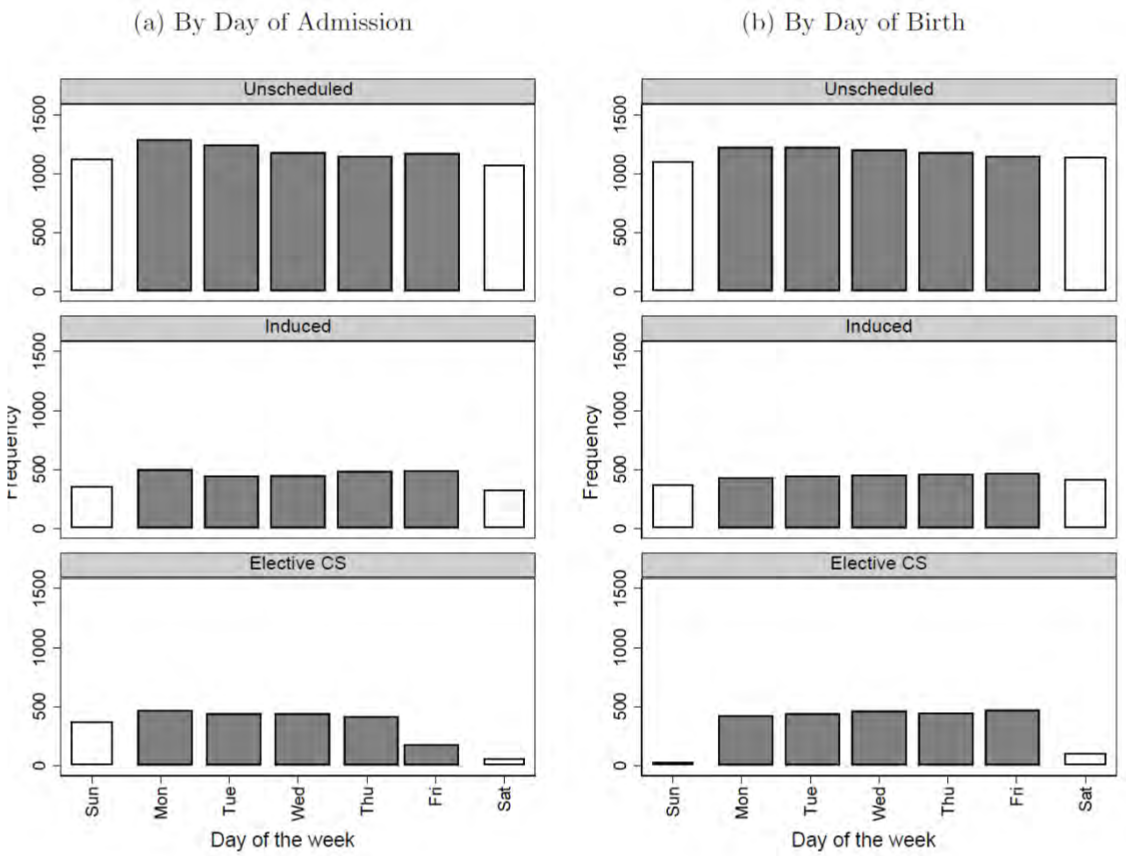


Figure A.2. Frequency of admissions and births by day



APPENDIX B: A MEASURE OF WORKLOAD WITHOUT ADJUSTING FOR SUPPLY SIDE FACTORS

The measure of workload used in the main paper is the ratio of patients-to-midwives, hence it takes into account both demand and supply side effects. Specifying the covariate of interest as a ratio may put some constraints on the estimated coefficient. This appendix repeats the main estimations but using instead the number of unscheduled patients waiting to give birth (without adjusting for the number of midwives).

Figures B.1 shows a histogram of the number of unscheduled patients observed by each patient at admission. The mode is 3, and the mean is slightly above at 3.34. As in the main paper, I divide this variable in quintiles to test for non-linearities in its effect on the probability of C-section. Table B.1 describes the number of observations and limits for each quintile.

Figure B.1. Histogram unscheduled patients

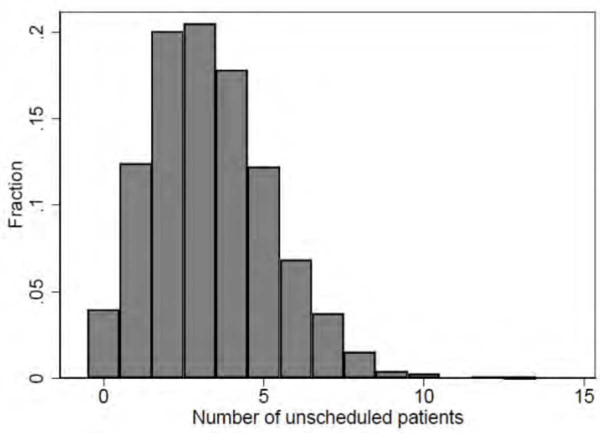


Table B.1: Descriptive statistics of quintiles

Quintile	N	Min	Max
1	953	0	2
2	536	3	3
3	466	4	4
4	320	5	5
5	338	6	13

Finally, Table B.2 presents the results from running the preferred model using the number of unscheduled patients as regressor. Similar to the findings in the main paper, there seems to be a non-linear relationship between workload and the probability of C-section. This effect starts to rise already in the second quintile and slowly declines in the fourth and fifth quintiles. This provides more assurance to the results using the ratio of patients-to-midwives.

Table B.2. Average marginal effect on probability of C-section

	Linear	Non-linear
Number of unscheduled	0.0044	
patients	(0.0036)	
2nd Quintile		0.0309*
		(0.0177)
3rd Quintile		0.0503**
		(0.0200)
4th Quintile		0.0384*
		(0.0223)
5th Quintile		0.0266
		(0.0212)
Observations	2,613	2,613
Mean dep.	0.119	0.119

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$

APPENDIX C: ROBUSTNESS TO ALTERNATIVE MODELS

In the main paper two functional forms are tested for the effect of workload on the probability of C-section: a linear specification, and a non-linear one using a categorical variable constructed from the 20th and 80th percentiles. This appendix elaborates further on the model selection and tests other specifications. Columns (1) to (4) in Table C.1 present the coefficients for different polynomial degrees of the ratio of patients-to-midwives, with the Akaike Information Criteria (AIC) reported at the bottom. It seems that, within these polynomial functional forms, the data at hand is better represented by either a squared or cubic polynomial, given their statistical significance and their low AIC.

Column five presents results using a categorical variable with the quintiles of the distribution of the ratio of patients-to-midwives (where the reference group is the first quintile). This specification gives the model more flexibility to fit the data, at the cost of estimating more coefficients. Results suggest that there is a sudden rise in the probability of C-section for patients who see a ratio of patients- to-midwives in the second quintile, which then falls slowly until the fifth quintile where is no longer statistically distinguishable from the reference group. This decay in the probability of C-section for higher workloads may be associated with capacity constraints on the operative theater (beds, number of gynecologists, etc.).

Given the previous, I created a variable with three categories where the 3 middle quintiles of the ratio of patients-to-midwives have been coded together in one group (<20th percentile, 20-80th percentile, >80th percentile). This specification has the advantage of capturing the higher level of C-sections that occurs in the middle of the workload distribution, while diminishing the number of coefficients to be estimated and augmenting precision. Results are presented in the sixth column.

Table C.1. Alternative model specifications

	(1)	(2)	(3)	(4)	(5)	(6)
Ratio	0.0004	0.0524**	0.1195**	0.1627		
	(0.0085)	(0.0240)	(0.0535)	(0.1283)		
Ratio Square		-0.0104**	-0.0369**	-0.0633		
		(0.0044)	(0.0182)	(0.0719)		
Ratio Cubic			0.0030*	0.0091		
			(0.0018)	(0.0157)		
Ratio Quadratic				-0.0005		
				(0.0011)		
2nd Quintile					0.0554***	
					(0.0199)	
3rd Quintile					0.0329*	
					(0.0196)	
4th Quintile					0.0299	
					(0.0211)	
5th Quintile					0.0167	
					(0.0216)	
20-80th Percentile						0.0408***
						(0.0156)
>80th Percentile						0.0197
						(0.0213)
Observations	2,613	2,613	2,613	2,613	2,613	2,613
AIC	1544.59	1542.58	1543.18	1545.06	1541.89	1539.65

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$

Appendix D: Other Graphs and Tables

Figure D1-2. Distribution of admissions and births

Figure D.1: Density distribution of Ratio

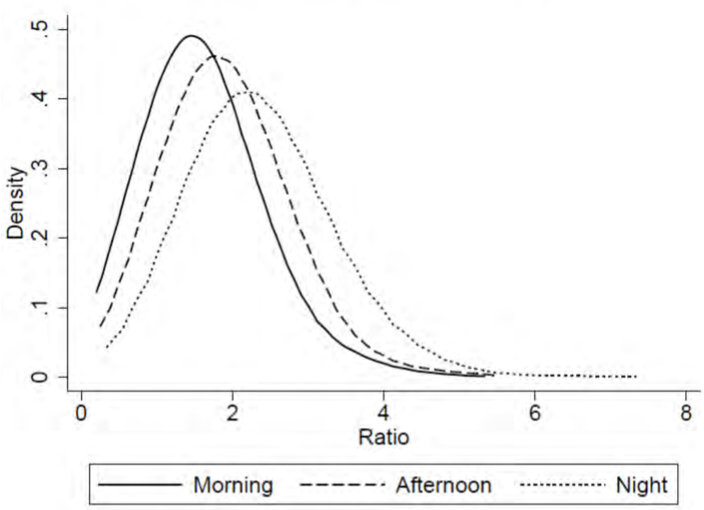
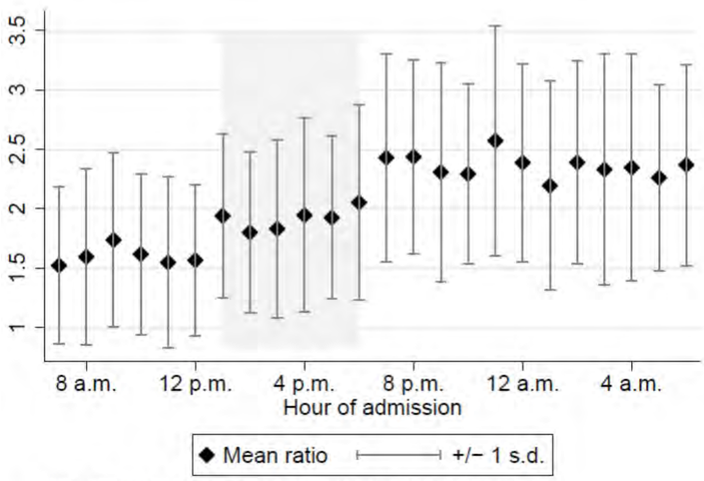


Figure D.2: Ratio by time of admission



Note: Shaded area marks the afternoon shift (1 pm – 7 pm)

Table D.1. Regression of pre-treatment characteristics on Ratio of patients-to-midwives

Dependent variable	Coef. of Ratio
Mother's characteristics	
With university degree	-0.0141
	(0.0129)
Above 36 yo	-0.0033
	(0.0117)
Pregnancy's characteristics	
Preterm (before 37th week)	-0.0080
	(0.0057)
At least 1 ER visit	-0.0005
	(0.0076)
Newborn's characteristics	
Male	-0.0045
	(0.0132)
Low weight at birth	-0.0054
	(0.0055)
Observations	2,613

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$

Table D.2. Pre-treatment variables balanced across treatments and control

	Level of Ratio				
	<20th Percentile	20-80th Percentile	>80th Percentile	(1) vs. (2)	(1) vs. (3)
Mother's characteristics					
% of mothers with university degree	0.386	0.358	0.340	0.271	0.112
	(0.022)	(0.013)	(0.018)		
% older than 36 yo	0.302	0.288	0.266	0.567	0.182
	(0.021)	(0.012)	(0.017)		
Pregnancy's characteristics					
% of births before 37 weeks of gestation	0.057	0.058	0.040	0.942	0.191
	(0.011)	(0.006)	(0.008)		
% of pregnancies with at least 1 ER visit	0.110	0.123	0.100	0.431	0.597
	(0.014)	(0.009)	(0.012)		
Newborn's characteristics					
% of male newborns	0.508	0.517	0.497	0.755	0.704
	(0.023)	(0.013)	(0.019)		
% of low-weight newborns (<2,500 grams)	0.055	0.051	0.042	0.746	0.305
	(0.010)	(0.006)	(0.008)		

Standard errors in parentheses. $***p < 0.01$, $**p < 0.05$, $*p < 0.1$

Table D.3. Probability of C-section using a Linear Probability and Probit Model

	LPM	Probit
Panel (A):		
Ratio patients to midwives	0.0007	0.0012
	(0.0083)	(0.0081)
Panel (B):		
20-80th Percentile	0.0410***	0.0437***
	(0.0154)	(0.0166)
>80th Percentile	0.0187	0.0222
	(0.0210)	(0.0221)
Observations	2,613	2,613
Mean dep.	0.119	0.119

Robust standard errors in parentheses
$***p < 0.01$, $**p < 0.05$, $*p < 0.1$

Table D.4. LPM of C-section by day and staff shift

	All	Weekday	Weekend	Morning	Afternoon	Night
20-80th Percentile	0.0410***	0.0377**	0.0608*	0.0467*	0.0341	0.0242
	(0.0154)	(0.0176)	(0.0325)	(0.0270)	(0.0300)	(0.0245)
>80th Percentile	0.0187	0.0195	0.0251	0.0054	0.0477	0.0014
	(0.0210)	(0.0245)	(0.0429)	(0.0574)	(0.0523)	(0.0282)
Observations	2,613	1,883	730	636	641	1,336
Mean dep.	0.119	0.114	0.132	0.104	0.137	0.118

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$

Table D.5. Effect of staffing for different windows of time since admission

	1 hour	2 hours	3 hours	4 hours	5 hours	6 hours
20-80th Percentile	0.0408***	0.0490***	0.0469***	0.0399**	0.0308*	0.0241
	(0.0156)	(0.0156)	(0.0163)	(0.0162)	(0.0164)	(0.0164)
>80th Percentile	0.0197	0.0213	0.0320	0.0136	0.0100	0.0077
	(0.0213)	(0.0209)	(0.0222)	(0.0220)	(0.0217)	(0.0216)
Observations	2,613	2,613	2,613	2,613	2,613	2,613

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$

Table D.6. Effect ofstaffing at admission and 24hs after

	Workload at admission		Workload 24hs after admission	
Panel (A)				
Ratio patients	0.0019	0.0004	0.0020	-0.0010
to midwives	(0.0074)	(0.0085)	(0.0073)	(0.0081)
Panel (B)				
20-80th Percentile	0.0432***	0.0408***	-0.0019	-0.0065
	(0.0149)	(0.0156)	(0.0160)	(0.0161)
>80th Percentile	0.0193	0.0197	0.0011	-0.0091
	(0.0196)	(0.0213)	(0.0202)	(0.0213)
Time FE	✓	✓	✓	✓
Controls		✓		✓
Other patients		✓		✓
Shift*DOW FE		✓		✓

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$

Table D.7. LPM using a continuous measure of staffing

	(1)	(2)	(3)	(4)
Ratio patients	0.0019	0.0038	0.0046	0.0007
to midwives	(0.0074)	(0.0074)	(0.0074)	(0.0083)
University degree		-0.0178	-0.0180	-0.0167
		(0.0134)	(0.0134)	(0.0135)
Old (>34 years old)		0.0489***	0.0492***	0.0510***
		(0.0153)	(0.0153)	(0.0153)
Male		0.0351***	0.0352***	0.0348***
		(0.0125)	(0.0125)	(0.0126)
Preterm (<37 wofg)		0.0079	0.0089	0.0055
		(0.0326)	(0.0326)	(0.0329)
Low weight at birth (< 2.5kg)		0.0284	0.0287	0.0318
		(0.0350)	(0.0350)	(0.0348)
ER visit		0.0062	0.0066	0.0116
		(0.0201)	(0.0201)	(0.0202)
>6 hours waiting		-0.0052	-0.0031	-0.0028
		(0.0148)	(0.0155)	(0.0156)
>14.2 hours waiting		0.0458***	0.0516**	0.0496**
		(0.0160)	(0.0204)	(0.0210)
Observations	2,613	2,613	2,613	2,613
Yea, Month, Dow FE	✓	✓	✓	✓
Controls		✓	✓	✓
# Scheduled CS			✓	✓
Hour FE				✓

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$

Table D.8. LPM using a categorical variable for staffing

	(5)	(6)	(7)	(8)
20-80th Percentile	0.0530***	0.0545***	0.0555***	0.0555***
	(0.0155)	(0.0154)	(0.0155)	(0.0166)
>80th Percentile	0.0413**	0.0467**	0.0485**	0.0477**
	(0.0187)	(0.0187)	(0.0188)	(0.0214)
University degree		-0.0172	-0.0175	-0.0162
		(0.0133)	(0.0133)	(0.0135)
Old (>34 years old)		0.0490***	0.0494***	0.0507***
		(0.0153)	(0.0153)	(0.0153)
Male		0.0350***	0.0351***	0.0346***
		(0.0125)	(0.0125)	(0.0126)
Preterm (<37 wofg)		0.0070	0.0079	0.0049
		(0.0324)	(0.0324)	(0.0327)
Low weight at birth (< 2.5kg)		0.0294	0.0296	0.0328
		(0.0350)	(0.0350)	(0.0347)
ER visit		0.0051	0.0055	0.0103
		(0.0201)	(0.0201)	(0.0201)
>6 hours waiting		-0.0045	-0.0023	-0.0017
		(0.0148)	(0.0155)	(0.0156)
>14.2 hours waiting		0.0471***	0.0532***	0.0523**
		(0.0159)	(0.0204)	(0.0209)
Observations	2,613	2,613	2,613	2,613
Yea, Month, Dow FE	✓	✓	✓	✓
Controls		✓	✓	✓
# Scheduled CS			✓	✓
Hour FE				✓

Robust standard errors in parentheses. $***p < 0.01, **p < 0.05, *p < 0.1$

GIANNI GHETTI

MODEL FOR THE ESTIMATION OF SOCIETAL COSTS FOR PERTUSSIS IN ITALY

Abstract. Objectives. To estimate age-specific costs associated with pertussis in Italy and the related Societal burden in the last decade (2006-2015), starting from available epidemiological data and healthcare charges. **Methods**. Age-specific data on notified pertussis cases were corrected to account for under-notification and hospitalization rates were estimated based on Italian data in the same period. Complications (pneumonia, seizures, and encephalopathy) rates were derived from a US study and adjusted. Assuming all complications are hospitalized, hospitalization-without-complications rates were estimated as a difference. The remaining part was considered outpatient cases. For outpatient, general practitioner (GP) consultations and antibiotics prescription rates were taken from a Dutch study. National DRG tariffs were used for inpatients; GP consultation cost was estimated from an Italian study and antibiotics (erythromycin, clarithromycin, and azithromycin) cost was estimated based on dosage units necessary to cover a full cycle of therapy, considering the formulations available on the Italian market. The Societal impact is measured in term of days of daily activities lost: globally (scenario 1) or medical-related (scenario 2). Age-specific values for days lost were obtained from an Italian study. **Results**. The estimated average direct medical cost ranges from €77 to €1,488 and is highest for infants. Average indirect costs (scenario 1) accounts for 70% of total costs and are lowest for adolescents. The estimated Societal burden amounts to, considering both direct and indirect costs, € 3.74 million per year for pertussis and to € 73.55 million for rotavirus, despite the similar cost per episode. **Conclusions**. The high level of vaccination coverage attained since 2000 has played a key role in diminishing both the incidence of the disease and the related burden for Society, as shown from the comparison with the burden associated with rotavirus.
Keywords. Pertussis, Cost-of-Illness, Societal burden, Italy

1. INTRODUCTION

A report issued by World Health Organization in 2001 (WHO 2002) estimated that, globally, there are 50 million cases of pertussis and 300,000 deaths every year, mostly among infants, in whom the risk of severe morbidity and mortality is highest. In addition, recent reports have noted an increase in the incidence of pertussis among adolescents and adults, in whom it is a considerable source of cough illness.
Several authors have reviewed the epidemiology of pertussis over a long period of time to describe the disease trends and to investigate the role of factors that may affect these trends (Clark *et al.* 2012; Hellenbrand *et al.* 2009; Jackson and Rohani 2014; Rohani and Drake 2011).These studies have focused on the epidemiology of

pertussis since the introduction of the immunization in the mid-1940s and have investigated factors potentially involved in the resurgence of pertussis, including increased awareness, diagnosis, and reporting, changes in vaccine composition or schedule, waning immunity, and evolution of the bacteria.

In addition to understanding the epidemiology and health burden of Bordetella pertussis infection and disease, it is also important to recognize the economic burden it poses.

1.1 Italian background

In Italy, recommendations for pertussis immunization were released in 1961, when whole cell vaccines became available. Nevertheless, vaccination coverage increased substantially only after the introduction of acellular pertussis vaccines in 1995 and, even further, after 2002, when the vaccine started to be offered free of charge by all Italian regions (Rota *et al.* 2005).

Based on routine surveillance data, Italy is currently a low incidence country and outbreaks or incidence peaks have been rarely reported after the achievement of a high immunization coverage (Gonfiantini *et al.* 2014).

The primary objective of this study was to estimate age-specific costs associated with the disease in Italy and describe, based on these results, the burden the Society had to bear in the last decade (2006-2015) because of pertussis.

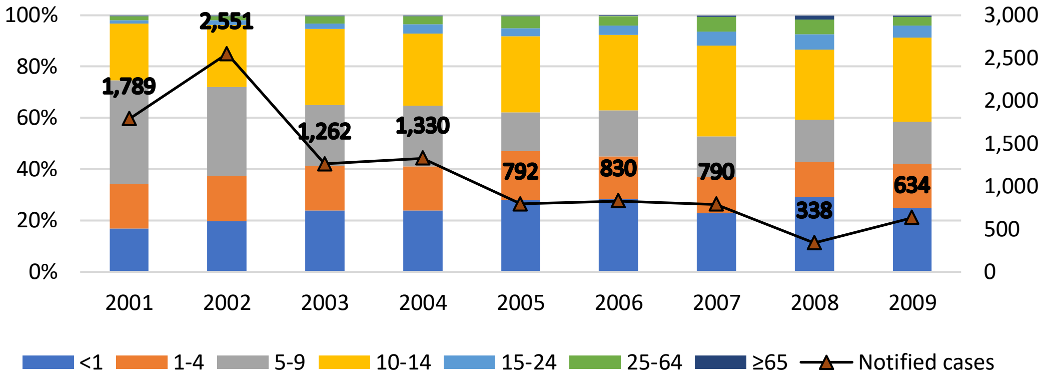
2. METHODS

A micro-simulation model was developed, starting from available epidemiological data and healthcare charges associated with pertussis episodes, to estimate age-specific costs and burden of disease. All the costs, expressed in euros, were inflated to 2016 using the inflation rates for Italy (Eurostat 2018).

2.1 Reported cases

Data on notified pertussis cases in Italy from 2001 to 2009 were obtained from the Ministry of Health, which collects notifications from the Surveillance System for Infectious Diseases in Italy (MdS). For the same period, information on the age percent distribution for the younger (0-14) was obtained from an article reviewing the epidemiology of pertussis in Italy in the last century (Gonfiantini *et al.* 2014) (Fig. 1) summarizes this information.

Figure 1. Age-distribution and number of reported cases of pertussis cases in Italy, 2001-2009



The Italian Surveillance system is passive, universal, mandatory and based on notifications. As a consequence, it is affected by limitations such as under-notification, under-diagnosis, and delay of notification. Moreover, pertussis incidence is under-estimated as many cases in adolescents, young adults, and adults are not identified because of their atypical clinical characteristics and the lack of lab assessment. Lab diagnosis and ad hoc epidemiological studies are not easily performed due to logistic and diagnostic issues (Guiso *et al.* 2011). Gonfiantini and colleagues (2014) found high under-notification in Italy when comparing incidence figures from a sentinel surveillance system with routine reports. Similar results have been reported for Poland (Stefanoff *et al.* 2014). To account for this issue, notified data were adjusted for the factors reported in Table 1.

Table 1. Age-specific under-notification correction factors

Age-class	Correction factor
<1	1.8
1-4	11.8
5-9	9.2
10-14	12.9
15-24	52.3
25-64	77.2
≥65	102.7

2.2 Hospitalization

Data on pertussis hospitalization (primary and secondary diagnoses) in Italy from 2001 to 2009 were obtained from an article analyzing the Italian hospital discharge database (Gabutti *et al.* 2012). Age-specific adjusted hospitalization rates were then estimated, as reported in Table 2.

Table 2. Age-specific hospitalization rates, corrected to account for under-reporting

Age-class	Hospitalization rate
<1	95.40%
1-4	3.09%
5-9	2.50%
10-19	2.09%
≥20	3.08%

2.2.1 Complications

The most frequent complication observed in children with pertussis is pneumonia, which occurs in 6% of cases. Other complications include sinusitis, otitis media, viral and bacterial superinfections, nutritional deficiencies resulting from repeated vomiting and neurologic complications, which are due mostly to hypoxia during coughing spells and apnea (Heininger *et al.* 1997)

In 1990 it was estimated that 50,000 children worldwide experienced long-term neurologic complications of pertussis (Ivanoff and Robertson 1997) and in the late 1990s, it was reported that 0.9 per 100,000 pertussis cases were complicated by encephalopathy (CDC 2002).

The risk of complications is higher among infants than among older children and adolescents. In a prospective case series in Germany, Heininger and colleagues (1997) report that the rate of complications among infants less than 6 months of age was 24%, as compared with 5% among older children.

Gonfiantini and colleagues (2014) report that during the last century, in Italy, pertussis mortality progressively decreased and reached so low numbers that from 1995 to 2001, only one pertussis death per year was reported and no deaths have been reported since 2002.

Given this, the present model considers the following complications: pneumonia, seizures, and encephalopathy.

In the absence of Italian data, assuming that complications rates are similar among similar countries, age-specific complication rates were then obtained from a study summarizing the characteristics and trends in reported cases of pertussis for the years 1980-1989 in the United States (Farizo *et al.* 1992). These rates were then adjusted in order to account for under-reporting (see section 2.1). Assuming that all pertussis cases with complications are hospitalized, age-specific hospitalization-without-complications rates were computed as the difference between the hospitalization rates (see section 2.2) and the sum of adjusted complication rates.

For last, age-specific percentages of outpatient cases were defined as the remaining part of patients. The estimated rates are reported in Table 3.

Table 3. Age-specific rates

Age-class	**Hospitalized patients**				**Outpatients**
	Pneumonia	Seizures	Encephalopathy	No complications	
<1	12.06%	1.67%	0.50%	81.17%	4.60%
1-4	1.04%	0.18%	0.03%	1.84%	96.91%
5-9	0.54%	0.10%	0.04%	1.82%	97.50%
10-19	0.22%	0.03%	0.01%	1.84%	97.91%
≥20	0.03%	0.01%	0.01%	3.03%	96.92%

2.2.2 Day Hospital

In Italy, according to Gabutti and colleagues (2012), pertussis hospitalizations treated in day hospital represent 10% of all hospitalizations (primary diagnosis) for pertussis.

For what concerns hospitalized patients, the model considers all cases with complications and infants (i.e. <1 years old) cases with no complications as treated with ordinary hospitalization. For older inpatients cases with no complications, the percentage of day hospital cases was defined so that the ratio between day hospital cases and total hospitalized cases is 10%.

2.3 Outpatients

The number of general practitioner consultations (GP) per pertussis case and the proportion of patients receiving antibiotics were estimated from a previous study on the disease burden of pertussis in the Netherlands (de Greeff *et al.* 2009) as reported in Table 4 and Table 5, respectively.

Table 4. Age-specific distribution of the number of gp consultations

Number GP consultations	0-9	≥10
1	3%	2%
2	39%	42%
3	38%	41%
4	19%	13%
5	2%	1%

Table 5. Age-specific percentage of outpatients treated with antibiotics

Age-class	% treated with antibiotics
<1	73%
1-9	69%
10-19	45%
20-44	61%
≥45	56%

For hospitalized cases, the model assumes no GP consultations and no antibiotics consumption, besides those received in the hospital.

3. Costs calculation

3.1 Direct medical costs

National Diagnosis-Related Group (DRG) tariffs (MdS 2013) were used to estimate costs related with hospitalized case: pneumonia [DRG 89, 90, 91], seizures [DRG 26, 562, 563], encephalopathy [DRG 22], ordinary hospitalization without complications [DRG 96, 97, 98] and day hospital [DRG 96, 97, 98] costs used in the model are reported in Table 6.

Table 6. National DRG tariffs

Age-class	Pneumonia	Seizures	Encephalopathy	Ord. Hosp. w/o complications	Day Hospital
<18	€ 1,948	€ 1,729	€ 2,989	€ 1,538	€ 185
≥18	€ 3,558	€ 3,289		€ 1,832	€ 197

For what concerns outpatients, even if theoretically the cost of a GP consultation is null, as, in Italy, the National Health Service remunerates GPs mainly on a capitation basis, the model follows the opportunity-cost approach, according to which the value of resources is defined by the benefit foregone when selecting one therapeutic alternative over the next best alternative use. Thus, in the model the cost of a GP consultation is estimated on the basis of Garattini and colleagues (2003) as the mean between the cost of office and home visits weighed for their respective frequencies.

Since a timely laboratory confirmation of pertussis diagnosis is problematic, administering an antibiotic on the basis of a clinical diagnosis is the practice. Antibiotics eradicate Bordetella pertussis from the airway but limit the severity of disease only if started in the catarrhal phase (Tozzi *et al.* 2005). The standard treatment of pertussis has been a full dose of erythromycin. More recently, many national agencies have tended to encourage the use of other macrolides such as clarithromycin and azithromycin.

Thus, the antibiotics considered in the model are erythromycin, clarithromycin, and azithromycin. The distribution of patients among these macrolides therapy, shown in Table 7 was obtained from a Canadian study documenting the morbidity of pertussis in adolescents and adult cases (De Serres *et al.* 2000).

Table 7. Distribution of patients treated with antibiotics among different therapies

Macrolides therapy	% of treatment
Erythromycin	55%
Clarithromycin	36%
Azithromycin	9%

The cost for each antibiotic treatment was estimated on the basis of dosage units necessary to cover a full cycle of therapy (see **Error! Reference source not found.**). The calculation takes into account the different formulations available on the Italian market and the specific dosage for adult and pediatric patients. The price of package – for those formulations not patented their reference price is considered – net of mandatory discounts (AIFA 2006a; 2006b) was obtained from the Italian Medicines Agency reimbursement lists (AIFA 2017); while the dosage from the summary of product characteristics of each active substance.

Table 8. Length and cost per cycle of antibiotic therapy

Antibiotics	Cycle	Cost per cycle
Erythromycin	10 days	€ 12.37
Clarithromycin	14 days	€ 15.89
Azithromycin	3 days	€ 6.70

3.2 Indirect costs

For what concerns indirect costs, the model estimates the social impact of pertussis in terms of loss of days of daily activities by one parent, if the person is less than 12 years old or by the person itself otherwise.

In the model, the age of the person who has to stay at home is estimated as the mean of the average age of fathers and mothers at birth (ISTAT 2015) plus the age of the infected person.

Concerning the number of days lost (De Serres *et al.* 2000) the model assumes that the distribution of those 12-17 years old also applies to younger (see Table 9 and 10). Finally, age-specific values for the loss of a day of daily activities were obtained from an Italian study estimating the societal costs of lost production in Italy (Pradelli and Ghetti 2017).

Table 9. Distribution of days lost – adolescents – scenario 1

# of days lost - <18 years	
Mean number of days lost	5
No days lost	22%
1-5 days	53%
6-10 days	17%
≥11 days	9%

Table 10. Distribution of days lost – adults – scenario 1

# of days lost - Adults	
Mean number of days lost	7
No days lost	33%
1-5 days	36%
6-10 days	13%
11-15 days	7%
≥16 days	10%

For what concerns the estimated number of days lost, in the model, there is also another, more conservative, scenario.

For outpatients, it is assumed a loss of a quarter of daily activities for every GP consultation by one parent, if the person is less than 12 years old or by the person itself otherwise. On the other hand, for inpatients, it is assumed a loss of a full day for day hospital and a loss equal to the length of hospital stay for those treated with ordinary hospitalization. Data on age-specific lengths of hospital stay, as reported in Table 11, were derived from an article examining demographics, case fatality rate, resource use and costs of hospital care related to pertussis in the US from 1996 to 1999 (O'Brien and Caro 2005).

Table 11. Age-specific length of hospitalization

Age-class >>>	<1	1-11	≥12
Mean length of hospital stay [days]	6	3.7	3.4
Median	4.5	3	3
Range (min, max)	1	1	1
	110	10	11

4. Probabilistic sensitivity analysis

A probabilistic sensitivity analysis (PSA) was performed in order to account for uncertainty in the parameters.

While the micro-simulation takes into account the variability in the population, the PSA allows considering the uncertainty on key parameters and its effect on the estimated outcomes. This is obtained by a two-level Montecarlo simulation: the inner loop (10,000 iterations) is the patient-level simulation, which is averaged and repeated 1,000 times (outer loop) to perform PSA on key model parameters.

In the absence of reliable data on the uncertainty of model parameters, standard deviation was set at 10% of their mean value, and adequate distributions were attributed according to the type of data (i.e. gamma or weibull distributions for age, number of days lost and length of hospital stay and beta distributions for probabilities). For conjugate probabilities (i.e. the probabilities of developing pneumonia, seizures, encephalopathy, of being hospitalized without complications or of being an outpatient, etc.) Dirichlet distributions were used.

The parameters on which the PSA was conducted and their distributions are presented in Table 12.

Table 12. Parameters and distributions for PSA

Variable	Expected value	Distribution	Alfa (1) o Lambda (2)	Beta (1) o K (2)
Distr. cases < 1 year				
Pneumonia	12.06%	Dirichlet		
Seizures	1.67%			
Encephalopathy	0.50%			
Hospitalization w/o complications	81.17%			
Outpatients	4.60%			
Distr. cases 1-4 years				
Pneumonia	1.04%	Dirichlet		
Seizures	0.18%			
Encephalopathy	0.03%			
Hospitalization w/o complications	1.84%			
Outpatients	96.91%			
Distr. cases 5-9 years				
Pneumonia	0.54%	Dirichlet		
Seizures	0.10%			
Encephalopathy	0.04%			
Hospitalization w/o complications	1.82%			
Outpatients	97.50%			
Distr. cases 10-19 years				
Pneumonia	0.22%	Dirichlet		
Seizures	0.03%			
Encephalopathy	0.01%			
Hospitalization	1.84%			

w/o complications				
Outpatients	97.91%			
Distr. cases ≥ 20 years				
Pneumonia	0.03%	Dirichlet		
Seizures	0.01%			
Encephalopathy	0.01%			
Hospitalization w/o complications	3.03%			
Outpatients	96.92%			
Age distribution of cases				
< 1	3.02%	Dirichlet		
1-4	15.67%			
5-9	12.83%			
10-14	27.31%			
15-19	6.75%			
20-29	5.63%			
≥30	28.80%			
Distr. # GP consultations < 10 years				
1	2.80%	Dirichlet		
2	38.80%			
3	37.80%			
4	18.80%			
5	1.80%			
Distr. # GP consultations ≥ 10 years				
1	2.20%	Dirichlet		
2	42.20%			
3	41.20%			
4	13.20%			
5	1.20%			
Distr. antibiotics				
Erythromycin	55.00%	Dirichlet		
Clarithromycin	36.00%			
Azithromycin	9.00%			
% treated with antibiotics				
< 1	73.00%	Beta	26.27	9.72
1-9	69.00%	Beta	30.31	13.62
10-19	45.00%	Beta	54.55	66.67
20-44	61.00%	Beta	38.39	24.54
≥45	56.00%	Beta	43.44	34.13
Day Hospital rate ≥ 1	26.52%	Beta	73.21	202.84
Age of parent at birth	33.50	Gamma	100	0.335
# days lost < 18 years	5.00	Gamma	0.69	5.41

# days lost ≥ 18 years	7.00	Gamma	0.37	14.48
Length of hospital stay [days]				
< 1	6	Weibull	0.37	0.62
1-11	3.7	Weibull	0.07	1.97
≥ 12	3.4	Weibull	0.10	1.78

Note: 1 – for beta and gamma distributions, 2 – for weibull distributions

In addition, in order to account for uncertainty in the parameters of the gamma and weibull distributions, these were further sampled from bivariate normal distributions, with mean equal to the value of parameters in the base case (as reported in the previous table) and assuming a standard deviation of 10% of their mean value and a correlation coefficient of -0.5. This last assumption was dictated by the fact that, by definition, the parameters of the gamma distribution are inversely proportional.

5. Results

The model simulates a population of 10,000 pertussis patients whose age distribution (see Table 12) was estimated on the data on notified pertussis cases in Italy from 2003 (i.e. the first year after vaccination was offered free of charge by all Italian regions) to 2015, obtained from European Centre for Disease Prevention and Control (ECDC) and corrected in order to account for under-reporting (see section 2.1). The results of the probabilistic sensitivity analysis are reported in Table 13.

Table 13. Age-specific estimates of direct medical costs and indirect costs per pertussis case

Age-specific estimates of direct medical costs and indirect costs per pertussis case			
	Medical costs	**Indirect costs – Sc. 1**	**Indirect costs – Sc. 2**
< 1 year	€ 1,488 (48.36)	€ 382 (67.01)	€ 844 (196.87)
1-4 years	€ 104 (10.80)	€ 340 (48.84)	€ 73 (3.48)
5-9 years	€ 94 (10.29)	€ 360 (41.63)	€ 76 (2.23)
10-14 years	€ 77 (5.60)	€ 155 (14.48)	€ 30 (0.98)
15-19 years	€ 87 (11.97)	€ 101 (11.50)	€ 17 (0.59)
20-29 years	€ 102 (13.31)	€ 299 (35.76)	€ 40 (1.80)
≥ 30 years	€ 93 (6.91)	€ 439 (44.52)	€ 59 (1.32)
Total	**€ 135 (24.78)**	**€ 302 (35.16)**	**€ 75 (13.05)**

Note: Data presented as mean (sd)

Direct medical costs are those costs that focus exclusively on health care resource utilization. For pertussis, this would include such things as hospitalizations, emergency room/physician visits, laboratory tests, and medications. Typically the direct costs of pertussis are higher in infants, for whom the disease burden is considerably greater and hospitalization is more common (Caro *et al.* 2005).

The average medical costs associated with pertussis range from € 77 to € 1,488 per patient and, as expected, these are particularly relevant for infants, where about 95% of cases are hospitalized and about 12% develop pneumonia as a complication.

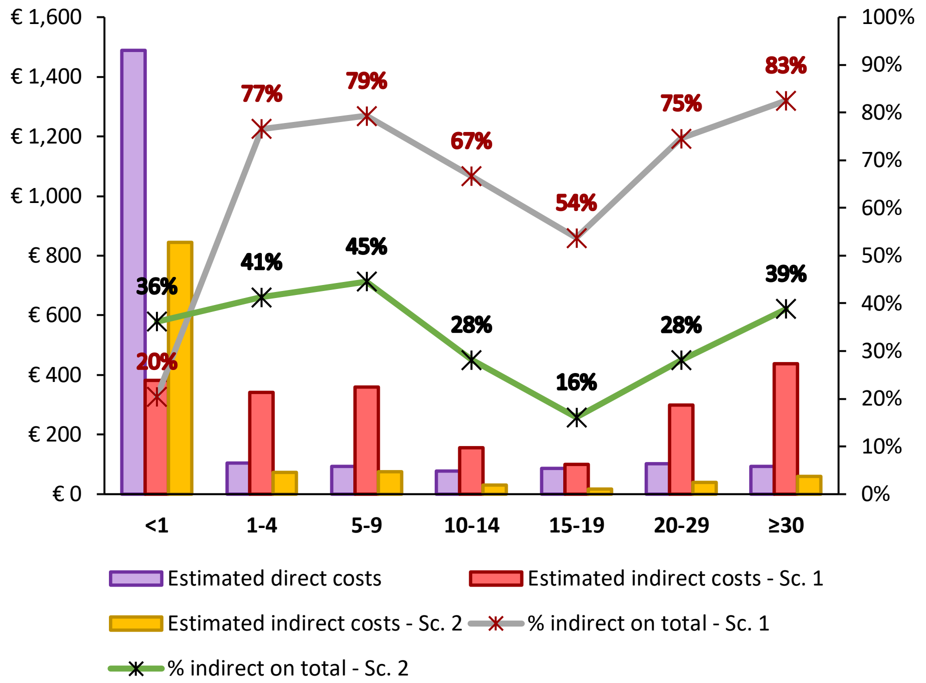
The indirect costs of pertussis are those incurred as a consequence of the illness, even though no direct expenditure has occurred. These include costs associated with time diverted from normal activities (e.g. as a consequence of visits to the general practitioner) and reduced work productivity, both of which may be caused by either individual illness or illness in a family member. Indirect costs can be expected to be relatively higher in adult cases, in whom illness is most directly linked to time lost from paid work activities, but can also be high for cases in infants and young children, where working parents are required to stay at home to care for their children (Lee and Pichichero 2000).

The results of the model are in line with what expected: indirect costs (scenario 1) are much higher for infants, young children, and adults than for adolescents.

For what concerns indirect costs, the model simulates two different scenarios. The second scenario has more conservative assumptions and, consequently, indirect costs are in general much lower with respect to the other scenario, with the exception of infants. This is due to the fact that the mean length of hospital stay for these patients may be also very long (up to 110 days). Therefore, valuing the number of days spent in the hospital, the estimated burden in this age-class is much higher than in the other scenario.

A graphical representation of the estimates obtained is given in Figure 2. For infants, medical costs are much higher than indirect ones, while the opposite holds true for all the other age-classes, considering the indirect costs of the first scenario. In the other scenario, indirect costs are always lower than direct costs, for every age-group.

Figure 2. Economic results of the simulation



5.1 Estimated burden in Italy, 2006-2015

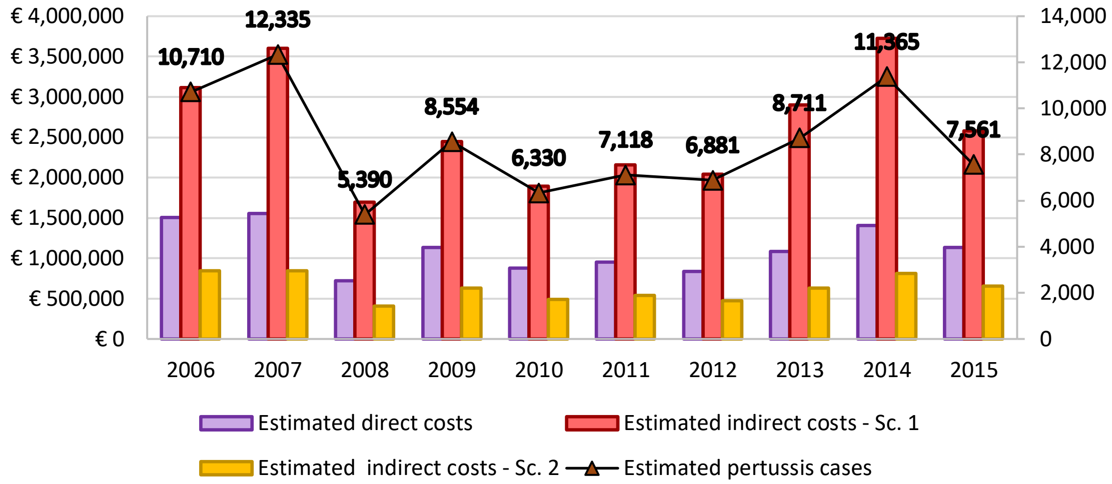
Data on the incidence of pertussis in Italy in the last decade (2006-2015) (ECDC) were corrected in order to account for under-reporting (see section 2.1).

With these data and the age-specific costs per pertussis case presented above (see Table 13), the burden the Society had to bear, in the last decade, because of pertussis amounts to, considering only direct medical costs, about € 11 million, corresponding to € 1.11 million per year. This result shows that annual costs for pertussis in Italy are still considerable and do not substantially deviate from those estimated for the Netherlands (approximately € 1.77 million) (de Greeff *et al.* 2009).

Pertussis in infants was responsible for 33% of these costs. The 1-4-year olds, 5-9-years olds, 10-14-year olds, 15-19-year olds, 20-29-year olds and ≥30-year olds accounted in for respectively 12%, 8%, 15%, 5%, 4% and 23% of the total direct medical costs.

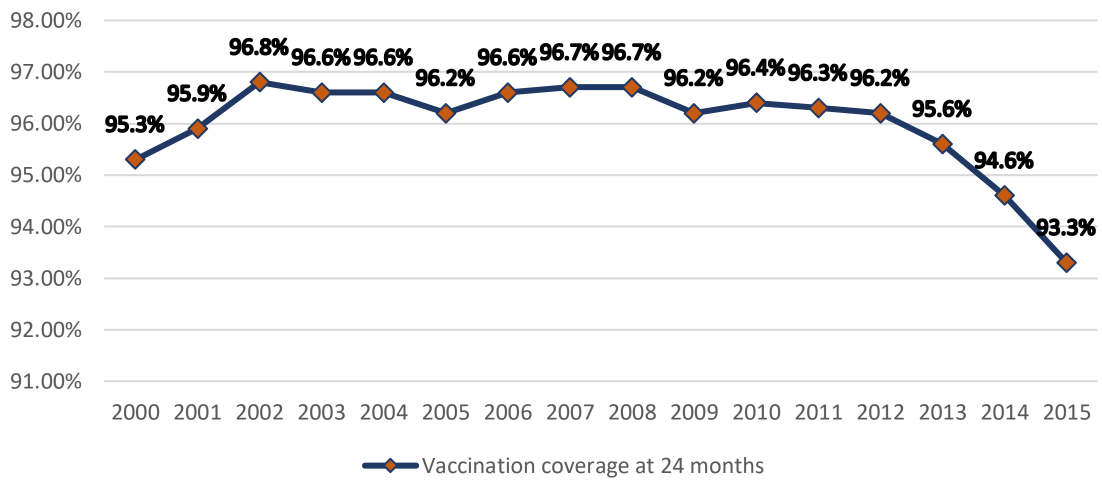
Considering also indirect costs, the mean annual burden rises to € 3.74 million (70% of which is due to indirect costs) in the first scenario and to € 1.75 million (where indirect costs amount to 36% of the total costs) in the more conservative one.

Figure 3. Estimated economic burden and estimated number of pertussis cases in Italy, 2006-2015



In order to account also for vaccination coverage, data were obtained from Istituto Superiore di Sanità (ISS). Figure 4 reports the level of vaccination coverage for pertussis at 24 months attained in Italy since 2000. Data do not show a clear trend: despite the vaccination coverage is almost constant since 2000, the burden has dropped in 2008 and peaked again in 2014, following a decrease in the vaccination coverage started between 2012 and 2013.

Figure 4. Pertussis vaccination coverage at 24 months in Italy, 2000-2015



In order to compare the annual burden of pertussis with that of other vaccine-preventable pediatric diseases, we took data from the REVEAL (Rotavirus Gastroenteritis Epidemiology and Viral Types in Europe Accounting for Losses in Public Health and Society) study; a prospective, multicenter, observational study of acute gastroenteritis in children <5 years of age in selected areas of 7 European countries (Belgium, France, Germany, Italy, Spain, Sweden, and the United Kingdom) that collected and analyzed data for each country for a 12-month period (1 October 2004 - 30 September 2005). For Italy, Padua was the selected study area (Van Damme *et al.* 2007; Giaquinto *et al.* 2007a; 2007b).

The estimated cost per episode of confirmed rotavirus gastroenteritis amounts to € 691 (considering both direct and indirect costs) (Giaquinto *et al.* 2007a), which is in line with the cost per pertussis episode estimated by the model for a person <5 years of age: € 690 (scenario 1). Notwithstanding this, in Italy, the estimated yearly burden of pertussis amounts to € 3.74 million while that of rotavirus to 73.55 million (Giaquinto *et al.* 2007b).

6. Discussion

From the findings of a comprehensive review of the literature (Caro *et al.* 2005), performed by the Global Pertussis Initiative to uncover data that quantify the economic burden of pertussis, it emerged that the cost of pertussis is highly variable. One important aspect is whether pertussis-related complications, such as pneumonia and encephalopathy, develop. According to published German estimates (Tormans *et al.* 1998), an uncomplicated case of pertussis is estimated to produce direct costs of € 210, whereas a case requiring hospitalization will be considerably more expensive, incurring an average direct cost of € 1700. Pertussis-related pneumonia has been estimated to increase the direct cost to € 3940, and cases leading to encephalopathy were estimated at € 5170.

It also emerged that estimates of direct costs vary according to age, being highest in infants. A US study estimated the direct costs of pertussis at US$ 2822 for infants (0-23 months), US$ 308 for children (2-12 years), US$ 254 for adolescents (13-18 years) and US$ 181 for adults (19 years of age or older) (Lee and Pichichero 2000).

In addition, a large-scale US study focusing on the societal costs (medical and nonmedical, excluding antibiotics to treat contacts) of pertussis in adolescents and adults found that mean direct medical costs in a cohort of 1679 adolescents (10-17 years old) and 936 adults (18 years old or older) were $242 and $326, respectively (Lee *et al.* 2004).

In this model, estimates of direct costs are much lower. Indeed, an uncomplicated case of pertussis is estimated to produce direct costs of € 58, whereas a case requiring hospitalization will be producing an average direct cost of € 1,471. Pertussis-related pneumonia has been estimated to increase the direct cost to € 2,015, and cases leading to encephalopathy were estimated at € 2,989. Age-specific estimates of direct costs are reported in Table 13.

Although charges for medical consumption differ across countries and exchange rates may fluctuate, hampering direct comparison of costs, the costs per case also depend on the estimated level of underreporting and the direct medical costs considered. For instance, the presented model includes neither specialist consultation, nor laboratory tests, nor cough medicine, nor vaccination costs.

Published data relating to the indirect costs of pertussis are less available than those for direct costs, but those available strongly suggest that, from a societal perspective, the indirect costs of pertussis and its management are substantial, particularly for working parents and working adults with pertussis. Indeed the economic consequences of pertussis in terms of reduced work productivity and absence from work likely match or exceed direct health care costs (Caro *et al.* 2005).
For instance, the estimate that, in the period 2006-2015 in Italy, indirect costs accounted for 70% of the total burden for pertussis is in line with the finding that the overall indirect costs associated with 87 cases of pertussis totaled US$ 107,025, equaling 73% of the total costs (Lee and Pichichero 2000).
For what concerns vaccination coverage, with the introduction in the Italian market of whole-cell vaccine, pertussis incidence showed a decreasing trend, although immunization coverage was still very low. When acellular vaccines replaced whole-cell ones, vaccine coverage increased dramatically and, within few years, incidence felt to so low figures that since 2002 epidemic cycles have been less clearly identifiable (Gonfiantini *et al.* 2014).
Bearing in mind that the estimated cost per episode for a person <5 years of age is almost identical for rotavirus and pertussis, the almost twenty-times higher annual Societal burden for rotavirus can be ascribed to the fact that this disease has a very low vaccination coverage. Vaccination is not mandatory yet and immunization policies still differ among regions.

This study has several limitations and I acknowledge there are still uncertainties around the proposed estimates of disease burden and assumptions on health care utilization. Concerning the number of pertussis cases, as said, passive surveillance systems, based on notifications, suffer from under-notification; I tried to overcome this issue correcting the available data. Regarding resources consumption, I neither collected data prospectively on the actual consumption of healthcare resources nor I collected information on the indirect costs, so that I had to rely on available data. For what concerns complications, I did not have access to Italian data, therefore I had to rely on US ones. The same holds true for data regarding the number of GP consultations, the percentage of outpatients cases treated with antibiotics, the number of days lost and the length of hospital stays.

7. CONCLUSION

The analysis presented in this article estimated age-specific costs associated with pertussis in Italy and the burden of pertussis the Society had to bear in the last decade (2006-2015). The results show that the average medical costs associated with pertussis range from € 77 to € 1,488 per patient and that these are particularly relevant for infants. Furthermore, the annual direct medical costs for pertussis in Italy are still considerable (approximately € 1.11 million) and do not substantially deviate from those estimated for the Netherlands (approximately € 1.77 million) (de Greeff *et al.* 2009).

The high level of vaccination coverage attained since 2000 has played a key role in diminishing both the incidence of the disease and the related burden for Society. Indeed, even though the estimated cost per episode of confirmed rotavirus gastroenteritis, amounting to € 691 (Giaquinto *et al.* 2007a), is in line with that of a pertussis episode, estimated in € 690, the annual societal burden associated with rotavirus gastroenteritis is almost twenty-times higher: € 3.74 million against 73.55 million (Giaquinto *et al.* 2007b).

The results highlighted the importance of indirect costs, representing 70% of total costs. Different studies suggest that the indirect economic burden of pertussis is substantial because of the economic expenditures required to care for ill family members, the prolonged recovery from illness and the consequential time lost from work.

Although the available studies provide some estimates of medical resource use and data on work time lost due to illness, a complete picture of the economic burden of pertussis is lacking. Major obstacles to the accurate assessment of the costs of pertussis are that the true incidence of pertussis is subject to considerable uncertainty, the severity and cost consequences of unreported and undiagnosed cases are unknown, and costs by the severity of symptoms have not been fully investigated.

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VALENTINA TONEI

MOTHER'S HEALTH AFTER CHILDBIRTH: DOES DELIVERY METHOD MATTERS?[1]

Abstract. The dramatic increase in the utilization of caesarean section has raised concerns on its impact on public expenditure and health. While the financial costs associated with this surgical procedure are well recognized, less is known on the intangible health costs borne by mothers and their families. We contribute to the debate by investigating the effect of unplanned caesarean deliveries on mothers' mental health in the first nine months after the delivery. Differently from previous studies, we account for the unobserved heterogeneity due to the fact that mothers who give birth through an unplanned caesarean delivery may be different than mothers who give birth with a natural delivery. Identification is achieved exploiting exogenous variation in the position of the baby in the womb at the time of delivery while controlling for hospital unobserved factors. We find that mothers having an unplanned caesarean section are at higher risk of developing postnatal depression and this result is robust to alternative specifications.
Keywords. Caesarean Section, Instrumental Variables, Maternal Health, Millennium Cohort Study, Postnatal Depression

1. INTRODUCTION

Over the past few decades a dramatic growth in the caesarean section (CS) rate has been recorded in many developed countries, regardless of the type of healthcare system (and relative incentives for physicians) and women's health needs (Bragg *et al.*, 2010; Gibbons *et al.* 2010). In England, for example, the overall rate was about nine

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per cent in the 1980s, while nowadays more than one-fourth of women gives birth through caesarean delivery (Health and Social Care Information Centre 2009, 2012). Similar patterns have been experienced by other OECD countries (OECD 2015), raising questions regarding the economic implications of alternative delivery methods.

Concerns about the increase in CS utilisation are justified by the high economic and health costs associated with this procedure (Koechlin *et al.* 2010). Indeed, while it is undeniable that caesarean deliveries have life-saving effects for mothers and children, especially for those who have concurrent health conditions (Gholitabar *et al.* 2011), it is also recognised that this procedure is very expensive, being the cost of a caesarean delivery between 66 and 88 per cent higher than the cost of a natural delivery (Gruber and Owings 1996; Petrou *et al.* 2002; Epstein and Nicholson 2009). Besides the financial impact, the World Health Organization (WHO) has highlighted the association of this procedure with short- and long-term health risks for the mother (WHO 2015).

This paper aims to contribute to the debate by analysing the causal impact of unplanned caesarean sections on mothers' mental health after childbirth. While the negative effects for mothers' physical health in terms of longer postpartum recovery and prolonged pain are well-known, less evidence is available on the effect on their psychological well-being. Mental health issues in general, and postnatal depression in particular, have been found to largely impact mother's life, being associated with a deterioration of her physical well-being and the relationship with her partner. Previous studies have also shown a strong link between maternal mental health and child development (Minkovitz *et al.* 2005; Propper *et al.* 2007; Kiernan and Mensah 2009; Coneus and Spiess 2012), rates of infections, hospital admissions and completion of recommended schedules of immunization for children (WHO 2015), child health (Perry 2008), long term child educational, labour market and criminal outcomes (Johnston *et al.* 2013).

For identification, we focus on unplanned deliveries. This is justified by two reasons. First, elective and unplanned caesarean deliveries may have different impacts on mother's mental health. Indeed, unplanned caesareans are unexpected, usually mentally and physically stressful, and associated with a loss of control and unmatched expectations. On the contrary, planned caesareans are scheduled in advance, allowing for the possibility for women to adjust (at least partially) their expectations for this event. The second motivation concerns a limitation of the data employed in the analysis. We cannot distinguish among planned caesarean deliveries, those that have been scheduled because of mothers' or babies' health needs and those requested by mothers for other reasons (the so-called *caesarean delivery on request*). Distinguishing between the two cases may be important because a different psychological impact of this procedure is expected depending on the reason why it has been implemented.

Medical studies investigating the relationship between the delivery method and maternal mental health find that caesarean deliveries are expected to carry higher risks for mothers' mental health compared to natural deliveries. Indeed, women who have a caesarean delivery are more likely to suffer from physical pain after childbirth and have longer and more difficult postnatal recovery, both conditions that also affect their psychological well-being. Additionally, caesarean deliveries may have a direct effect on mothers' mental health due to separation of mothers and their babies in the instants after the delivery. However, previous literature investigating this topic has not reached a unanimous consensus on whether having a caesarean delivery increases the risk of postnatal depression. This may depend on the limitations that characterise some of these studies, such as the small sample usually restricted to a particular geographic location or a population cohort, which does not allow to generalise the results to the entire population (Fisher *et al.* 1997; Koo *et al.* 2003). Failure to distinguish between elective (i.e. planned) and unplanned caesarean deliveries might also represent an issue, given that people tend to adjust better to traumatic events when they can predict or prepare for them (Clement 2001). Additionally, the variability in the source of information on mothers' mental health (e.g. medical visits, self-completion questionnaires) and in the length of the postnatal period during which mothers develop depression (from a few weeks to one year after childbirth) can contribute to explain such variability (Robertson *et al.* 2004; Patel *et al.* 2005; Carter *et al.* 2006).

These studies are in general characterised by the common assumption that women who have an unplanned caesarean do not differ from those who give birth naturally except through observable characteristics for which we can control. However, because of data limitations and the multiplicity of factors which can have an impact on both the delivery method and mothers' mental health, it is very unlikely to be the case. As a result, the estimates reported in these studies may be (downward) biased.

This paper builds on the previous research by addressing some of these issues. It investigates the effect of caesarean deliveries on the risk of postnatal depression by employing a nationally-representative sample of British children (and mothers) obtained from the first sweep of the UK Millennium Cohort Study. We use a medium-term measure of maternal postnatal depression, which captures a period of sadness in the first nine months after childbirth. More importantly, this study represents the first attempt to identify the *causal* link between unplanned caesarean deliveries and mothers' mental health, by accounting for unobserved differences between mothers who give birth through different delivery methods (endogeneity). While other papers in the economic literature have dealt with this issue when analysing the causal effect of the delivery method on mothers' (Halla *et al.* 2016) and children's outcomes (Jensen and Wüst 2015; Costa-Ramon *et al.* 2018), none of them have looked to psychological consequences for mothers and the econometric methods they have employed are different from those employed in this study.

We identify two main sources of endogeneity. One is due to the unobserved hospital characteristics that can both affect the choice of the delivery method as well as the risk of developing postnatal depression. For example, the level of resources available in a hospital in terms of staff and operating rooms may affect the level and standard of care. More specifically, in a hospital with a low nurse-to-patient ratio, women may receive less attention both during the labor and after the delivery. This may translate into (a) more compilations during labor, and therefore, into a higher risk of having an unplanned caesarean and (b) less psychological support after the delivery. At the same time, because the caesarean section is a surgical procedure, which requires to be performed in operating rooms by surgeons, medical staff may decide to opt for this delivery method only in extreme cases and they may prefer to perform a natural (or instrumental) delivery whenever possible.

The second potential source of endogeneity is related to the fact that mothers who have an unplanned caesarean section might be systematically different from mothers who give birth naturally, in terms of their own health and of the health of their babies. While we can control for some characteristics (e.g. maternal age and child's health at birth) which the literature has identified as driving the risk of having an unplanned caesarean delivery, there may be other factors we cannot observe due to data limitations, such as mother's mental and physical health before and during the pregnancy. Also, we do not have detailed information on what occurred during the delivery and whether pain relief were used. However, epidural anesthesia has been found to increase the risk of unplanned caesarean delivery, and it is also associated to a reduction of pain during labour, which turns into a lower psychological impact from this event. As a result, failing to account for this factor can affect the estimation results.

In order to overcome these problems, we adopt an instrumental variable approach combined with hospital fixed effects, the latter to control for time-invariant characteristics at hospital level. As main source of exogenous variation we exploit, for the first time in the literature, the position of the baby in the womb at the time of delivery. It has been shown that, conditional on mother's observable characteristics, the probability of having babies in abnormal position (i.e. with shoulders or feet first) is random and mothers cannot affect it with their behavior. We control for mothers' health conditions which may be (weakly) associated to the position of the baby and we argue that the variation left is exogenous. As an additional source of variation we use information on whether the mother suffered by pre-eclampsia during pregnancy. This is a health condition which may affect mothers after the 34th week of gestation and it is associated with some chronic conditions, such as diabetes and kidney diseases. While we control for these health problems, we cannot rule out completely that pre-eclampsia is uncorrelated with mother mental health after childbirth. As a result, we decide to rely on baby position in the womb and to use information on pre-eclampsia as a sensitivity only.

Our results show that having an unplanned caesarean delivery increases the risk of postnatal depression. Without accounting for endogeneity, we find that a woman who gives birth through this procedure is 3 percentage points more likely to experience postnatal depression. The sign and statistical significance of this effect is confirmed by IV estimates, even if in this case marginal effects are larger (15.3 percentage points). Results are robust to a number of specifications.

The rest of the paper is organized as follows. Section 2 presents our empirical strategies while data are described in Section 3. Section 4 shows the main results and discusses the validity of the instruments employed in the analysis. Section 5 provides some sensitivity checks and Section 6 concludes.

2. EMPIRICAL STRATEGIES

We study the effect of the mode of delivery on mother's mental health after childbirth. In particular the empirical model explaining the risk of postnatal depression is specified following the standard health capital models as theorized by

Grossman (1972) and firstly estimated by Rosenzweig and Schultz (1983). We adapt such framework by defining a maternal mental health production function that includes medical, as well as socio-economic factors associated with the risk of postnatal depression.

$$PD_m = f(CS_m, \mathbf{SES_m}, \mathbf{X_m}, \mathbf{H_m^H}, \mathbf{H_m^C}) \quad (1)$$

The outcome variable is PD_m, a binary variable denoting whether mother *m* suffered from postnatal depression in the first nine months after childbirth. CS_m represents the causal variable of interest and indicates whether the mother gave birth through an unplanned caesarean delivery (as compared to a natural delivery). SES_m is a vector including all socio-economic variables that may be related to mother's mental health. X_m includes information on pregnancy; while $H_m{}^M$ and $H_m{}^B$ refer to mother's and child's health respectively.

In order to overcome the fact that the delivery method is not randomly assigned we combine two econometric approaches: hospital fixed effect models and instrumental variables.

The inclusion of hospital fixed effects into the model allows us to control for hospital characteristics, such as internal organisation, resources availability (e.g. medical staff and operating rooms) and quality of care which might affect both the probability of giving birth through an unplanned CS and the risk of developing postnatal depression. However, there may still be mothers-specific unobservable characteristics correlated with the mode of delivery and their mental health (e.g. mother's health status during pregnancy and delivery experience) which would bias the results. We account for this issue by adopting an instrumental variable approach which exploits two sources of variation: (a) the position of the baby in the womb before the delivery and (b) mother's health status during the pregnancy, namely whether she suffered from pre-eclampsia.

We define a binary variable, Pos_m equal to one if the baby presents feet or shoulders first, head at the back or other abnormal positions at birth, a situation called *breech position*. Full breech position at term means that the baby has not turned head down in the womb by week 37 of the pregnancy. Among babies at term, breech position is present in three to four per cent of all births (Royal College of Obstetricians and Gynaecologists 2006).

This approach requires two conditions to be met. The first is that having a baby in a breech position at the time of delivery is uncorrelated with unobserved characteristics of the mother (and her pregnancy). Tharin *et al.* (2011) argue that breech babies can be considered as a good random subgroup of all babies since there is no clear evidence of maternal or baby's characteristics that can predict the probability of breech position. This is also supported by Jensen and Wüst (2015), who show that breech and non-breech mothers are similar in a range of observable characteristics such as level of education, pregnancy conditions unrelated to breech (e.g. pre-eclampsia and diabetes). Also the Royal College of Obstetricians and Gynaecologists (2006) states that, while persistent breech presentation may be associated with biological factors (such as amniotic fluid volume, the placental localisation and the uterus), it may be due to chance as well. This result is also generally confirmed by the medical literature, although there is some evidence of a weak association between the position of the baby and some predictors of postnatal depression, such as mother's age, parity, health behaviors and birth weight (Rayl *et al.* 1996; Fruscalzo *et al.* 2014). When using our sample to compare characteristics of mothers with breech and not-breech babies, we also find some differences in maternal characteristics (parity, ethnicity, socio-economic status) and in child health at birth. The latter result is not surprising since breech babies are more likely to be underweight and preterm (Cammu *et al.* 2014). However, when we regress baby's position on the full set of covariates (including unplanned CS), we find that apart from the mode of delivery, only parity, baby's health at birth and mother's ethnicity are significantly associated with the position of the baby. Nonetheless, to make sure the instrument is actually exogenous, we control for these factors in the model. In Section 4 we provide further evidence of the validity of the instrument computing an over-identification test after estimating linear IV models.

The second condition to satisfy is the relevance of the instrument, i.e. POS_m must be a strong predictor of the mode of delivery. The National Institute of Clinical Excellence (NICE) guidelines encourage resorting to a caesarean delivery if a breech position occurs at the end of the gestational period to reduce the risk of perinatal mortality and neonatal morbidity (Gholitabar *et al.* 2011). Similar guidelines have been issued in other countries, especially after the publication of the results from the *Term Breech Trial*, the largest randomized control trial evaluating the adequate mode of delivery for breech babies that has shown the superiority of planned caesarean delivery with respect to a planned natural delivery (Hannah *et al.* 2002). As a result, a large proportion of breech babies are born through a caesarean delivery in the United

Kingdom every year (Bragg *et al.* 2010), and similar rates are observed in other countries, e.g. USA (Lee *et al.* 2008), Sweden (Alexandersson *et al.* 2005), Denmark (Jensen and Wüst 2015) and the Netherlands (Rietberg *et al.* 2005). Given the strong compliance with guidelines, we expect a strong and positive correlation.

Our additional instrument exploits variation in the probability of experiencing pre-eclampsia during the pregnancy. This health condition is mainly characterised by high blood pressure (hypertension) and it usually occurs after the 34th week of gestation. We define a binary variable, ECL_m, equal to one if the mother suffered from this condition during the pregnancy.

We expect this condition to be highly correlated with the mode of delivery as women with pre-eclampsia are at higher risk of stroke and heart attack, and since a natural delivery is a very stressful event with a strong impact on the woman's body, physicians tend to opt for a (unplanned) caesarean delivery to avoid such risk.

This approach also requires that unobserved characteristics do not affect jointly the probability of suffering from pre-eclampsia and the mode of the delivery. As in the case of breech position, women cannot directly affect their probability of experiencing this condition. However, differently from the previous case, there are some biological and behavioral factors that the medical literature has identified as strong predictors of this condition. These include parity, obesity, multiple deliveries and antenatal visits. Moreover, women with existing long-term medical problems, such as diabetes, kidney diseases or high blood pressure, are at higher risk of pre-eclampsia. While we can easily control for the first set of covariates, this is not completely possible for the others. Indeed, our data report only whether the woman has ever been affected by these long-term illnesses during her life, but no information is available on the onset of such conditions. As a result, even if the woman reports that she suffers (or has suffered) from such health conditions, we are not able to determine whether this happened during the pregnancy. While imperfect, we include these measures in the model to isolate the exogenous variation in mothers' probability of suffering from pre-eclampsia. However, because of the limitations discussed above, our preferred specification is the one exploiting variation in the position of the baby at the womb.

We use POS_m, in combination with ECL_m, as a robustness check in order to obtain an over-identified model and be able to test formally the validity of the instruments.

In terms of estimation methods, we implement an instrumental variable approach by adopting a two-stage least squares estimation using POS_m as the main instrument and including hospital fixed effects in the model. Then we re-estimate the model adding ECl_m to the vector of instruments to obtain an over-identified model, which allows testing the validity of the instrument using a Sargan test.

The main advantages of using linear IV methods concern the possibility to test the validity and relevance of the instruments and to interpret the coefficients in terms of marginal effects. However, as in the OLS case, this specification ignores the binary nature of the health outcome and the endogenous variable. To account for that, as a robustness check, we also estimate bivariate probit models (Heckman 1978), following the approach suggested by Nichols (2011).

3. Data

The data come from the UK Millennium Cohort Study (MCS), a multidisciplinary longitudinal data set on a cohort of children born between 2000 and 2002. For the purpose of this study, we use only the first sweep, which contains detailed information on circumstances of pregnancy and birth, as well as socio-economic background and health conditions of the family where children were born.

Our initial sample is characterised by 18,818 children, born from 18,552 women. We exclude women who had a multiple delivery (two or more babies) because they are more likely to have health complications after childbirth and their babies are systematically different in terms of (lower) birth weight, gestational age at birth and other birth characteristics in comparison to single-pregnancy babies. We also drop observations with missing or incomplete information on the variables included in the model. This lead us with a final sample of 14,221 women.

3.1 Variables

The outcome variable is a binary indicator that takes a value equal to one if the mother reports to have experienced a period of sadness lasting two weeks or more after childbirth. Previous literature focusing on mother's postnatal depression shows a high degree of heterogeneity in the definition of this condition. In particular, while there is a general agreement in the medical community on the symptoms that identify postnatal depression (e.g. low mood, loss of enjoyment and pleasure, anxiety), the length of the period after delivery and the time of onset that should be taken into account are less clear. In this paper, we follow the definition of postnatal depression suggested by Mcintosh (1993), by considering postnatal depression as the experience of a depressed mood for a period of at least two weeks at some stage during the first nine months after delivery. Compared to other measures, this can be considered a medium- to long-term indicator of maternal postnatal depression. MCS does not provide details on the severity of this condition, therefore mothers reporting symptoms of postpartum depression may be affected by this condition differently.

A potential limit of this measure is that it is built using self-reported information. The general concern about this type of health outcomes is that it can measure individuals' health with error, being affected by unintentional (e.g. recall bias) and intentional bias (stigma associated with mental disorders may lead mothers to under-report mental illnesses). Nonetheless, more 'objective' measures, such as postnatal depression diagnosed by doctors, are not necessarily more appropriate. A report from the Royal College of Obstetricians and Gynaecologists shows that about half of the mothers who experienced mental health problems were not referred to services or offered any further information about where to go for support. Additionally, the probability of being diagnosed with depression depends on the frequency of the contacts with her GP. If the woman tends to not attend GP visits, there exists a risk of underestimating the incidence of this condition.

We classify mode of delivery in three mutually exclusive groups: natural, elective caesarean and unplanned caesarean delivery. Natural deliveries are defined as those that can be classified as medical procedures according to the Healthcare Resource Groups (HRG) system (therefore, this category also includes instrumented deliveries), while caesarean sections are distinguished in elective and unplanned, the latter usually associated with unexpected complications at the time of delivery. Since the analysis focuses on the effect of unplanned caesarean deliveries compared to natural births, we drop from the sample elective caesareans and we define a binary variable, CS_m, taking a value equal to one if the woman had an unplanned caesarean delivery and zero in case of a natural delivery.

There exists a variety of factors that may be associated with both the probability of experiencing postnatal depression and giving birth through unplanned caesarean delivery. We classify them in the following categories: (i) socio-economic factors, (ii) pregnancy-related attitude, (iii) maternal health, and (iv) child health at birth.

Among the socio-economic factors, mother's age is one of the most important. Teenagers are, on average, more likely to experience postnatal depression (Brown 1996;

Deal and Holt 1998; Robertson *et al.* 2004). A dummy defining mother's ethnicity is included to account for the heterogeneous composition of the UK population. Ethnic minorities are less likely to be affected by anxiety and/or depression, probably because of their propensity to under-report health problems (Fiscella *et al.* 2002; Harris *et al.*, 2005). Household income and mothers' level of education have been identified by the economic literature as strong predictors of psychological wellbeing. They are also positively associated with unplanned caesarean sections, even after controlling for standard covariates (Gresenz *et al.* 2001; Segre *et al.* 2007). We measure mother's highest level of education with a set of dummies that indicate her highest national vocational qualification (defined in three categories). Marital status is included as a proxy for perceived social support. A married woman is expected to have psychological support from her husband in taking care of the child, which, in turn, reduces her likelihood of becoming depressed (Stewart *et al.* 2003).

As covariates related to the pregnancy, we include an indicator that measures if the pregnancy was planned and a measure of parity (i.e. whether the cohort member is the first child of the woman).

Physical and mental health before pregnancy are strong predictors of postnatal depression. Unfortunately, being the MCS a child-focused dataset, it does not include any direct information on maternal health before the child was born. To proxy her physical health before pregnancy, we use information about the mother's smoking behaviour and her body mass index (including both a linear and a quadratic term to allow for non-linearities in the effect). Admission to hospitals and whether she had a paid job during pregnancy are included to control for mother's health during childbearing. Furthermore, we include dummies to measure whether she suffers/has suffered in the past from diabetes, gestational hypertension or kidney diseases.

According to the literature (e.g. McLennan *et al.* 2001), mothers of unhealthy children are negatively affected by their babies' health and, as a consequence, they are more likely to report a depression status. We account for this association by including binary indicators for underweight and overweight baby at birth and gestational age (in weeks) as measures of baby's health. However, these variables may be endogenous if affected by the mode of delivery. In Section 5 we discuss the results when we exclude these variables from the model.

Finally, our empirical approach exploits the exogenous variation in the position of the baby at birth and whether the mother suffered from pre-eclampsia during the pregnancy. The instruments we derive are two binary variables, *Pos* and *Ecl*. The former takes the value of one if the baby was in a breech position or in another abnormal position requiring a surgical intervention. Pre-eclampsia, hereafter called *Ecl*, measures whether the mother has suffered from this hypertension disorder during the pregnancy and/or at the time of delivery.

3.2 Descriptive statistics

Table 1 provides some descriptive statistics of the variables used in the analysis for the full sample and by treatment status (i.e. the delivery method). 34.5 per cent of the women in the sample have experienced a period of sadness lasting at least two weeks after delivery (*PD*=1). This percentage is above the documented prevalence rate for postnatal depression in the U.K., estimated to be around 15 per cent. However, this difference could be ascribed to variability in the definition of postnatal depression and in the length of the postnatal period considered. Additionally, previous literature has shown that higher rates of incidence are observed when employing non-clinical definition of depression (e.g. self-reported measures as the one employed in this study), compared to the case in which standard instruments for the detection of depression are used.

When looking at other predictors of CS, we find a significant difference in mothers' health status during pregnancy, with CS mothers more likely to experience poor health (in terms of hospitalisation during pregnancy, diabetes and hypertension). Additionally, 64.4 per cent of the women who had an unplanned CS had no previous pregnancies, while only 39.4 per cent of those who had a natural delivery have no other children. This may suggest that a mechanism through which unplanned CS negatively affects mothers' mental health is the lack of information (or wrong expectations) they may have about the delivery experience.

4. RESULTS

Equation (1) is initially estimated with Ordinary Least Squares (OLS), treating the mode of delivery as exogenous. This specification can be viewed as a descriptive regression which sheds light on whether the effect of unplanned caesarean delivery persists after controlling for other observed factors. Also, it provides a benchmark against which to compare the results from fixed effects and IV models. Table 2 presents OLS estimates obtained by adding gradually different sets of covariates in the regression equation.

In all the specifications, we find a positive and significant association between unplanned caesarean delivery and postnatal depression. When including measures of maternal health and pregnancy experience, the magnitude of the coefficient decreases (moving from 0.046 to 0.041), suggesting that part of the effect is due to the poorer level of health of mothers who gave birth through this procedure. The inclusion of measures of child health at birth further reduces the magnitude and such reduction is even larger. Overall, when controlling for the full set of regressors, we find that having an unplanned caesarean delivery increases the likelihood of postnatal depression by 3 percentage points and this effect is significant at the 5 per cent level.

As discussed above, OLS estimates may be biased if there are omitted variables at mother or hospital level. We first control for time-invariant hospital factors by including hospital fixed effect in the model. Results (Column 4, Table 2), are not statistically different from those obtained in the OLS specifications. This may be because unobservable hospital characteristics affecting both the delivery method and the risk of developing postnatal depression are not an issue in this context. Another explanation is that what matters to explain the risk of developing postnatal depression is the relationship of the mother with the nurse that follows her during and after the pregnancy, rather than factors at hospital level, such as quality of care and resources available. Under the second scenario, hospital FE would not solve the endogeneity issue, being this method only able to account for hospitals characteristics which do not vary over time and across mothers. However, the IV strategy can represent a solution, accounting for any source of endogeneity, provided that the instruments employed are valid.

4.1 First stage results

The first stage of linear IV models (Table 3) estimated using only POS_m as an exclusion restriction, show that the partial correlation between baby's position in the womb and unplanned CS is equal to 0.316 (0.320 when hospital fixed effects are included), and it is strongly statistically significant. This is in line with what obtained comparing the proportion of breech babies born through an unplanned caesarean section with those who are born naturally, which are 0.185 and 0.045 respectively. Further evidence of the instrument relevance is by the F-statistic testing for the significance of the excluding restriction equal to 324.1 and the test on the weak identification of the IV model (robust Kleibergen-Paap Wald rk F statistic).

Interestingly, the estimated effects of the covariates are in line with other studies (Halla *et al.* 2016). Maternal age and ethnicity are the most important socio-economic predictors with older women and women from minorities groups being at higher risk of having an unplanned caesarean delivery (see column 1 in Table A1 in the Appendix). Most of the measures of mother's health before and during the pregnancy have also predictive power. For example, hospitalisation during pregnancy, smoking before the pregnancy as well as having suffered from diabetes and hypertension also increase the risk of an unplanned caesarean delivery. A U-shape relationship between gestational age and the probability of unplanned caesarean is also found, with premature and post term babies more likely to born through this procedure. This is explained by the fact that a premature birth usually occurs because of the development of un expected health condition affecting either the mother or the baby. On the contrary, a post term delivery may come up to be an unplanned CS if an unexpected event occurred during the delivery, for example in case of delivery induction. Similar findings are obtained when we include hospital fixed effects to the model (results available from the author).

Adding ECl_m to the set of instruments (with or without hospital fixed effects) does not change the main conclusions. In terms of relevance of this instrument, we find that women suffering from pre-eclampsia have a higher risk of giving birth through an unplanned caesarean (coefficient equal to 0.08) and the effect is strongly statistically significant. Also in this case, the tests confirm the strong relevance of the instruments

4.2 Second stage results

Using POS_m as the only instrument we find that an unplanned caesarean delivery increases the risk of developing postnatal depression by 15.3 per cent (18.7 per cent) when we include (exclude) hospital fixed effects. The effect is not negligible, given the underlying risk of postnatal depression of 34.5 per cent. Same results are obtained when using both the instruments (columns 3 and 4 of Table 3).

When focusing on the coefficients associated with the covariates included in the model, we find that they behave as expected (see Table A2). For example, married women have a 3.9 percentage points lower chance of being affected by postnatal depression than unmarried women. Income also shows a strong negative association with the probability of experiencing postnatal depression. This result is consistent with the *family stress model*, as defined by Conger *et al.* (2000), which states how economic hardship and pressure negatively impact parents' mental health.

Among the health and pregnancy variables, we find that women who smoked before the pregnancy are at higher risk of developing postnatal depression compared to non-smoking mothers. Poor physical health, measured by hospitalization during pregnancy, strongly predicts postnatal depression, coherently with the literature that shows a strong association between physical and mental health (Canadian Mental Health Association 2008). In the same line we find that working during pregnancy is negatively associated with the probability of developing depression after childbirth (negative coefficient equal to 0.035). On the contrary, diabetes and hypertension are not significantly associated with postnatal depression (and the negative effect of having kidney problems is significant at 10 per cent only). This result goes in the opposite direction with respect to what is highlighted by the descriptive statistics, suggesting that once accounting for other concurrent health issues, these conditions are no longer relevant in explaining mother's mental health.

Results also show that having planned the pregnancy in advance is associated with a decrease in the probability of postnatal depression by 3.5 percentage points. Finally, as expected, poor baby's health (proxied by birth weight and gestational age) are strong predictors of maternal postnatal depression.

Overall, when comparing the coefficient associated with the delivery method in the OLS model with those found in linear IV models, it seems that failing to account for selection into the delivery method leads to an underestimation of the impact of this procedure on maternal mental health. But, is it really the case? An explanation for our results may be that the linear IV model produces upward biased results because the instrument employed in the analysis do not satisfy the exclusion validity assumption. However, as we discussed in Section 2, this is unlikely to be the case and the Sargan-Hansen test (reported in Table 3) further supports the validity of the instruments, so we rule out this hypothesis. Another explanation is that what we obtain when exploiting the exogenous variation in the position of the baby in the womb (and mother's pre-eclampsia) is a local average treatment effect (LATE), i.e. the average treatment effect for a defined *subgroup* of women who had an unplanned caesarean delivery as a consequence of the position in the womb of their baby at birth (the so-called compliers, see Angrist and Pischke 2008). That said, the effect is still interesting and policy relevant being the number of breech babies born every year not negligible (about 3-4 per cent of all deliveries).

5. Sensitivity checks and heterogeneity analysis

5.1 Model specification

Linear IV models may not be an appropriate model specification because of the binary nature of CS_m and PD_m. We account for this issue by estimating a bivariate probit model and comparing results from this model with those obtained using the linear IV specification. Because the coefficients of the bivariate probit model are not directly interpretable we report average marginal effects (AME) from the bivariate probit (reported in Table 4). When only breech position is excluded from the outcome equation, we find that having an unplanned delivery increases the probability of postnatal depression by 11.6 percentage points. Adding pre-eclampsia to the vector of exclusion restrictions does not significantly change the results (average marginal effect equal to 11.5 percentage points).

When interpreting the estimates in the two models, it is important to keep in mind the different source of identification that the two models exploit. The bivariate probit model depends, in addition to the exclusion restrictions, also on the functional form assumptions. The fact that in this case the magnitude of linear IV estimates and AME from the bivariate probit are similar suggests that the stricter assumption of joint normality of error terms in the bivariate probit is consistent with the data. More generally, these findings could be interpreted as evidence of the robustness of results, that do not depend on parametric assumptions.

5.1 Excluding baby health mesures

Gestational age at birth and birth weight may be endogenous as they are potentially affected by the mode of delivery. This may be the case if, for example, women with low birth weight babies and women at high risk of giving birth before their due date are more likely to receive intensive care before the birth of their child and are also more likely to end up with an unplanned caesarean. For this reason, as a sensitivity, we estimate linear IV models excluding these variables from the vector of covariates.

Results do not change significantly compared to the model with the full set of regressors. In particular, the instruments are still strongly associated with the endogenous variable, CS_m, suggesting that the relevance of the instruments is not affected by the exclusion of baby health measures. Similarly, the second stage estimates are not statistically different from those obtained including baby health measures. Finally, the Sargan Hansen test confirms the validity of the instruments (p-value equal to 0.9230). However, because breech position may be correlated with baby's health at birth, we prefer to include in the models these controls.

5.2 Heterogeneity

Women with more resources or knowledge may be more able to mitigate for negative events they experience. We test this assumption by splitting the sample into two groups of mothers, those with a university degree and those with a lower qualification (or any). When estimating the linear IV model for the two groups, we find no differences in the relationship between unplanned caesarean delivery and maternal mental health.

Similarly, we explore a mechanism which can explain the negative impact of unplanned caesarean deliveries on mothers' mental health, namely the difference between what women may expect or imagine to be the delivery and their actual experience. It may be that women who had previous pregnancies (regardless of the mode of delivery) have more information on the delivery process and, as a result, their expectations are more similar to the reality they experience. If this is the case, the psychological impact of having an unplanned caesarean would be smaller than the impact for women with no previous experience. However, when we test this hypothesis by splitting the sample by parity, we find no evidence of differences in the impact of unplanned caesarean deliveries. This may imply that having an unplanned caesarean delivery carries a lot of stress, regardless of the previous delivery experience.

6. Discussion and conclusions

This study contributes to the growing economic literature on the determinants and consequences of the increased utilisation of surgical procedures such as caesarean deliveries. It represents the first attempt to identify the causal effect of unplanned caesarean sections on mothers' mental health, looking at whether this procedure is associated with an increase in the mothers' risk of developing postnatal depression in the first nine months after childbirth.

Results show that unplanned caesarean deliveries carry significant psychological risks, with women who give birth through this procedure being more vulnerable to post-traumatic distress and depression (by 15 percentage points when estimated using linear IV models combined with hospital fixed effects). These findings are important for a number of reasons. First, caesarean deliveries have spread remarkably in recent years, becoming one of the most frequent surgical procedures, with 165,000 deliveries performed every year in England (among these, about 25,000 are unplanned caesarean deliveries). Second, depression can be a very severe condition limiting mothers' everyday life and their ability to take care of their children. Because mothers are usually the main childcare givers, poor mother's mental health is likely to negatively affect their baby's health and development. Additionally, postnatal depression is likely to become a chronic health condition, associated with high costs for families as well as for society (e.g. inability to work, see Schultz *et al.* 2013).

This paper contributes also to explain the effect of caesarean deliveries on women's subsequent fertility decisions. Halla *et al.* (2016) find that mothers who give birth through a caesarean section are less likely to have other children and they mention psychological problems after childbirth as a possible explanation. While we cannot argue that this is necessarily the case, this paper shows evidence of the existence of a negative psychological impact of unplanned caesarean deliveries, which can explain their findings.

A limitation of this study, which opens the door to future research in this area, relates to the data used to shed light on this phenomenon. A longitudinal administrative dataset with detailed information on mothers' previous pregnancy experiences and their health conditions before and after the pregnancy would allow to identify the causal effect of unplanned caesarean deliveries using alternative econometric approaches and fewer assumptions. In addition, it would allow the comparison of results from this study with those obtained using objective measures of mother's mental health. Reaching these goals would require a link of hospital records on maternity events to other data sources containing information on primary care visits, being depression usually diagnosed by general practitioners (at least in a first instance), and census data providing details on mothers' income, education, working condition and other socio-economic variables. However, such linkages are not currently available, at least for English data.

Another aspect left for the future is the extension of this study to elective caesarean deliveries. As discussed in the paper, from a theoretical point of view, we may expect elective caesarean deliveries to have a smaller impact compared to unplanned caesarean deliveries, being planned in advance and giving mothers the opportunity to adjust their expectations. However, they can still have a negative impact on their body, and as a consequence, make more difficult their post-partum recovery. Looking also to elective caesareans would provide a more complete picture of this phenomenon.

From a policy perspective, this study highlights the importance of accounting for the psychological costs of unplanned caesarean deliveries when evaluating the costs and benefits of this procedure. Failing to account for these factors would lead to inaccurate evaluations of this procedure and, as a consequence, to the implementation of inappropriate health policies (Drummond 2005). Additionally, it suggests the importance to provide appropriate services, such as professionally-based home visits and peer-based telephone support, to prevent the development of postnatal depression (Royal College of Obstetricians and Gynaecologists 2017).

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Appendix

Table 1. Mother and child characteristics by mode of delivery

	Full sample	*Natural delivery*	*Unplanned CS*	*Statistical difference*
Breech position	0.053	0.032	0.185	***
Pre-eclampsia	0.072	0.061	0.144	***
Mother's age	28.1	27.92	29.2	***
Married	0.577	0.572	0.606	**
Household income	15.37	15.03	17.48	***
No qualification	0.154	0.16	0.12	***
GCSE/O-level (or eq.)	0.394	0.398	0.367	**
A-level or higher (but no uni)	0.153	0.153	0.151	
University qualification	0.299	0.288	0.362	***
White	0.871	0.873	0.86	
Smoker (before pregnancy)	0.363	0.366	0.347	
BMI (before pregnancy)	23.59	23.43	24.57	***
Parity	0.443	0.411	0.644	***
Planned pregnancy	0.542	0.537	0.575	**
Hospitalisation (during pregnancy)	0.182	0.169	0.26	***
Employed (during pregnancy)	0.653	0.637	0.748	***
Diabetes	0.019	0.016	0.035	***
Hypertension	0.004	0.003	0.011	***
Kidney diseases	0.009	0.009	0.011	
Gestational age	39.69	39.793	39.043	***
Underweight baby	0.065	0.05	0.16	***
Overweight baby	0.121	0.114	0.161	***
Postpartum depression	0.345	0.341	0.37	*
No. Observation	14,221	12,249	1,972	

Notes: The table reports the mean of the variables employed in the analysis. Mother's age is measured in years, while baby's gestational age at birth is measured in weeks. Income is measured in thousands of GB pounds. All the other figures indicate percentages. Differences between groups are tested by means of 2-independent-sample t tests. *** p<0.01, ** p<0.05 and * p<0.10.

Table 2. Effect on Postnatal Depression estimated using Linear Probability Models with and without hospital fixed effects

	LPM (1)	*LPM* (2)	*LPM* (3)	*Hospital FE* (4)
Unplanned CS	0.046***	0.041***	0.030**	0.031***
	(0.011)	(0.011)	(0.012)	(0.012)
No. observations	14,221	14,221	14,221	14,221
Socio-economic variables	X	X	X	X
Pregnancy variables		X	X	X
Mother's health variables	X	X	X	X
Child's health variables			X	X

Notes: Robust standard errors clustered at hospital level in parentheses. Linear probability models estimated using OLS. *** p<0.01, ** p<0.05 and * p<0.10. Socio-economic variables include mother's age, ethnicity, marital status and household income. Pregnancy variables include parity and planned pregnancy. Mother's health include BMI, whether she was a smoker, employment status and hospitalisation during pregnancy, whether she ever suffered from diabetes, hypertension and kidney diseases. Child's health at birth include birth weight and gestational age.

Table 3. Effect on Postnatal Depression using Linear Instrumental Variable (IV) models

	IV: Breech position		*IV: Breech position & Pre-Eclampsia*	
	without FE	*with FE*	*without FE*	*with FE*
	(1)	*(2)*	*(3)*	*(4)*
First stage				
Breech position	0.316***	0.320***	0.317***	0.321***
	(0.018)	(0.017)	(0.018)	(0.017)
Pre-eclampsia			0.077***	0.081***
			(0.016)	(0.016)
Second stage				
Unplanned CS	0.186***	0.153***	0.187***	0.152***
	(0.050)	(0.049)	(0.049)	(0.049)
No. observations	14,221	14,189	14,221	14,189
Post-estimation tests				
F-test (first stage)	324.1	337.7	178.0	178.0
Weak identification test	324.1>16.38	337.7>16.38	178.0>19.93	178.0>19.93
Endogeneity test of CS[a] (p-value)	0.0017	0.0135	0.0014	0.0125
Sargan Hansen test (p-value)	-	-	0.9829	0.9631

Notes: Robust standard errors clustered at hospital level in parentheses. Linear IV models. *** p<0.01, ** p<0.05 and * p<0.10. All model specifications include information as in column (3) of Table 2. Column (2) and (4) include hospital fixed effects (FE). For the weak identification test we report Stock-Yogo critical values in case of one or two exclusion restrictions, allowing for 10 per cent of IV maximum distortion with respect to OLS. a: Under the null hypothesis of exogeneity, the statistic is distributed as a chi-squared with 1 degree of freedom.

Table 4: Effect on postnatal depression using linear IV models and bivariate probit models

	IV: Breech position		*IV: Breech position & Pre-eclampsia*	
	Linear IV model	*Bivariate model*	*Linear IV model*	*Bivariate model*
	(1)	(2)	(3)	(4)
Coefficients				
Unplanned CS	0.186***	0.328***	0.187***	0.326***
	(0.05)	(0.125)	(0.049)	(0.119)
Marginal effects				
Unplanned CS	0.186***	0.116***	0.187***	0.115***
	(0.005)	(0.044)	(0.049)	(0.042)
No. observations	14,221	14,221	14,221	14,221

Notes: Standard errors in parentheses. Linear IV models and bivariate probit models. *** p<0.01, ** p<0.05 and * p<0.10. All model specifications include information on socio-economic status, pregnancy, mother's health and baby's health as in column (3) of Table 2. Average marginal effects estimated with the Stata command margins. Delta-method standard errors.

APPENDIX

Table A1. Linear IV Models: first stage results

	IV: Breech position		IV: Breech position & Pre-eclampsia	
	without FE	*with FE*	*without FE*	*with FE*
Breech position	0.316***	0.320***	0.317***	0.321***
	(0.018)	(0.017)	(0.018)	(0.017)
Pre-eclampsia			0.077***	0.081***
			(0.016)	(0.016)
Mother's age	0.007***	0.007***	0.007***	0.007***
	(0.001)	(0.001)	(0.001)	(0.001)
Married	0.000	0.003	-0.000	0.002
	(0.007)	(0.007)	(0.007)	(0.007)
Household income	0.001	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)
GCSE/O-level(or eq)	0.001	0.002	0.000	0.001
	(0.008)	(0.008)	(0.008)	(0.008)
A level or more but no uni	-0.006	-0.006	-0.007	-0.007
	(0.010)	(0.010)	(0.010)	(0.010)
University qualification	0.000	-0.001	0.000	-0.001
	(0.011)	(0.011)	(0.011)	(0.011)
White	-0.035***	-0.029**	-0.037***	-0.031**
	(0.009)	(0.014)	(0.009)	(0.014)
Smoker	0.020***	0.020***	0.021***	0.021***
	(0.007)	(0.007)	(0.007)	(0.007)
Parity	0.129***	0.128***	0.126***	0.125***
	(0.007)	(0.007)	(0.007)	(0.007)
Planned pregnancy	-0.002	-0.002	-0.002	-0.001
	(0.006)	(0.006)	(0.006)	(0.006)
BMI	0.006	0.006	0.005	0.006
	(0.005)	(0.005)	(0.005)	(0.005)
BMI (sq.)	0.006	0.006	0.005	0.006
	(0.005)	(0.005)	(0.005)	(0.005)
Hospitalisation	0.041***	0.038***	0.021**	0.018*
	(0.009)	(0.009)	(0.009)	(0.009)
Employed	0.004	0.005	0.004	0.004
	(0.006)	(0.006)	(0.006)	(0.006)

Diabetes	0.067***	0.066**	0.067***	0.066***
	(0.026)	(0.025)	(0.025)	(0.025)
Hypertension	0.122**	0.118**	0.081	0.076
	(0.052)	(0.052)	(0.052)	(0.052)
Kidney diseases	-0.016	-0.018	-0.016	-0.017
	(0.03)	(0.03)	(0.03)	(0.03)
Underweight baby	0.138***	0.140***	0.136***	0.138***
	(0.017)	(0.017)	(0.017)	(0.017)
Overweight baby	0.071***	0.072***	0.071***	0.072***
	(0.009)	(0.009)	(0.009)	(0.009)
Gestational age	-0.138***	-0.130***	-0.141***	-0.133***
	(0.034)	(0.034)	(0.034)	(0.034)
Gestational age (sq.)	0.002***	0.002***	0.002***	0.002***
	(0.000)	(0.000)	(0.000)	(0.000)
Constant	2.543***		2.596***	
	-0.66		-0.655	
No. Observations	14221	14189	14221	14189

Table A2. Linear IV Models: second stage results

	IV: Breech position		IV: Breech position & Pre-eclampsia	
	without FE	*with FE*	*without FE*	*with FE*
	(1)	*(2)*	*(3)*	*(4)*
Unplanned CS	0.186***	0.153***	0.187***	0.152***
	(0.05)	(0.049)	(0.049)	(0.049)
Mother's age	-0.001	-0.001	-0.001	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)
Married	-0.038***	-0.039***	-0.038***	-0.039***
	(0.01)	(0.01)	(0.01)	(0.01)
Household income	-0.003***	-0.003***	-0.003***	-0.003***
	(0.000)	(0.000)	(0.000)	(0.000)
GCSE/O-level(or eq)	0.018	0.019	0.018	0.019
	(0.015)	(0.015)	(0.015)	(0.015)
A level or more but no uni	0.012	0.012	0.012	0.012
	(0.017)	(0.016)	(0.017)	(0.016)
University qualification	0.011	0.012	0.011	0.012
	(0.019)	(0.018)	(0.019)	(0.018)
White	-0.003	-0.007	-0.003	-0.007
	(0.015)	(0.017)	(0.015)	(0.017)
Smoker	0.085***	0.082***	0.085***	0.082***
	(0.010)	(0.010)	(0.010)	(0.010)
Parity	-0.049***	-0.044***	-0.049***	-0.044***
	(0.012)	(0.012)	(0.013)	(0.013)
Planned pregnancy	-0.035***	-0.035***	-0.035***	-0.035***
	(0.008)	(0.008)	(0.008)	(0.008)
BMI	-0.003	-0.004	-0.003	-0.004
	(0.006)	(0.006)	(0.006)	(0.006)
BMI sq.	0.000	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Hospitalisation	0.082***	0.082***	0.081***	0.081***
	(0.010)	(0.010)	(0.010)	(0.010)
Employed	-0.036***	-0.035***	-0.036***	-0.035***
	(0.010)	(0.010)	(0.010)	(0.010)
Diabities	-0.022	-0.021	-0.022	-0.021
	(0.029)	(0.029)	(0.029)	(0.029)
Hypertension	-0.016	-0.015	-0.016	-0.014
	(0.066)	(0.066)	(0.065)	(0.065)
Kidney diseases	0.074*	0.069*	0.074*	0.069*
	(0.040)	(0.039)	(0.040)	(0.039)

Underweight baby	0.014	0.018	0.014	0.018
	(0.021)	(0.021)	(0.021)	(0.021)
Overweight baby	-0.037***	-0.038***	-0.037***	-0.038***
	(0.013)	(0.014)	(0.013)	(0.013)
Gestational age	-0.085***	-0.086***	-0.085***	-0.087***
	(0.031)	(0.032)	(0.031)	(0.032)
Gestational age (sq.)	0.001***	0.001***	0.001***	0.001***
	(0.000)	(0.000)	(0.000)	(0.000)
Constant	2.148***		2.147***	
	(0.592)		(0.592)	
No. Observations	14221	14189	14221	14189

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