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Social interactions in inappropriate behavior for childbirth services: Theory and evidence from the Italian hospital sector

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Abstract

Empirical evidence supports the conjecture that social interactions among agents can produce both positive and negative effects. We build on this literature by exploring the role of social interactions in the hospital sector using the large incidence of cesarean sections, usually considered an inappropriate outcome in the childbirth service. In doing so, we lay out a simple model of hospitals' behavior where the effect of peers' behavior emerges simply by sharing the same institutional authority responsible for auditing inappropriate behavior. In this setting, enforcement congestion induces a peer effect among hospitals that could make inappropriate behaviors more likely. Then, using the risk-adjusted cesarean section rate of a large panel of Italian hospitals, we empirically investigate whether the behavior of each hospital is affected by the behavior of hospitals within the same region, after controlling for demand, supply, and financial factors. In particular, our empirical test employs both peer effects estimate and the spatial econometric approach, exploiting the panel dimension of our data. Both estimates show a significant and strong presence of peer effects among hospitals. We interpret this evidence as a presence of constraint interactions within the hospital sector, which has important implications for healthcare policies against inappropriateness.

JEL Classification: I11, C31

Keywords: social interactions, peer effects, cesarean section, spatial econometrics.

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1. Introduction

Over the past several decades, the role of social interactions has become increasingly important in economic discussions. Even if the origin of social interactions can be found in the sociological literature (Crane, 1991; Mayer, 1991), by now it is acknowledged that the interaction among agents can produce both positive and negative effects. For instance, economic literature in education investigates the positive effects on students' outcomes of the interaction with classmates (Epple and Romano, 1998, 2011; Sacerdote, 2001; Zimmerman, 2003). Similarly, in health economics literature Apouey and Picone (2014) show the existence of social interactions for malaria preventive behaviors. On the other hand, negative effects of interaction have been found in crime (Glaeser et al., 1996), tax evasion (Galbiati and Zanella, 2012) and health behavior (Trogdon et al., 2008; Auld, 2011).

In this study, we explore the role of social interactions in inappropriate behavior in hospitals. More specifically, we first derive a model of hospitals' behavior where the effect of peers' behavior emerges simply by sharing the same institutional authority responsible for auditing inappropriate behavior. In particular, the enforcement congestion generated by the audit system induces a peer effect among hospitals that could make inappropriate behaviors more likely. Then, we test our predictions using the risk-adjusted cesarean section rate in a large panel of Italian hospitals for the period 2007–2012 as a case study. Our focus on this form of social interaction in the Italian hospital sector, as opposed to others such as social norms and trust, reflects the institutional setting of the Italian NHS, which is characterized by a high degree of decentralization at the regional level[1]. In fact, each region is responsible for its healthcare services and cannot afford to prosecute inappropriate behavior in all hospitals within the region; therefore, inappropriate behavior of hospitals within a region makes the open road to the inappropriate behavior of their peers.

Since our empirical test aims to ascertain whether being "surrounded" by other hospitals who are more likely to use (inappropriate) medical procedures has an impact on hospitals' behavior, focusing on cesarean sections seems particularly appropriate. The recent worldwide upward trend in cesarean rates (OECD, 2011) has drawn the attention of both scholars and

[1] The Italian NHS is an especially interesting case for testing our hypothesis, as decentralization processes have made the Regional Health Authorities the main institutional authorities for each hospital and, in particular, the third-party payers for the health services provided. Moreover, not only Italy exhibits one of the highest cesarean rates among the OECD countries but, as a result of the high decentralization, great variation exists across regions both in the regulation and in the delivery of childbirth services (e.g., Francese et al., 2014; Cavalieri et al., 2014).

policymakers, raising concern about the clinical appropriateness of some cesarean deliveries[2]. Furthermore, cesarean section rates exhibit an extraordinarily high variation among regions (Grant, 2005; Baicker et al., 2006), even after adjusting for risk factors. In Italy, about 38% of all child deliveries are performed via cesarean section, a percentage well beyond the WHO (1985) recommended level of 15%. In this regard, among others[3], a recent national report by the Italian Ministry of Health concluded that 43% of the cesarean sections executed in 2010 appeared unjustified, based on information in the patients' discharge records.

Along with clinical factors, many explanations for high cesarean rates have been explored in the literature, including maternal age (Abdul-Rahim et al., 2009), physicians' perceptions of the safety of the procedure (Hopkins, 2000; Kabakian-Khasholian et al., 2007), "defensive medicine" (Dubay et al., 1999; Grant and McInnes, 2004), social and cultural factors (Lo, 2003; Hsu et al., 2008). However, together these factors do not account for the majority of the observed variation. For this reason, many studies have investigated the hypothesis that providers are motivated by financial incentives in their choice of child delivery method, finding that they play a significant role in explaining cesarean section rates (Gruber et al., 1999; Grant, 2009). With regard to Italy, Cavalieri et al. (2014) find that whenever the regional reimbursement policy favors cesarean sections, providers have an incentive to shift deliveries to the more highly reimbursed cesarean procedure.

Indeed, the problem of financial incentives motivating inappropriate behavior in hospitals is not specific to cesarean sections, rather it is a more general problem associated with the tariff systems[4]. Therefore, many NHSs have tried to combine the tariff system with some form of auditing to combat the unintended consequences of tariff mechanisms (Busse et al., 2011). Nonetheless, while diverting considerable human and financial resources, audit systems do not seem to be effective in stemming inappropriate behavior in health systems (Lomas et al., 1991; Ivers et al., 2012), especially in medical procedures in the gray area of medicine (Chandra et al., 2011), such as childbirth medical services.

In our empirical analysis, we first perform a traditional peer effect estimate, really close to our model of hospitals' behavior. As will be shown, our empirical test is a particularly fortune case of peer effects analysis, since our non-linear model does not suffer from the

[2] In particular, in absence of specific therapeutic reasons, the alternative vaginal delivery is generally considered a more appropriate treatment (e.g., Althabe et al., 2006; Betrán et al., 2007; Belizán et al., 2007). Moreover, medically unjustified cesarean deliveries have implications not only for patients but also for the overall society, as they impose a financial burden on the system, while diverting resources from other public services.

[3] See e.g., Fortino et al. (2002), Rusticali and Di Virgilio (2010).

[4] For a recent discussion on the typical incentives, both desirable and undesirable, associated with the tariff systems, see e.g. Cots et al. (2011).

"reflection problem" of linear-in-mean models (Manski, 1993; Brock and Durlauf, 2007; Blume et al., 2011). Finally, we conduct the more recent (but less microfounded) spatial econometric analysis, where the spatial weights matrix is based again on the sharing of the same institutional authorities, in line with a few contributions claiming the primary importance of institutions respect to geography (Rodrik et al., 2004; Arbia et al., 2009; Atella et al., 2014). Both estimates show a significant and strong presence of peer effects among hospitals, with important implications for healthcare policies against inappropriateness.

The remainder of the paper is structured as follows. In Section 2, we describe our model of hospitals' behavior and derive the main implication in terms of peer effects. In Section 3, we describe our data, along with the Italian hospital system. The empirical strategy is presented in Section 4, followed by our results in Section 5. Finally, Section 6 concludes with a discussion on the implications for healthcare policies against inappropriateness.

2. The model

In this section we lay out a model of hospitals' behavior where the effect of peers' behavior emerges by sharing the same institutional authority. In this study, hospitals choose between two competing treatments (i.e., vaginal and cesarean section) for each patient, which will result in the hospital's cesarean section rate. Following previous studies on hospital behavior (Ellis and McGuire, 1986; Chandra et al., 2011), we assume that hospitals select treatments to maximize their objective function. Therefore, the resulting cesarean section rate might not be "appropriate" for the overall healthcare system.

Consider a population of N risk neutral hospitals $i = 1, 2, \ldots, N$, distributed across R institutional health authorities $r = 1, 2, \ldots, R$, each of size n_r. The health authorities are the reference third-party payers for the health services provided by each hospital, as well as being responsible for the "appropriateness" of the healthcare system in each group r. However, they operate under a stringent budget constraint, implying that only a fraction of hospitals in each r can be audited by the reference health authority (RHA).

For each patient, hospitals choose between two competing treatments $\{0, 1\}$, where 0 is usually considered the most appropriate treatment in the absence of specific therapeutic reasons[5]. Each treatment produces a different benefit $B(\sigma)$ for patients, according to patient's characteristics σ. In particular, the utility of patient *k* from the two treatments is given by[6]:

$$U_k(0) = B^0(\sigma_k) + \varepsilon_k^0 \tag{1}$$

$$U_k(1) = B^1(\sigma_k) + \varepsilon_k^1 \tag{2}$$

where the error terms capture the heterogeneity in the benefits of each treatment to that patient, and follow a standard extreme value distribution $\varepsilon^{0,1} \sim EV(0,1)$. On the other hand, each treatment implies a different cost $C(\sigma, X)$, according to the patient's characteristics σ and the hospital's characteristics X.

If hospital *i* chooses the appropriate treatment for patient *k*, it chooses 1 over 0 provided that the improved benefit compensates for the increased costs, that is:

$$\begin{aligned} Pr\{treat_{ik} = 1\} &= Pr\{B^1(\sigma_k) - \lambda C^1(\sigma_k, X_i) + \varepsilon_k^1 > B^0(\sigma_k) - \lambda C^0(\sigma_k, X_i) + \varepsilon_k^0\} = \\ &= Pr\{[B^1(\sigma_k) - B^0(\sigma_k)] - \lambda[C^1(\sigma_k, X_i) - C^0(\sigma_k, X_i)] + \xi_k > 0\} = \\ &= Pr\{\xi_k < [B^1(\sigma_k) - B^0(\sigma_k)] - \lambda\,[C^1(\sigma_k, X_i) - C^0(\sigma_k, X_i)]\} = \\ &= \frac{e^{[B^1(\sigma_k)-B^0(\sigma_k)] - \lambda\,[C^1(\sigma_k,X_i)-C^0(\sigma_k,X_i)]}}{1 + e^{[B^1(\sigma_k)-B^0(\sigma_k)] - \lambda\,[C^1(\sigma_k,X_i)-C^0(\sigma_k,X_i)]}} \end{aligned} \tag{3}$$

where we have exploited the fact that $(\varepsilon_k^1 - \varepsilon_k^0) = \xi_k \sim Logistic(0,1)$. The parameter λ in (3) is usually called the *value of life* and captures the trade-off made by the third-party payer (RHA) between improved benefit and increased costs (Murphy and Topel, 2006). Equation (3) represents the probability for patient *k* of receiving treatment 1 in hospital *i*. Therefore, integrating (3) over the distribution $f(\sigma)$ of patient's characteristics σ produces the appropriate cesarean section rate of hospital *i*:

$$CS_i^{app} = \int_\sigma \left\{ \frac{e^{[B^1(\sigma_k)-B^0(\sigma_k)] - \lambda\,[C^1(\sigma_k,X_i)-C^0(\sigma_k,X_i)]}}{1 + e^{[B^1(\sigma_k)-B^0(\sigma_k)] - \lambda\,[C^1(\sigma_k,X_i)-C^0(\sigma_k,X_i)]}} \right\} f(\sigma)\, d\sigma \tag{4}$$

However, hospitals select treatments to maximize their objective function. Following previous studies (McGuire and Pauly, 1991; Chandra et al., 2011), while hospitals consider the benefit for patients, they also consider the financial incentives associated with the two treatments. In particular, the welfare for hospital *i* from providing the two treatments to patient *k* are given by:

$$W_{ik}(0) = \beta_i B^0(\sigma_k) + V^0(F_r^0) - C^0(\sigma_k, X_i) + Z_r^0 + \varepsilon_k^0 \tag{5}$$

$$W_{ik}(1) = \beta_i B^1(\sigma_k) + V^1(F_r^1) - C^1(\sigma_k, X_i) + Z_r^1 + \varepsilon_k^1 \tag{6}$$

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[5] As we discussed above, in our specific case the vaginal birth, in absence of specific therapeutic reasons, is usually considered the more appropriate treatment (e.g., Althabe et al., 2006; Betrán et al., 2007; Belizán et al., 2007), not only for medical but also for economic reasons.

[6] Notice that we are implicitly assuming that patients do not pay any price for the health service they get; or, alternatively, they pay exactly the same price regardless the treatment, which is essentially the Italian case.

where β_i reflects the relative importance of patient benefit respect to financial aspects, $V(F)$ capture the expected revenues from the two treatments as a function of fees, and Z denote other contextual factors affecting hospitals in each group *r*. Specifically, whenever hospitals claim for treatment 0 (usually considered the most appropriate), they always receive the established fee, $V^0(F_r^0) = F_r^0$. In contrast, when hospitals claim for treatment 1, what they receive depends on being audited by the RHA. In particular, when hospital *i* is not audited, it receives F_r^1 for each claim. Instead, when hospital *i* is audited, it receives a reduction in fee proportional to the "inappropriate" claims. Therefore, the expected revenue from a treatment 1, in each group *r*, is given by:

$$V^1(F_r^1) = (1 - p_{ir})\, F_r^1 + p_{ir}\, \min\left\{1, \frac{CS_i^{app}}{CS_i^*}\right\} F_r^1 \quad (7)$$

where p_{ir} is the probability for hospital *i* of being audited by the RHA for group *r*. Reasonably, if hospital *i* claims an appropriate share of treatment 1 (i.e., $CS_i^* \leq CS_i^{app}$)[7], according to (7) it does not receive any reduction in fees. Instead, if $CS_i^* > CS_i^{app}$, then hospital *i* might be subject to a reduction if audited.

As noted earlier, each RHA operates under a stringent budget constraint, implying that each hospital is audited only with some positive probability. However, the probability of being audited is not completely random, as the behavior of hospitals provides information to the RHAs. For the sake of simplicity, we assume the following linear specification, which captures the effect of hospitals' behavior within the group *r* on hospital *i*'s audit probability:

$$p_{ir} = \frac{n_r^{control}}{n_r} + \gamma Pr\left(CS_i^* > CS_i^{app} \middle| CS_i^*\right) - \frac{\theta}{n_r - 1} \sum_{j \in r, j \neq i} Pr\left(CS_j^* > CS_j^{app} \middle| CS_j^*\right) \quad (8)$$

The first term of (8) is the random probability of being audited, equal to the fraction of hospitals the health authority r can afford to audit, with $n_r^{control} < n_r$. The other terms capture the idea that RHAs infer the probability of inappropriate claims from hospitals' behavior (i.e., from CS^*) and, accordingly, shift their audit resources to more suspicious hospitals. Indeed, we do not want to model explicitly the way in which RHAs estimate the probability of inappropriate claims, as it might differ across *r*. Nonetheless, it is reasonable to assume that higher is the amount of claims for treatment 1, higher is the estimated probability of inappropriate claims, that is:

$$\frac{\partial Pr\left(CS_i^* > CS_i^{app} \middle| CS_i^*\right)}{\partial CS_i^*} > 0 \qquad \forall\ i \in r,\ \forall\ r \in R \quad (9)$$

[7] Since $F_r^1 > F_r^0\ \forall\, r \in R$ (that is, caesarean section fees are higher than vaginal fees), there should be no incentive for hospitals to claim $CS_i^* < CS_i^{app}$, as the profit associated with cesarean sections is usually higher than that associated with vaginal delivery. Certainly, this is the case for hospitals in the Italian NHS (e.g., Francese et al., 2014; Cavalieri et al., 2014).

Interestingly, as long as $CS_i^* > CS_i^{app}$, (9) it is enough to imply that:

$$\frac{\partial V^1(F_r^1)}{\partial CS_i^*} = \left(\frac{CS_i^{app}}{CS_i^*} - 1\right) F_r^1 \frac{\partial Pr(CS_i^* > CS_i^{app} | CS_i^*)}{\partial CS_i^*} - F_r^1\, p_{ir} \frac{CS_i^{app}}{CS_i^{*2}} < 0 \tag{10}$$

$$\frac{\partial V^1(F_r^1)}{\partial CS_{j \in r, j \neq i}^*} = \left(1 - \frac{CS_i^{app}}{CS_i^*}\right) F_r^1 \frac{\theta}{n_r - 1} \frac{\partial Pr\left(CS_j^* > CS_j^{app} \middle| CS_j^*\right)}{\partial CS_j^*} > 0 \tag{11}$$

Therefore, when other hospitals within the same group become more suspicious for the RHA, this reduces hospital *i*'s probability of being audited and, in turn, increases the expected value of a marginal inappropriate claim. In other words, the last term of (8) represents the effect of peers' behavior on the other hospitals sharing the same RHA.

In contrast to (3), hospital *i* chooses treatment 1 over treatment 0 for patient *k* provided that $W_{ik}(1) > W_{ik}(0)$, that is:

$$\begin{aligned} Pr\{treat_{ik} = 1\} = \; & Pr\{W_{ik}(1) > W_{ik}(0)\} = Pr\{I_{ik}(1) + \varepsilon_k^1 > I_{ik}(0) + \varepsilon_k^0\} = \\ & = Pr\{I_{ik}(1) - I_{ik}(0) + \xi_k > 0\} = Pr\{\xi_k < I_{ik}(1) - I_{ik}(0)\} = \\ & = \frac{e^{[I_{ik}(1) - I_{ik}(0)]}}{1 + e^{[I_{ik}(1) - I_{ik}(0)]}} \end{aligned} \tag{12}$$

where $I_{ik} = \beta_i B(\sigma_k) + V(F_r) - C(\sigma_k, X_i) + Z_r$ and $(\varepsilon_k^1 - \varepsilon_k^0) = \xi_k \sim Logistic(0, 1)$. Then, integrating (12) over the distribution $f(\sigma)$ of patient's characteristics σ, we obtain:

$$CS_i^* = \int_\sigma \left\{\frac{e^{[I_{ik}(1) - I_{ik}(0)]}}{1 + e^{[I_{ik}(1) - I_{ik}(0)]}}\right\} f(\sigma)\, d\sigma \equiv G(CS_i^*) \tag{13}$$

Note that (13) represents the hospital *i*'s reaction function (as a fixed point), because it tells us the optimal cesarean section rate as a function of all other hospitals' (within *r*) cesarean section rates. Interestingly, from (10) we have that:

$$\frac{\partial G(CS_i^*)}{\partial CS_i^*} = \frac{\partial V^1(F_r^1)}{\partial CS_i^*} \int_\sigma \left\{\frac{e^{[I_{ik}(1) - I_{ik}(0)]}}{\left[1 + e^{[I_{ik}(1) - I_{ik}(0)]}\right]^2}\right\} f(\sigma)\, d\sigma \begin{cases} < 0 & if \quad CS_i^* > CS_i^{app} \\ = 0 & if \quad CS_i^* \leq CS_i^{app} \end{cases} \tag{14}$$

Therefore, as shown in Figure 1, (13) gives a unique CS_i^* best response:

- FIGURE 1 about here –

Moreover, applying the implicit function theorem to the equilibrium condition (13), we have the main implication of the model concerning the presence of peer effects on hospitals' inappropriate behavior, that is:

$$\frac{\partial CS_i^*}{\partial CS_{j \in r, j \neq i}^*} = -\frac{\partial D(CS_i^*, CS_{j \in r, j \neq i}^*)/\partial CS_{j \in r, j \neq i}^*}{\partial D\left(CS_i^*, CS_{j \in r, j \neq i}^*\right)/\partial CS_i^*} > 0 \tag{15}$$

where we have posed the equilibrium condition (13) in the implicit form:

$$D\left(CS_i^*, CS_{j\in r, j\neq i}^*\right) = CS_i^* - \int_\sigma \left\{\frac{e^{[I_{ik}(1)-I_{ik}(0)]}}{1+e^{[I_{ik}(1)-I_{ik}(0)]}}\right\} f(\sigma)\, d\sigma \tag{16}$$

Figure 2 shows the effect of peers' behavior on hospital i's cesarean section rate.

- FIGURE 2 about here –

The main intuition behind (15) is that, once other hospitals within r exhibit higher cesarean section rates, they become relatively more suspicious for the RHA, which accordingly focuses on them. In other words, peers with higher cesarean section rates reduce the hospital's probability of being audited, which increases the expected value of inappropriate claims.

Regardless the extent of peer effects[8], the model has a unique Nash equilibrium (as a fixed point) with many interesting empirical implications. Firstly, the inappropriate behavior of hospitals and, in particular, cesarean rates should be spatially correlated according to the sharing of the same institutional authority rather than the geographic distance among peers. Indeed, this view of peer effects among hospitals differs from the usual interpretation of learning from the reference school treatment style (Epstein and Nicholson, 2009). Secondly, the presence of peer effects should generate an excess variance in equilibrium, meaning that even a small difference in fundamentals among hospitals (belonging to different institutional authorities) might produce large differences in cesarean section rates. Therefore, peer effects contribute to explain the large regional variation in health services not explained by differences in fundamentals (Skinner, 2011).[9]

To some extent, the aforementioned process generating peer effects is similar to models of enforcement congestion developed in other strands of literature, such as crime behavior (Sah, 1991) and tax evasion (Galbiati and Zanella, 2012). However, our model presents the significant advantage that each hospital can have a limited number of peers within the

[8] Linear-in-mean models would need a stability requirement on the magnitude of peer effects, such as the impact of peers' behavior on the probability of being audited θ has to be somewhat lower than the impact of own behavior γ (e.g., Galbiati and Zanella, 2012). Provided that probabilities are well-behaved, in our non-linear-in-mean model such requirement is not needed for guarantee a unique Nash equilibrium, as the equilibrium would converge in any case, as can be gathered from Figure 2. Nonetheless, this restriction on the magnitude of peer effects remains certainly reasonable also in our context and, in particular, we will see below that estimated marginal peer effects are strictly less than 1 in all estimates.

[9] The other implications of the model are more standard in the literature on hospital behavior (Chandra et al., 2011). In particular, to the extent that different hospital types (private vs. public) might evaluate patients utility differently in their objective function (i.e., $\beta_i \neq \beta_j$), they could exhibit different equilibrium cesarean section rates. Similarly, those institutional authorities with a higher fee differential (i.e., $F_r^1 - F_r^0 > F_l^1 - F_l^0$) could have hospitals with higher equilibrium cesarean section rates. Finally, as long as the provision of childbirth services is characterized by economies of scale and/or *learning-by-doing* effects, hospitals with different characteristics and/or different degrees of specialization (i.e., $C^{0,1}(\sigma_k, X_i) \neq C^{0,1}(\sigma_k, X_j)$) could also exhibit different cesarean rates, *ceteris paribus*.

reference group, making more reasonable the idea that agents somehow guess the probability of being audited by the RHA after observing the behavior of their peers.

An important feature of our model is that we focus on one source of social interaction only (i.e., *constrain interactions* within the same RHA), although other social forces may be at work. Nonetheless, we have good reasons to do so in our context. Firstly, the usual argument that widespread inappropriate behaviors violate social norms, the strength of which decreases with the diffusion of such behavior within the reference group, does not seem to apply in our context, as high cesarean rate is not necessarily viewed as negative by patients (Grant, 2005; Fusco et al., 2010). Similarly, previous literature on physicians' treatment styles (a potential source of *preference interactions* among peers) shows that physicians do not significantly change their style of childbirth treatments because of local peer interactions (Epstein and Nicholson, 2009). Moreover, many studies emphasize that higher tariff differentials among alternative treatments induce hospitals to shift deliveries to the more highly reimbursed procedure (Grant, 2009; Cavalieri et al., 2014); thus, it seems reasonable to investigate whether peers' behavior within the reference institutional authority represents a significant constraint for hospitals' inappropriate behavior. Finally, the source of social interaction emphasized here identifies the reference group exogenously for each hospital, as opposed to being the outcome of an arbitrary choice of the researcher. This should represent a significant advantage of our model, especially for the subsequent empirical analysis, as it is well known that wrong assumptions on reference groups can be strongly misleading (Conley and Topa, 2003).

3. Data description

Our empirical analysis uses data provided by several sources. Our dependent variable is provided by the National Program for Outcome Assessment (*PNE*) run by the National Agency for Regional Health Services (*AGENAS*), together with the Italian Ministry of Health. Since its inception in 2009, the program has assessed the healthcare activity of all Italian hospitals, public and private, providing several performance and outcome indicators[10]. In particular, the indicator employed here is the risk-adjusted cesarean section rate at the hospital

[10] Overall, 45 performance indicators (32 related to hospital services and 13 to hospitalization) are computed, mostly using discharge data gathered through the Informative Hospital System (*SIO*). For more information on PNE see *http://151.1.149.72/pne11_new*.

level for first-time mothers[11], aged 10–55 years. The cesarean rate for first-time mothers provided by the *PNE* is adjusted for maternal age and comorbidities (main and secondary diagnoses for admissions during the last two years), as well as for *a-priori fetus* risk factors[12]. Other variables related to the structural characteristics of providers (number of beds, yearly birth deliveries, hospital type) are provided by the Italian Ministry of Health. Among the demand factors potentially relevant, we consider the female employment rate at the province level, provided by the Italian National Institute of Statistics (*ISTAT*). Finally, data on regional fees for childbirth deliveries come from the *AGENAS* and, in some cases, directly from Italian RHAs.

All data have been checked for reporting errors, outliers, and missing values, and refer to the period 2007–2012. Therefore, the final data set consists of a panel of 492 public and private Italian hospitals within the study period, yielding 2,952 observations. Table I shows the summary statistics for the variables employed in the analysis, while Appendix 1 contains some more descriptive statistics for the risk-adjusted cesarean rates.

- TABLE I about here –

To investigate the presence of peer effects among hospitals belonging to the same RHAs, Figure 3 plots hospital risk-adjusted cesarean rates against the average cesarean rate of their peers. Despite the significant heterogeneity, probably due to the different hospital types in the Italian NHS, the fitted values seem to reveal a positive relationship between hospital cesarean rates and the (inappropriate) behavior of their peers belonging to the same RHAs (correlation coefficient $\rho = 0.67^{***}$).

- FIGURE 3 about here –

With regard to regional payment policies for childbirth deliveries, we match data from several sources, mainly from the *AGENAS*, but also directly from RHAs. We consider tariffs for ordinary admissions (longer than one day) for two specific DRGs (Medical Disease Classification 14), namely DRG 370 (cesarean section with complications and comorbidities)

[11] Compared to the overall rate of cesarean deliveries, this indicator is considered better suited to capturing the phenomenon of clinical inappropriateness, since it is not strongly influenced by the high risk of cesarean delivery for those women who have already experienced a cesarean section and by their distribution among hospitals.
[12] Risk adjusted cesarean rates for first time mothers are reported by the *PNE* only for those hospitals with at least 10 childbirths in the selected year.

and DRG 372 (vaginal delivery with complications)[13], for each RHA in the study period. In the Italian NHS, each RHA is free to apply the national tariffs or to set their own fees. As shown in Table I, tariffs for cesarean sections are higher than those for vaginal deliveries, because the former is a surgical intervention, performed in an operating room by a surgeon. Even higher tariffs are set for childbirth deliveries in the presence of complications and comorbidities.[14]

In particular, in our empirical analysis we consider the following tariff differential indicator:

$$FEEDIFF = \frac{(RegFEE_{DRG370}/RegFEE_{DRG372})}{(NatFEE_{DRG370}/NatFEE_{DRG372})}$$

The idea behind this indicator is that hospital cesarean rates are driven more by the tariff differential than by the amount of each tariff. Therefore, a higher regional tariff differential between cesarean and vaginal deliveries implies a greater incentive for providers to behave strategically by opting for a cesarean section[15]. More specifically, a value of 1 for *FEEDIFF* indicates that a RHA has applied the national tariffs for both cesarean and vaginal deliveries. A value greater (less) than 1 designates a RHA where the ratio between the two tariffs is higher (lower) than the corresponding national one, implying a relative financial benefit of executing cesarean sections. Figure 4 provides an overview of *FEEDIFF* by region for the last year in our sample, showing that RHAs in Italy have opted for different tariff policies.

- FIGURE 4 about here -

4. Empirical strategy

In this section we present our empirical strategy to test for the presence of peer effects in cesarean section rates among Italian hospitals. We first perform a traditional peer effects estimate, really close to our model of hospitals' behavior. Then, we carry out also the more recent (but less microfounded) spatial econometric analysis, where we emphasize the primary importance of institutions respect to geography.

[13] Indeed, we might consider also the two DRG tariffs for childbirth deliveries without complications and comorbidities, namely DRG 371 (cesarean section without complications and comorbidities) and DRG 373 (vaginal delivery without complications). However, in Cavalieri et al. (2014) it is found that only DRG tariffs with CC are significant in driving providers' behavior, probably because of the greater difficulty for providers to "induce" and justify a cesarean delivery in absence of complications and comorbidities.

[14] For a more detailed description of the Italian NHS and the specific hospital financing system for childbirth services, see e.g. Cavalieri et al. (2014, Section 2.2).

[15] By asserting this, we are implicitly assuming that costs are relatively homogeneous at the national level, at least among the same type of providers, which seems quite reasonable in the Italian context (e.g., Francese et al., 2014; Cavalieri et al., 2014).

4.1 Logit model

From the equilibrium condition (13), we estimate the following logit model for the risk-adjusted cesarean rates:

$$RACR_{irkt} = \frac{e^{\left[\beta * Dem_{irt} + \tau * Bed_{irkt} + \gamma * Birth_{irkt} + \delta * FEEDIFF_{rt} + \rho \frac{\sum_{j \in r, j \neq i} RACR_{jrkt}}{n_r - 1} + \theta_k + \varphi_r + \mu_t\right]}}{1 + e^{\left[\beta * Dem_{irt} + \tau * Bed_{irkt} + \gamma * Birth_{irkt} + \delta * FEEDIFF_{rt} + \rho \frac{\sum_{j \in r, j \neq i} RACR_{jrkt}}{n_r - 1} + \theta_k + \varphi_r + \mu_t\right]}}$$

that is

$$Logit\ RACR_{irkt} = \alpha + \beta * Dem_{irt} + \tau * Bed_{irkt} + \gamma * Birth_{irkt} + \delta * FEEDIFF_{rt} + \\ + \rho \frac{\sum_{j \in r, j \neq i} RACR_{jrkt}}{n_r - 1} + \theta_k + \varphi_r + \mu_t + \varepsilon_{irkt} \quad (17)$$

The dependent variable *RACR* is the risk-adjusted cesarean rate in hospital *i* of type *k* in region *r* in year *t*. The risk-adjustment procedure ensures that we have already considered the demographic and clinical factors driving the differences in cesarean rates among hospitals. The first group of explanatory variables *Dem* controls for potential differences in preferences among patients driven by socio-economic factors rather than risk factors. We have the following demand factors: the female employment rate *FER* at the provincial level, usually considered the catchment area for hospitals providing childbirth services; a regional capital dummy, where *RC* = 1 if hospital *i* is located in a province that is the regional capital; and a province capital dummy, where *PC* = 1 if hospital *i* is located in a municipality that is the provincial capital. Then, we control for the following supply factors potentially affecting cesarean rates: *Bed* is the total number of beds; and *Birth* is the total number of childbirth deliveries in hospital *i* in year *t*, capturing the size and the level of specialization of hospitals[16]. Considering the previous evidence on the relevance of *learning-by-doing* effects in the provision of health care services (Birkmeyer et al., 2002; Chandra et al., 2011), these supply factors might be important in explaining the shares of cesarean sections. To follow, the explanatory variable *FEEDIFF* is our DRG tariff differential indicator, as described in the previous section[17].

[16] In the first version of this paper we considered also some other demand and supply factors potentially affecting cesarean sections, as local female tertiary education rate, local household income, number of personnel units. Not surprisingly, all these variables turn out to be highly correlated with the others already considered, inducing a multicollinearity problem in our estimates. Therefore, we removed them from the estimates, knowing that we are already controlling for the underlying factors potentially driving differences among cesarean rates.

[17] The variable *FEEDIFF* only considers the variability of DRG tariffs among Italian regions, capturing the relative financial incentive to execute cesarean sections in each region. Unfortunately, data availability prevents us from also considering the variability of DRG tariffs among different hospitals within the same region. Therefore, the magnitude of the impact of *FEEDIFF* on hospital cesarean rates may be underestimated.

The variable $\frac{\sum_{j \in r, j \neq i} RACR_{jrkt}}{n_r - 1}$ is our main variable of interest, and it is the average risk-adjusted cesarean rate of all hospitals belonging to the same institutional authority *r*, excluding hospital *i* (i.e., the average inappropriate behavior of the peers of hospital *i*). Consequently, the coefficient ρ is the peer effect on inappropriate behavior among hospitals.

Finally, we include a large set of hospital (θ), regional (φ), and time (μ) fixed effects to capture those unobservable differences among hospitals, regions, and years that could affect cesarean section rates. This large set of fixed effects should help us to alleviate omitted variables bias and model misspecification.

As long as exogeneity $E(\varepsilon_{irkt}|X_{irkt},) = 0$ holds, we could interpret the estimated coefficients as consistent. In equation (17), we are particularly interested in estimating ρ. Here, [exp(ρ) – 1]*100 can be interpreted as the percentage change in the odds of the share of cesarean sections due to a marginal increase in the average cesarean rate of the reference peers. As far as the identification of ρ is concerned, this is a particularly fortune case of peer effects analysis, since our non-linear model does not suffer from the "reflection problem" of the linear-in-mean models. In fact, the non-linearity in our model (17) breaks the linear dependence between the outcome variable $RACR_{irkt}$ and the endogenous effect $\frac{\sum_{j \in r, j \neq i} RACR_{jrkt}}{n_r - 1}$, which is the basis of Manski's (1993) result of nonidentification in the linear case (see also Blume et al., 2011). Moreover, the use of panel data in (17) allows us to consistently address group level unobservables φ potentially correlated with the endogenous effect, which represents a potential source of nonidentification even in non-linear models[18]. In particular, Brock and Durlauf (2007, Proposition 7, p. 67) provides the formal result of the identification of endogenous effects in non-linear models with panel data[19].

Because the number of cross-sectional observations is significantly larger than the number of time-series ones, heteroscedasticity could be a problem in our estimates. In particular, the share of cesarean sections might exhibit a different variability by hospital size,

[18] In this regard, the identification of endogenous effects in non-linear models with group level unobservables would require specific restrictions on the joint distribution of observables and unobservables, likely not reasonable in many contexts under analysis (e.g., Brock and Durlauf, 2007; Blume et al., 2011).

[19] Notice that Brock and Durlauf (2007, Proposition 7, p. 67) provides the formal result of partial (not full) identification in non-linear models with group level unobservables, meaning that not all parameters can be identified. However, the unique parameter not being identified is that relating to the time invariant group-specific characteristics; indeed, this is not surprising because there is no way to distinguish between the time invariant group-specific characteristics and group level unobservables. Looking at our model (17), however, we do not have any time invariant group specific characteristics, but only time variant group specific characteristics which are fully identified, along with the endogenous effect ρ.

specialization and type, eventually implying heteroscedastic residuals. Therefore, for all our estimates, we provide standard errors that are robust to the presence of heteroscedasticity.

4.2 Spatial econometric model

An alternative empirical approach to study social interactions among agents is the more recent spatial econometric. Several studies have performed spatial analysis to infer similar effects in agents' behavior (Moscone et al., 2012; Gravelle et al., 2014; Atella et al., 2014). An interesting empirical implication of our model is that cesarean section rates should be spatially correlated according to the sharing of the same institutional authority. Therefore, as further evidence of the significant presence of peer effects among hospitals, we carry out the following spatial econometric analysis, exploiting the panel dimension of our data.

Looking at the specific model, when spatially lagged dependent variable, regressors, and error term are included simultaneously, the model is not identified unless at least one of these interaction effects is excluded (Manski, 1993). As suggested by LeSage and Pace (2009), the choice of the excluded effect should be driven by the specific research question. According to our model in Section 2, hospitals' behavior within the same region should be affected by their peers' behavior rather than their characteristics. Specifically, hospital i's cesarean section rate (13) is affected by its peers' cesarean section rates rather than their supply factors. Therefore, we specify the following spatial autoregressive model with autoregressive disturbances (SAC), provided that we perform standard model selection tests:

$$RACR_{irkt} = \beta * X_{irkt} + \delta * FEEDIFF_{rt} + \rho \sum_j w_{ij} RACR_{irktj} + \theta_k + \varphi_r + \mu_t + \varepsilon_{irkt} \quad (18)$$

$$\varepsilon_{irkt} = \phi \sum_j m_{ij} \varepsilon_{irktj} + \upsilon_{irkt} \quad (19)$$

where X_{irkt} is the vector of demand and supply factors, and w_{ij} and m_{ij} are the (i,j) elements for each *t* of the spatial matrices ***W*** and ***M***, respectively[20].

In contrast to the standard spatial analysis, the spatial weights matrix is not based on the geographic distance among hospitals. Rather, it is based on the sharing of the same institutional authorities, to be consistent with our model of hospitals' behavior. More specifically, the row-standardized spatial weights matrices ***W*** = ***M*** are as follows:

$$w_{ij} = m_{ij} = \begin{cases} 0 & if\ i = j \\ \frac{1}{n_r - 1} & if\ r_i = r_j \\ 0 & if\ r_i \neq r_j \end{cases} \quad (20)$$

[20] According to the spatial econometric literature (e.g., Anselin, 1988), the two spatial matrices ***W*** and ***M*** (respectively, for the dependent variable and disturbances) can be different; however, in the following spatial analysis we consider ***W*** = ***M***.

This implies that each hospital is correlated only with other hospitals within the same region. Therefore, the spatial weights matrix (20) emphasizes the primary role of institutions respect to geography in affecting agents' behavior (Rodrik et al., 2004; Arbia et al., 2009; Atella et al., 2014). In particular, this reflects our interpretation of the effect of peers among hospitals as a *constraint interaction* within the same institutional authority.

5. Results

First, we show the estimates for the logit model, along with the generalized linear model (GLM), potentially an econometric specification even more appropriate for cesarean section rates. Then, for the spatial analysis, we present our estimates for different spatial econometrics models, along with standard model selection tests.[21]

5.1 Logit model

Table II reports the estimates from the logit model (17). As discussed earlier, heteroscedasticity may be present in our estimates. Therefore, we use a generalized least squares (GLS) estimator for panel data[22]. As shown in column (1), the point estimate of our main variable of interest *Peers RACR* is positive and strongly significant. In particular, the coefficient of 0.027 implies a marginal effect of 0.006, meaning that an increase of one percentage point in peers' cesarean rates would increase hospitals' cesarean rates by about 0.6 percentage points[23]. Therefore, our estimate of *Peers RACR* suggests a significant presence of peer effects among Italian hospitals.

- TABLE II about here –

With regard to the demand and supply factors, while the female employment rate at the provincial level (*FER*) does not seem to play a role[24], both the regional (*RC*) and province

[21] We have also conducted different sensitivity analyses to check the robustness of our estimates, finding results fully in line with both the logit and the spatial econometric estimates reported here in the paper. For a matter of space, we have not reported our robustness checks, but the interested reader can easily find them in the working paper version of the paper (Guccio and Lisi, 2014).
[22] We have also tried to run a pooled OLS estimator with robust standard errors (POLS results available upon request), obtaining coefficients fairly close to GLS in Table II but, not surprisingly, estimates were less precise.
[23] In particular, we are considering the standard marginal effect at means (MEMs), that is

$$\frac{\partial RACR}{\partial Peers\ RACR} = \rho\big(F(\overline{\boldsymbol{X}}\boldsymbol{\beta}) - F(\overline{\boldsymbol{X}}\boldsymbol{\beta})^2\big) = \rho\,\frac{e^{(\overline{\boldsymbol{X}}\boldsymbol{\beta})}}{\big(1 + e^{(\overline{\boldsymbol{X}}\boldsymbol{\beta})}\big)^2}$$

capital (*PC*) dummies are positive and significant at the 1% significance level. This implies that, on average, patients in big provinces and big cities tend to prefer more cesarean deliveries, *ceteris paribus*. Among the supply factors, the number of childbirth deliveries (*Birth*) is significant, while hospital size (*Bed*) does not seem to play any role, probably because we control for hospital type. In particular, the negative coefficient of *Birth* indicates that more specialized hospitals tend to exhibit lower cesarean rates, suggesting the presence of a *learning-by-doing* effect in the provision of childbirth services.

With regard to the financial factor, the point estimate of *FEEDIFF* is positive and significant at the 1% significance level, implying that higher tariff differentials are associated with higher cesarean rates. In particular, the coefficient of 0.279 implies a marginal effect of 0.058, meaning that a marginal increase of *FEEDIFF* would increase the probability of a cesarean delivery by 5.8 percentage points. Thus, in those regions with a relatively greater financial incentive to execute cesarean sections, providers respond in a strategic way by shifting procedures towards more cesarean deliveries.

Finally, our estimates show interesting regularities in the differences among hospital types. In particular, directly managed public hospitals (*Hospital Units*), where financial factors should not play a big role, tend to experience lower cesarean rates. Conversely, accredited private hospitals (*Private Hospitals*), where financial aspects should be crucial, tend to execute more cesarean sections, *ceteris paribus*. These findings suggest that providers behave strategically in accordance with financial incentives, which is consistent with the previous literature.

From an econometric point of view, although the logit model (17) is more appropriate than the simple linear probability model, it might not be the most appropriate for cesarean rates. In this regard, Papke and Wooldridge (1996) note that the log-odds type procedures implicitly assume a standard normal distribution for the error term, which might not be appropriate for regression models with fractional dependent variables. Therefore, since fractional variables are the result of a dichotomous process, they propose a more attractive quasi-likelihood estimation method in the framework of generalized linear models (GLM), using the logit transformation as link function, but assuming a binomial distribution for the

[24] We find similar results if we include local female tertiary education rate or local household income, instead of female employment rate; indeed, this is not surprising as all these demand factors tend to be positively correlated in our sample.

error term[25]. To the extent that the share of cesarean sections is the result of the dichotomous choice between vaginal/cesarean delivery, GLM could be even more appropriate.

Therefore, in the second column of Table II, we run the same model, but for the described GLM estimator. Nonetheless, column (2) shows that all the results are also confirmed by the GLM estimates. In particular, the coefficient of *Peers RACR* is still positive and strongly significant, implying a marginal effect of about 0.5 percentage points in hospitals' cesarean rates. Similarly, all other results concerning the demand, supply, and financial factors are in line with the logit estimates.

Overall, both the logit and GLM estimates seem to support our predictions on hospitals' behavior. Provided that we control for demand, supply, and financial factors in (17), we interpret our estimate as the (reduced-form) hospitals' reaction function (13) to the inappropriate behavior of their peers within the RHA. Therefore, our estimate of *Peers RACR* suggests a significant presence of peer effects among hospitals. In particular, our estimate implies that an increase of one percentage point in peers' cesarean rates leads to an increase of about 0.6 percentage points in hospitals' cesarean rates.

Following the classification proposed by Manski (2000), we interpret this evidence as a presence of substantial *constraint interactions* in the hospital sector, meaning that the behavior of peers represents a constraint for the inappropriate behavior of hospitals within the same authority. Looking at our theoretical model, this interpretation would seem fairly reasonable for the specific context of the Italian healthcare sector.

5.2 Spatial econometric model

We now move to the spatial analysis. Since the implication of our theoretical model is that cesarean rates should be spatially correlated within the same institutional authority, our spatial analysis could provide further support for our model, thus, providing evidence of peer effects among hospitals.

First, we test pre-emptively whether cesarean section rates show spatial dependence within the same RHA. Table III shows the results of the Moran's I test, using (20) as the spatial weights matrix. As can be seen, we find evidence of spatial dependence among

[25] Indeed, Papke and Wooldridge (1996) proposed the so-called "fractional logit" in the cross-section context. However, there are no serious drawbacks in applying their GLM approach with panel data, provided that one "... can account for unobserved heterogeneity that is possibly correlated with the explanatory variables ..." (Papke and Wooldridge, 2008). In this regard, we are confident that in our study the large structure of fixed effects should be sufficiently able to account for the unobserved heterogeneity, without suffering from the incidental parameters problem. For a study applying the "fractional logit" with panel data see e.g. Hausman and Leonard (1997), where they use a similar strategy to account for the unobserved heterogeneity.

cesarean section rates, regardless of whether we compute the Moran's I test on the whole data set or per year.

- TABLE III about here –

Then, Table IV shows the results of different spatial econometric models. With regard to model selection, hospitals' behavior should be affected by their peers' behavior, not their characteristics. Accordingly, we first estimate a spatial autoregressive model with autoregressive disturbances (SAC), as specified in (18) and (19). As can be seen, the spatial effect in the dependent variable (*RHO*) is positive and strongly significant, whereas the spatial effect in the error term (*LAMBDA*) is not statistically significant. Therefore, we estimate a spatial autoregressive model (SAR), excluding the autoregressive disturbances and compute standard model selection tests. For the sake of completeness, we also estimate a spatial Durbin model, including supply factors as spatial regressors. The *LR* tests, the Akaike information criterion (*AIC*), and the Bayesian information criterion (*BIC*) show a clear preference for the SAR. Therefore, we consider the SAR as the preferred model in our spatial analysis.

As Table IV clearly shows, the empirical results support the prediction of our theoretical model. In particular, the spatial coefficient *RHO* in the dependent variable is positive and strongly significant, implying dependence among hospitals' behavior within the same RHA. On the other hand, hospital cesarean rates do not seem to have any relation with their peers' characteristics, as our model predicts. In terms of the magnitude of our estimated spatial effect, we find that *RHO* is 0.396, which is consistent with previous studies estimating spatial correlation among health providers' behavior within the same institutional authority (e.g., Atella et al., 2014). In terms of the other factors, the empirical results are similar to those of the logit model for demand, supply, and financial factors, as well for the differences among hospital types.

- TABLE IV about here –

Therefore, our spatial analysis apparently confirms the prediction of our model concerning the spatial correlation among hospital cesarean rates within the same RHA. Although the spatial econometric models are less microfounded than the logit model (17), the evidence in Table IV provides further support for our theoretical model.

6. Conclusion

In this study, we explored the presence of social interaction effects in the inappropriate behavior among hospitals sharing the same institutional authority. The main intuition of our theoretical prediction is that each institutional authority cannot afford to tackle the inappropriate behavior of all hospitals under its authority. Therefore, higher hospitals' inappropriate behavior can produce an enforcement congestion effect, which makes the open road to the inappropriate behavior of their peers. Subsequently, we tested the implications of our model in the Italian hospital sector, controlling for demand, supply, and financial factors. Here, our estimates show a significant and strong presence of peer effects among hospitals. Following the classification proposed by Manski (2000), we interpret this evidence as a presence of substantial *constraint interactions* in the hospital sector, meaning that the behavior of peers represents a constraint on the inappropriate behavior of hospitals within the same institutional authority responsible for auditing inappropriate behavior.

Our results have important implications for healthcare policies against inappropriateness. First, through the social interaction effect, health authorities could reduce inappropriate behavior at a cost less than that of auditing all hospitals under their authority. However, consistent with previous literature, the significant presence of *constraint interactions* among hospitals and the high level of inappropriateness in Italian regions suggest that auditing policies do not appear to be effective in reducing inappropriate behavior. Instead, policies that set resource allocation on outcome targets (e.g. pay for performance), regardless of audits on appropriateness, has the potential to work better as a measure against inappropriateness, especially in medical procedures in the gray area of medicine, such as the childbirth medical services. Moreover, policies based on outcome targets could lead to consistent savings in human and financial resources respect to auditing systems on the appropriateness of each medical procedure provided by hospitals.

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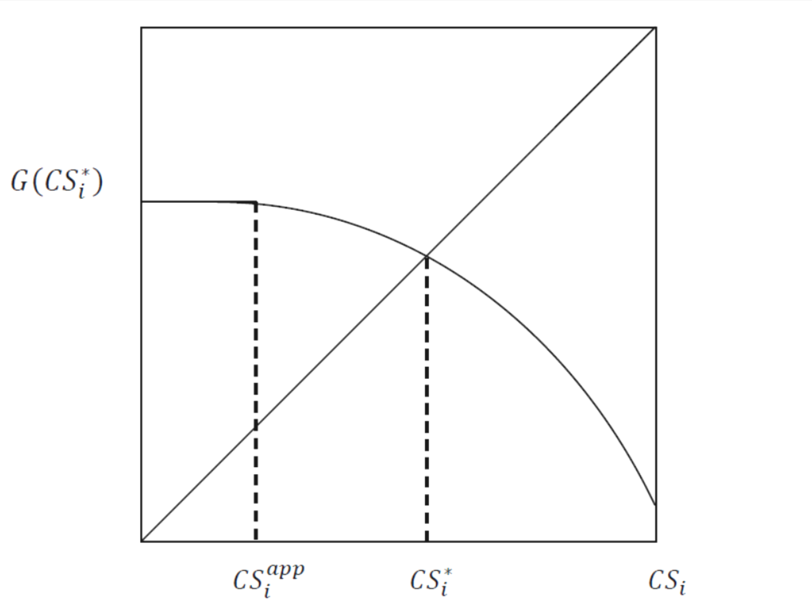
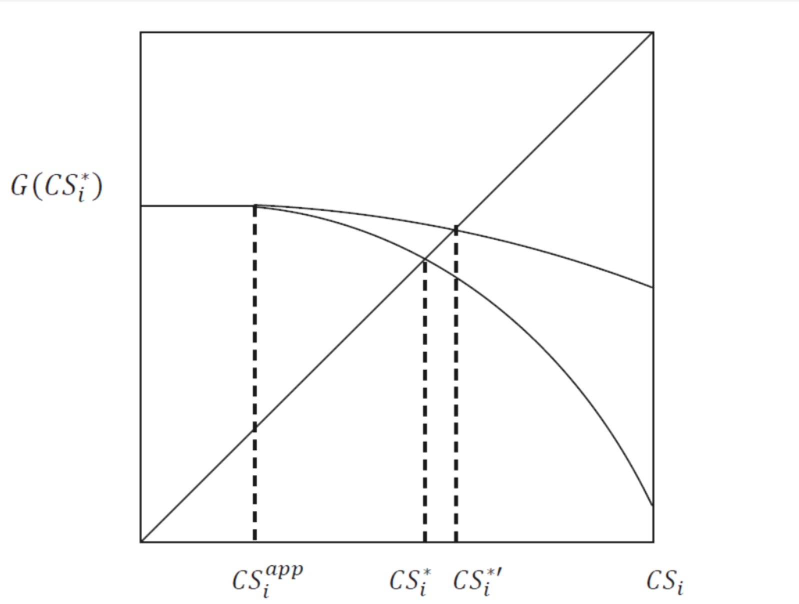
TABLES AND FIGURES

Figure 1: Optimal CS_i^* best response



Note: The figure depicts the CS_i^* best response (as a fixed point) given by the equilibrium condition (13), showing that the hospital i' s optimal cesarean section rate is unique.

Figure 2. Peers effect on equilibrium CS_i^*



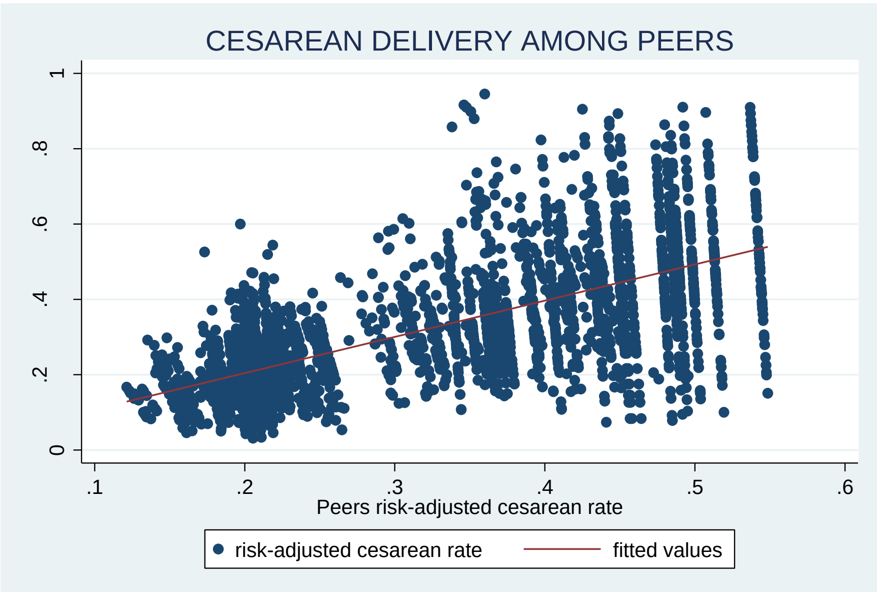
Note: The figure depicts the effect of peers' behavior on the hospital i' s cesarean rate given by equation (15), showing that being "surrounded" by peers with higher cesarean section rates increases the optimal CS_i^* best response.

Table I: Descriptive statistics

Variables	Obs.	Mean	Std. Dev.	Min	Max
Risk-adj. cesarean rate (*RACR*)	2952	0.31	0.16	0.03	0.94
Vaginal birth fee W CC	120	2280.18	416.08	1370	3180
Cesarean section fee W CC	120	3588.74	647.66	2316	4955
Tariff differential index (*FEEDIFF*)	120	1.01	0.14	0.67	1.42
Female employment rate (*FER*)	654	47.21	11.74	22.71	64.82
Number of beds (*Bed*)	2952	397.97	328.54	25	1719
Number of births (*Birth*)	2952	830.97	662.85	90	7313

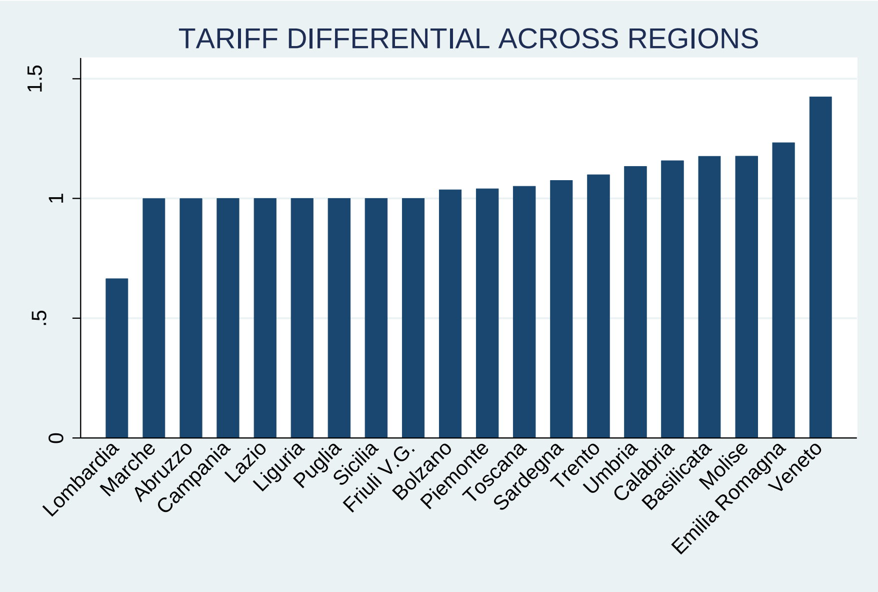
Note: The table presents the main descriptive statistics of the variables employed in our empirical analysis. The risk-adjusted cesarean rate considers only first-time mothers, because it is usually considered a better index of inappropriateness respect to the overall cesarean rate. *FEEDIFF* is the index of fee differential between cesarean and vaginal DRG tariffs W CC, as described in section 3.

Figure 3: Cesarean section rates and correlation among hospital peers



Note: The figure plots the hospital risk-adjusted cesarean rates against the average cesarean rate of their peers within the same RHA. The slope of the superimposed regression line reveals a positive relationship between hospital cesarean rates and the (inappropriate) behavior of their peers belonging to the same RHAs.

Figure 4: Tariff differential indicator across Italian regions



Note: The figure provides an overview of our tariff differential indicator (FEEDIFF) across the Italian regions. Specifically, a value of 1 of *FEEDIFF* indicates that a RHA has applied the same tariffs as the national ones for both cesarean and vaginal deliveries. Differently, a value higher (lower) than 1 designates a RHA where the ratio between the two DRG tariffs is higher (lower) than the corresponding national one, implying a relative financial convenience to execute a cesarean section. Overall, the figure shows that different RHAs in Italy have opted for different tariff policies.

Table II: Risk-Adjusted Cesarean Rates for First-Time Mothers (LOGIT Model)

	(1)	(2)
	LOGIT[a]	GLM[b]
Peers RACR	0.027	0.025
	(0.002)***	(0.002)***
FER	-0.003	-0.003
	(0.002)	(0.002)
RC	0.074	0.111
	(0.020)***	(0.022)***
PC	0.085	0.081
	(0.022)***	(0.023)***
Bed	0.000	0.000
	(0.000)	(0.000)
Birth	-0.001	-0.001
	(0.000)***	(0.000)***
FEEDIFF	0.279	0.226
	(0.089)***	(0.086)***
Hospital Unit	-0.061	-0.098
	(0.036)*	(0.039)**
Private Hospital	0.626	0.632
	(0.038)***	(0.037)***
Constant	-1.883	-1.670
	(0.162)***	(0.181)***
Hospital type dummies	YES	YES
Regional dummies	YES	YES
Year dummies	YES	YES
Observations	2952	2952

LOGIT: logit model; GLM: generalized linear model for fractional regression; FER: female employment rate; RC: regional capital; PC: province capital; FEEDIFF: index of fee differential (differential between cesarean and vaginal DRG tariffs W CC);
[a] GLS estimator for panel data. [b] Generalized linear estimator for fractional variable by Papke and Wooldridge (1996).
Robust standard errors in brackets. * significant at 10%, ** significant at 5%, *** significant at 1%.

Table III: Moran's I Tests

Test	**I Statistics**	**p-value**
Moran's I_{Panel}	0.473	0.000
Moran's I_{2007}	0.473	0.000
Moran's I_{2008}	0.451	0.000
Moran's I_{2009}	0.451	0.000
Moran's I_{2010}	0.440	0.000
Moran's I_{2011}	0.498	0.000
Moran's I_{2012}	0.508	0.000

Note: The table shows the results of the Moran’s I tests, using (20) as the spatial weights matrix. The first row concerns the Moran’s I test computed to the whole panel, whereas the other rows concern the Moran’s I tests computed to each year one by one in our dataset.

Table IV: Risk-Adjusted Cesarean Rates for First-Time Mothers (Spatial Model)

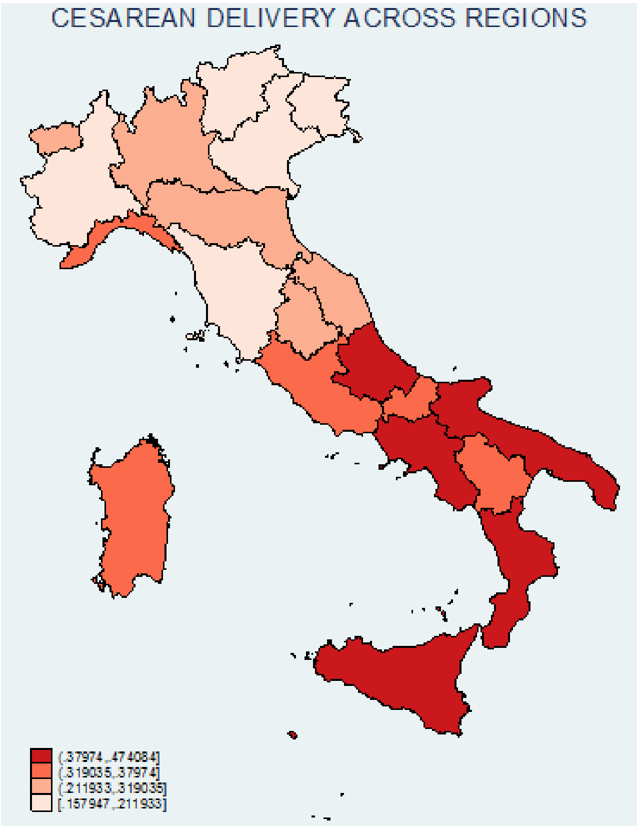
	(1)	(2)	(3)
	SAC	SAR	DURBIN
FER	-0.001	-0.001	-0.001
	(0.001)	(0.001)	(0.001)
RC	0.027	0.027	0.029
	(0.005)***	(0.005)***	(0.005)***
PC	0.014	0.013	0.014
	(0.005)***	(0.005)***	(0.005)***
Bed	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)
Birth	-0.001	-0.001	-0.001
	(0.000)***	(0.000)***	(0.000)***
FEEDIFF	0.031	0.033	0.031
	(0.018)*	(0.019)*	(0.019)*
Hospital Unit	-0.017	-0.015	-0.016
	(0.008)**	(0.008)*	(0.008)**
Private Hospital	0.152	0.153	0.151
	(0.007)***	(0.007)***	(0.007)***
RHO	0.430	0.396	0.394
	(0.058)***	(0.034)***	(0.034)***
Bed			0.000
			(0.000)
Birth			0.000
			(0.000)
LAMBDA	-0.082		
	(0.126)		
LR Test (SAC vs. SAR)		p-value = 0.511	
LR Test (Durbin vs. SAR)		p-value = 0.999	
AIC	-4589.063	-4590.632	-4577.962
BIC	-4487.229	-4494.788	-4476.128
Hospital type dummies	YES	YES	YES
Regional dummies	YES	YES	YES
Year dummies	YES	YES	YES
Observations	2952	2952	2952

DURBIN: spatial durbin model; SAR: spatial autoregressive model; SAC: spatial autoregressive model with autoregressive disturbances; FER: female employment rate; RC: regional capital; PC: province capital;
Clustered standard errors in brackets. * significant at 10%, ** significant at 5%, *** significant at 1%.

APPENDIX 1: RISK-ADJUSTED CESAREN SECTION RATE IN THE SAMPLE

In this appendix we provide further details on cesarean section rate in our sample. As shown in Figure A.1, a first look at the data reveals great regional variability in cesarean section rates for first-time mothers in Italy. At the regional level, the risk-adjusted cesarean rate varies between a minimum of 14.6% in Alto Adige and a maximum of 50.1% in Campania, with an average national value of 31%. More specifically, cesarean rates are higher in southern regions (42.7% on average) than in the rest of the country (23% on average), as clearly emphasized by Figure A.1. At the hospital level, variation in cesarean rates ranges from almost 3.2% in Lombardy to 94.5% in Lazio.

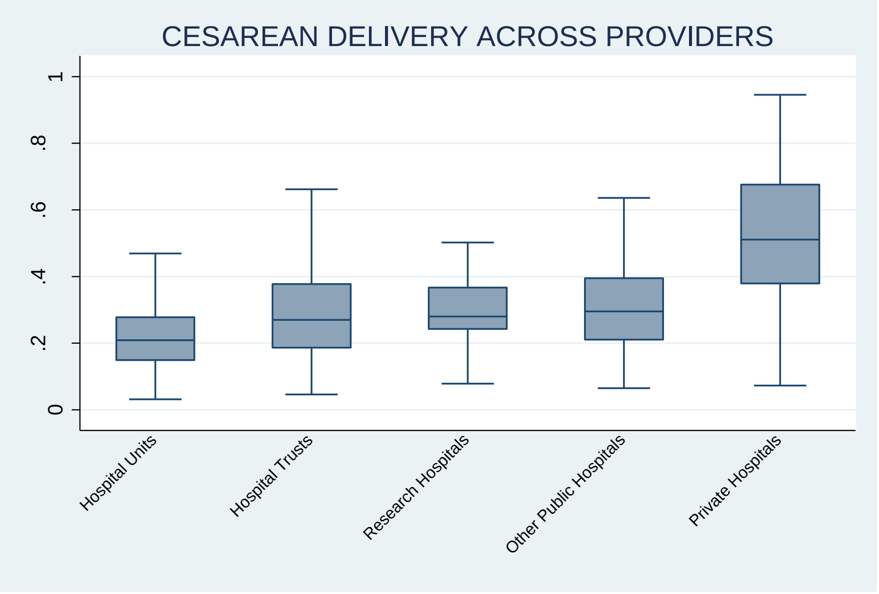
Figure A.1: Cesarean delivery for first-time mothers across Italian regions



Note: The figure shows the intensity of cesarean sections across the Italian regions and, specifically, a higher color intensity means a larger use of cesarean sections. The figure reveals that there is great variability across Italian regions in the use of cesarean section, even after controlling for risk factors.

With regard to the typology of hospitals, as expected, the directly managed public hospitals (Hospital Units) display a lower median value of the risk-adjusted cesarean rate and less variability (see Figure A.2). Not surprisingly, the opposite is true for private hospitals (Private Hospitals), which display the higher median cesarean rate. Looking at the other categories of public hospitals, that is independent (Hospital Trusts) and research (Research Hospitals) hospitals, the latter show a higher median value but less variability of risk-adjusted cesarean rates. Overall, Figure A.2 shows that there is a great variability in cesarean section rates also across providers, which partly explains the high heterogeneity emerging in Figure 3.

Figure A.2: Cesarean delivery for first-time mothers across providers



Note: The figure shows the great variability of risk-adjusted cesarean rates across the different hospital types providing childbirth services in the Italian NHS.