Physician labor supply, financial incentives, and healthcare accessibility

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Abstract

Do physicians respond to financial incentives ? A wide-scale reform in France, where most physicians' fees are regulated, raised the price of a visit from €23 to €25 in 2017. Relying on granular data covering the universe of patients, physicians, and visits, our estimates show that physician activity grew by nearly 9% after the price increase, yielding a unitary price elasticity of care provision. The number of patients seen increased substantially (by the same relative magnitude), when the provision of medical services per patient hardly changed, resulting in a slight increase in physicians' workload. Drugs prescriptions per patient are also shown to decrease, suggesting that the policy was cost-effective and enhanced healthcare accessibility, with limited adverse effects. Early-career physicians respond strongly to higher financial incentives, while later-career physicians remain largely unaffected.

Keywords: Physician labor supply; Financial incentives; Healthcare accessibility; Medical spending.

JEL Classification: I11, I18, J44.

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1 Introduction

How do physicians react to financial incentives? The resulting estimates, across countries and specialties, are (to say the least) dispersed. A precise answer goes beyond labor supply considerations since it impacts healthcare accessibility and individuals' health. All the more so that access to physicians is ever more difficult in most OECD countries because of aging, an increase of obesity, cancer, or diabetes, and more generally an expansion of chronic diseases as well as a deterioration of individuals' mental health.

In the U.S., the government has designated primary care Health Professional Shortage Areas (HPSAs) to address the issue of unequal access to healthcare, which basically consists in encouraging physicians to locate in deprived areas based on financial incentives. A recent working paper by Pilvar (2024) evaluates a similar program in the UK. In France, this policy instrument has also been used, but apparently with limited success (Cour des Comptes, 2024).

Experts in the field have emphasized that the extensive margin (namely, the creation of new practices in medically deprived areas) may be more difficult to incentivize than the intensive margin (providing more medical services, conditional on already practicing in a deprived area) in this respect. To that aim, bargaining between doctors' unions and the public health insurance resulted in the latter committing to revalue physicians fees according to some predetermined scheme, from €23 in 2016 to €25 in 2017, and up to €30 in December 2024.

Our research design exploits the shift in the reimbursement rate of an office visit from €23 to €25, namely a +8.7% price change, which occurred on May 1st, 2017, viewed as an exogenous shock on concerned doctors' financial incentives. This was a large-scale policy change since each and any general practitioner (GP) working in France and affiliated with the public health insurance, the French Cnam, according to so-called *secteur 1* (sector 1) regulating rules was affected by that reform. Sector 1 GPs cannot charge any other price than the regulated one. By contrast, *Secteur 2* (sector 2 hereafter) GPs may charge higher prices, but do not benefit from any rebate on their payroll taxes. Beyond the sole self-employed sector 2 GPs, specialists and salaried doctors were not affected by this reform either, which leaves room for various identification strategies. Our empirical analysis, a DiD approach, precisely compares sector 1 GPs as a treated group with sector 2 GPs as

a comparison group. A possible objection to that strategy relates to the latter competing with the former in the healthcare market. We alleviate concerns about SUTVA being possibly violated in our context by investigating whether sector 2 GPs have been indirectly affected by the reform, hereby relying on an alternate comparison group composed of direct access specialists. We also leverage a comparison among sector 1 GPs, considering 2016 as a baseline or 'comparison' year and 2017 as a 'treated' year, in the spirit of identification strategies that have been used to evaluate the impact of COVID-19 pandemics on various outcomes (mortality, consumption, savings, etc.). We further check that our empirical strategy passes falsification tests, and we implement synthetic DiD approaches to assess its robustness with respect to the common trend assumption (CTA).

Our analysis is based on granular data from the *Système National des Données de Santé* (SNDS), which contains the comprehensive set of all doctor visits for all patients in France between 2016 and 2018. We further avail of exhaustive information on each and any healthcare provider in this country over that period. Exploiting such unique databases allows us to cover the universe of both patients and healthcare providers in the French system, and thus to have a complete picture of this market. To the best of our knowledge, this contrasts with numerous studies based on a subsample of patients or physicians. We observe sociodemographics of both patients and physicians, and we have information on doctors' legal status (self-employed, salaried, or mixed), specialty, not to omit practices (including their location at the municipality level). The data also contains information at the visit level on actual fees charged, as well as on the exhaustive list of drug prescriptions. Importantly, our data permits to obtain precise estimations within a causal inference framework like the one based on the exogenous shock mentioned above.

Equipped with that research design and such data, our identification strategies all conclude to an estimated price elasticity of healthcare provision of about 1, which is lower than in the U.S. (Clemens and Gottlieb, 2014) on the set of Medicare enrollees, but which contrasts with long-lasting empirical evidence of null, small, or even negative elasticities. As a result, we estimate that medical expenditures, which are almost fully covered by the public health insurance, rose by nearly 18% following the regulatory change. Interestingly, we are able to unravel the mechanisms by which sector 1 GPs have accordingly modified their behavior. In particular, we investigate whether they have adjusted their workload (measured in working days), the intensity of care (proxied by the number of medical services provided

to each patient). We also wonder whether healthcare accessibility (namely, the number of patients seen each month) has changed, as well as drug prescription and retirement behaviors (the extensive margin response). Our estimates indicate that doctors' workload slightly increased, the supplementary effort in terms of working days each year accounting for nearly 1/4 of the effect on activity, that the intensity of care hardly varied, but that the reform mostly enhanced healthcare accessibility (about 3/4 of previous effect), presumably through a shorter visit duration. After the reform, GPs have more patients, whom they see as their regular GPs, and that number increases faster than the overall number of patients, consistent with these doctors admitting new patients, and becoming their regular GP. We find little change in travel times from patients to doctors, which suggests that new patients are roughly located in similar areas as former ones. Drug prescription per patient also diminished, especially as antibiotics are concerned, consistent with a positive correlation between the length of a visit and drug prescription in the current setting. The reform seems to have induced concerned GPs to postpone their retirement, but that effect is not large. Overall, our results suggest that increasing doctors' fees per visit, which are rather low by OECD standards, could be a cost-effective public policy in order to enhance the accessibility of healthcare. It should for instance be noted that the direct costs of higher healthcare provision are partially offset by a thriftier drug prescription. A more complete cost-benefit analysis is nonetheless beyond the scope of this paper since it would require monetizing patients' well-being related to the mere fact of going to the doctor, and the improvement (hopefully) of their health conditions as a consequence.

Literature At the source of the literature devoted to physician labor supply, Feldstein (1970), Brown and Lapan (1972), Fuchs and Kramer (1972) generally found small negative elasticities, i.e. backward-bending supply curves that are consistent with relatively strong income effects that dominate substitution effects. Yet those studies were based on aggregate time series data: they typically relied on small sample sizes, which raises concerns as regards the precision of the measure. Sloan (1975) used U.S. census data from 1960 and 1970, and concluded to small positive wage elasticities, on average, combined with a backward-bending labor supply curve for a minority of doctors at the top of the income distribution. The first causal evidence based on micro data dates back to Rizzo and Blumenthal (1994) who instrument for physicians' wages by experience, and focus on young physicians aged less than 40 in order to abstract from retirement considerations. Based on a measurement

of labor supply in working hours, they find a short-run wage elasticity of 0.27, and a larger one for female doctors. According to their estimates, the income elasticity amounts to –0.17, but leisure is an inferior good for female physicians. Showalter and Thurston (1997) rely on another instrument for wages, namely exploiting geographic variation in the maximum marginal tax rate. They find that self-employed physicians have a 0.33 short-run wage elasticity, whereas employed physicians are inelastic. In Norway, Baltagi et al. (2005) leverage a natural experiment in which some male hospital doctors received a 15% wage increase while others did not. Their results point out to an elasticity ranging from 0.3 to 0.34. Another natural experiment in Scotland led Ikenwilo and Scott (2007) to estimate a 0.12 elasticity in the case of NHS hospital consultants. Other studies including Sæther (2005) and Andreassen et al. (2013) adopt structural approaches to estimate physicians' labor supply elasticity, and tend to find even smaller estimates. By contrast, Clemens and Gottlieb (2014) exploit a 1997 Medicare geographic boundary re-adjustment, which yielded area-specific price shocks in physician payments. They strikingly obtain a larger value for that elasticity of labor supply since a +2% increase in reimbursement rates is associated with a +3% increase in healthcare provision to the concerned population of patients aged more than 65. In a later work, Gottlieb et al. (2023) estimate a smaller value comprised between 0.2 and 0.4, though. In Canada where the typical fee per visit is €40, Fortin et al. (2021) find that an increase in the price of all clinical services caused physicians to supply less labor, consistent with both leisure being a normal good and income effects being large there. In France, where doctors' visits tend to be more affordable than in comparable OECD countries, the most recent evidence we are aware of is provided by Coudin et al. (2015) who also find a negative price elasticity of doctors' activity. Based on the 'sector 2-freeze' 1990 reform, which consisted in enforcing new restrictions to join the sector 2, their research design takes advantage of a fuzzy RDD which relies on a drop in the share of newly graduated sector 2 doctors. Their empirical analysis makes use of a pool of two cross-section datasets, and their identification is local since it relies on a subsample of compliers. They conclude that the effect of that tighter regulation for these presumably high-income GPs, previously eligible to the sector 2, has been to decrease prices by 54% as well as to increase activity by 41%.

The rest of the article is organized as follows. Section 2 presents the institutional background. Section 3 is devoted to the data. Our empirical analysis is exposed in section 4.

Section 5 contains our main results and various robustness checks including answers to concerns about threats to identification, falsification tests, and an analysis of the heterogeneity of treatment effects. Section 6 concludes.

2 Institutional setting

In France, one half of doctors are self-employed, including 60% of the 100,000 GPs, according to Table 7 in Appendix 2 of Cour des Comptes (2024), and 45% of the 125,000 specialists, according to Drees. Almost all of them, roughly 115,000, are fee-regulated by the *Caisse nationale de l'assurance maladie* (Cnam), the French public health insurance that provides both mandatory and universal healthcare. Each and any physician must opt for a payment system that is either sector 1, which corresponds to a fixed reimbursement rate set by the regulator, or sector 2, a more flexible regime whereby prices for medical services may exceed that rate, up to some threshold above which they are capped. 85,000 physicians, namely more than 90% of GPs and a small majority of specialists, choose the sector 1 option while 30,000 opt for the sector 2.

In France, the healthcare market is regulated so that patients' out-of-pocket (OOP) expenditures are almost zero when they visit their regular general practitioner. As regards sector 1 physicians, the rate of an office visit is set by Cnam. Patients usually pay for medical services at the end of the visit, and are then reimbursed by Cnam. Exceptions include patients who benefit from *tiers payant*: these patients are exempted from upfront payment, depending on their socio-economic status and the coverage provided by their supplementary health insurance. 96% of French households were covered by supplementary health insurance in 2017. In the *tiers payant* case, healthcare providers are directly paid by the public health insurance. Poorest patients are eligible to the *Complémentaire santé solidaire* (CSS), a means-tested supplementary health insurance which is free of charge, which covers against OOP medical expenditures, and which insures that the concerned patients do not have to pay upfront, *de facto* offering universal healthcare along with a *tiers payant* payment scheme. To mitigate moral hazard concerns and to temper the increasing trend in health expenditures, a €1 deductible had been implemented by the public health insurance since 2005. By construction, this deductible implies that OOP

expenditures are non zero. That deductible has doubled since March 31st, 2024.[1]

In 2016, a nationwide bargaining session, the so-called *Convention nationale 2016-2021*, took place between the regulator, i.e. Cnam, and doctors' unions. At the end of these negotiations, a statement was publicly released on August 25, 2016. It was then announced that a supplementary fee of €2 per visit would prevail from May 1, 2017 onward. Before that date, the reimbursement rate of an office visit amounted to €23. The reform we currently study thus raised that rate to €25 (Figure 1). We view this 8.7% increase in fees for medical services as an exogenous shock on doctors' financial incentives.

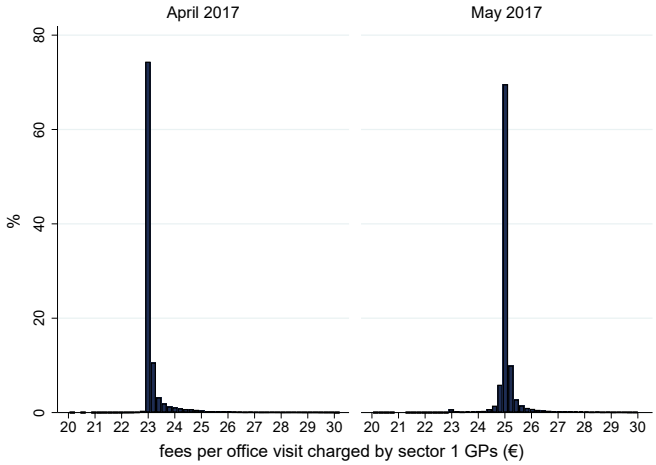


Figure 1: Actual fees charged by sector 1 GPs

Though the agreement concerned sector 1 GPs only, some sector 1 specialists were affected by that regulatory change, too. In particular, pediatricians have benefited from an increase in the reimbursed fee of their office visits, depending on (essentially decreasing with) the age of the child. Mandatory procedures specific to infants were raised by €8, as opposed to other procedures that experienced a €1 increase only, combined with no price change for pupils aged 6 to 16 (when not referred to by their regular GP). Other changes happened, but with a different timing than the one that prevailed for office visits to sector 1 GPs.

Importantly as regards our identification strategy exposed below, the overall demand

[1] Young individuals and foreigners, among others, are exempted from that deductible.

for healthcare has remained stable over the period, i.e. slightly above than 250 million visits per year (see Figure 2 in Cour des Comptes (2024)).

3 Data

Our empirical analysis is based on the *Système National des Données de Santé* (SNDS), a French comprehensive database, and the administrative source that records each and any medical medical service for every patient who ever utilizes healthcare. Indeed, our data cover the universe of both patients and healthcare providers in the French system. It contains detailed information on office visits, patients' home visits, or video visits, but also on prescription behavior, hospitalization, death, etc. Our observation period runs from 2016 to 2018, but we choose to exclude November and December 2018 from our sample because of right-censoring. Indeed, the processing time for claim files may last up to 40 days, and the date of record may thus well fall after the end of 2018. We aggregate our data at the doctor-month level. In January 2016, our working sample contains about 52,000 self-employed GPs, either fee-regulated (sector 1) or not (sector 2), defined as such based on their specialty identifier,[2] with a positive activity along the 34 months when they may be observed (see Appendix Figure A1). Activity is measured here by the number of medical services (office visits, patients' home visits, telemedicine) provided by doctors. We restrict our attention to doctors aged 30 to 79 practicing in mainland France, namely metropolitan France (excluding overseas) at the exclusion of Corsica. For technical reasons, we also impose that their practice be located in a municipality for which information on population and density is non-missing in databases provided by Insee, the French institute in charge of statistics. Similarly, for each of these municipalities, one should have a non-missing continuous index called APL (for *Accessibilité Potentielle Localisée*, that is, a location-based index for potential healthcare accessibility) that helps Drees, the statistical institute of the Department of Health define medical deserts. Our working sample contains about 1.7 million observations at the physician-month level. Depending on our identification strategy, self-employed direct access specialists (gynecologists, ophthalmologists, and stomatologists) who fulfill the above inclusion criteria may also be included in our sample.

There is a pervasive trend to rarefaction of supply since the number of GPs keeps

[2] General medicine corresponds to codes 1, 22, and 23 in the corresponding classification.

on decreasing over the period until about 50,000 in October 2018, and this diagnosis is in line with Cour des Comptes (2024). About 20% of these GPs are not active every month, as shown by Appendix Figure A2. An age-sex pyramid is provided by Figure A3, which indicates that female doctors are much younger than their male colleagues. It hence suggests that in the near future that profession will become more feminine, and this is confirmed by panel a) of Figure A4: the share of women increased from 35% in January 2016 to 39% in October 2018, i.e. +4pp, more than 11% within less than three years. Panel b) of the same Figure shows that a vast majority (more than 92%) of GPs are sector 1 doctors.[3] We learn from panel c) that nearly 35% of GPs' practices are located in dense areas, and that almost the same proportion of practices can be found in intermediate areas, while slightly less than 30% are in weakly dense areas. Practices in non dense areas are extremely scarce: they account for only 0.2% of the total.

The distribution of medical services provided per doctor is given by the histogram of Figure 2. A typical GP makes 400 visits per month, yet there is some heterogeneity in this regard. The median number of visits per month is higher for men (Figure 3, panel a). It exhibits an inverted U-shape with respect to age (Figure 3, panel b), which is almost quadratic with an acme of nearly 460 monthly visits reached at 54 (Figure A10). Denser areas are seemingly associated with lesser visits (Figure 3, panel c), remembering that it is not statistically meaningful to put much weight on what happens in non dense areas.[4] A peak of doctors' activity[5] occurs every Mondays, and activity keeps on decreasing along the rest of the week with a rebound just before the week-end. From the viewpoint of care provision, Saturdays correspond to 1/3 of an ordinary weekly working day, and Sundays are rather anecdotal.[6]

[3] This share exhibits an upward trend, which is partly mechanical and due to the fact that women enroll more often in sector 1 than men.

[4] Activity tends to be higher in medical deserts (see Figure A11). Medical deserts are defined according to the classification into ZIP or ZAC areas, based on the APL. Almost 3/4 of the French population resides either in ZIP or in ZAC.

[5] See Figure 3, panel d) that does not separate GPs from specialists.

[6] Figure A5 further confirms that 1/3 of doctors -again, that Figure relates to both GPs and specialists- work 6 days a week (hence including Saturdays, most of the time), 1/4 of doctors work 5 days a week, while 1/4 have a part-time activity in the sense that they work 4 days or less. On average, GPs work 22 days each month but, again, heterogeneity prevails in this regard. As before, being a male aged 50 to 59 and having a practice located in weakly dense areas is associated with a higher number of working days, both per month (Figure A7) and per year (Figure A8). It is also related to working more often on Saturdays (Figure A9). The number of distinct patients seen per month covaries similarly along those dimensions (Figures A13 and A14).

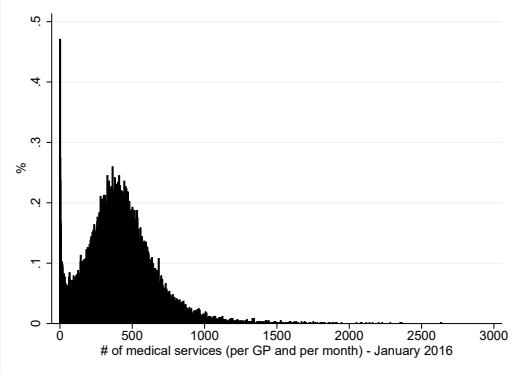


Figure 2: Physician activity

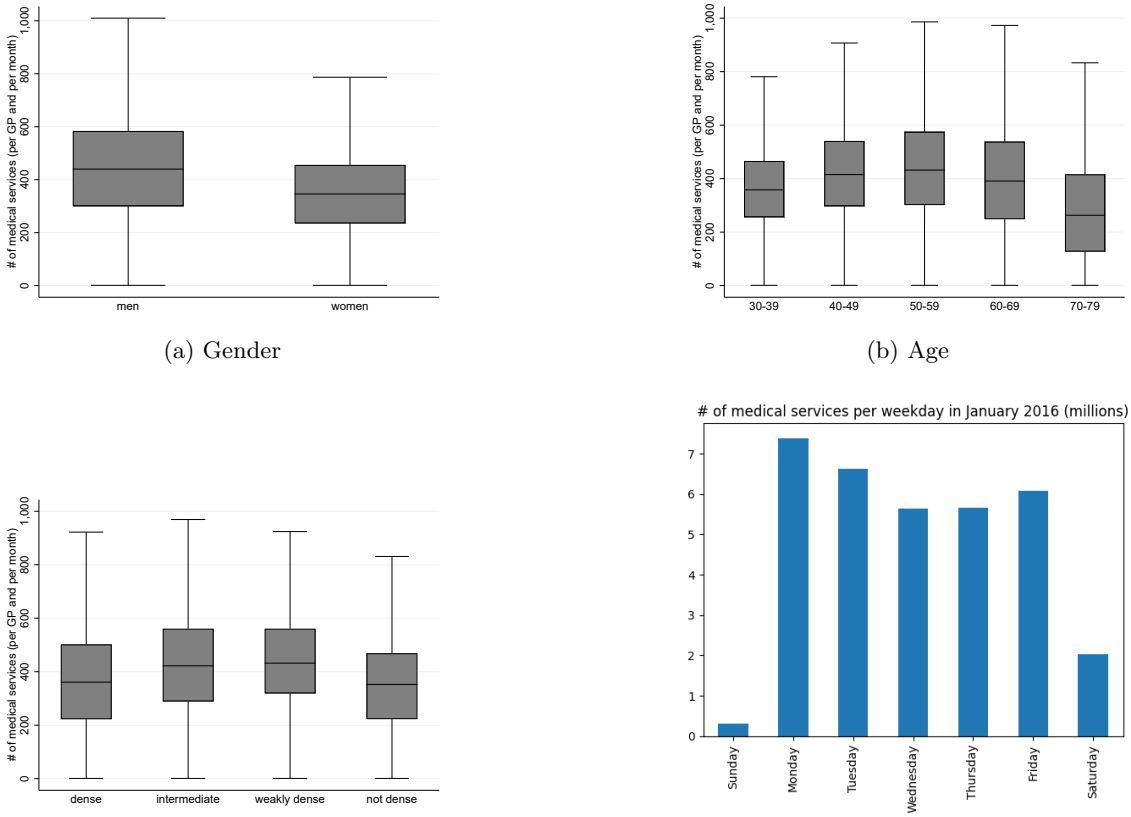
As regards the evolution of GPs' activity over time, a proxy for the aggregate demand for healthcare in the absence of binding capacity constraints, Figure 4 indicates that it decreased at a quiet pace from January 2016 to October 2018.

A proxy for the intensity of care, the median number of medical services per patient (defined as the ratio of the number of medical services over the number of patients seen within a month), tends to be quite homogeneous around slightly less than 1.2 per month, regardless of the location of doctors' practices (Figure A15, panel c), but it monotonically increases with age (Figure A15, panel b) and it is higher for men (Figure A15, panel a).

4 Empirical strategy

4.1 Identification

Our research design relies on a policy-induced shock on doctors' financial incentives, the 2017 increase in the regulated price charged by sector 1 general practitioners for an office visit. We adopt a differences-in-differences (DiD) approach, whereby we consider fee-regulated GPs as a treatment group. Our comparison group is composed of sector 2 GPs who were not concerned by this change in financial incentives. Our identification strategy therefore rests on two assumptions: the absence of spillovers, namely the Stable Unit Treatment Value Assumption (SUTVA), an hypothesis stating that the comparison group was not affected by the regulatory change, and the common trend assumption (CTA), according to which all the outcomes we consider would have evolved similarly in both



(a) Gender

(b) Age

(c) Location

(d) Weekday

Figure 3: Doctors' activity, depending on various dimensions

groups in the absence of that policy. Based on pre-event data, Figure 5 suggests that this assumption cannot be dismissed as regards our two main outcomes of interest, medical activity and corresponding expenditures since no differential pre-trend can be observed, an empirical assessment that indeed comforts the plausibility of the CTA in our setting.

We adopt an alternate identification strategy by considering another comparison group composed of direct access specialists (gynecologists, ophthalmologists, and stomatologists), which are equally reimbursed to the patient whether or not she is referred to the specialist by her regular GP. Specialists were not affected by the reform, regardless of their affiliation status with the public health insurance, i.e. of whether they belong to sector 1 or sector

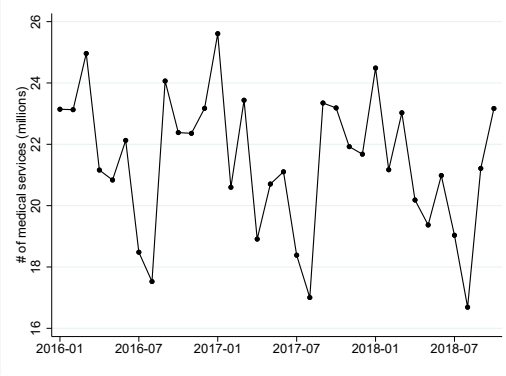
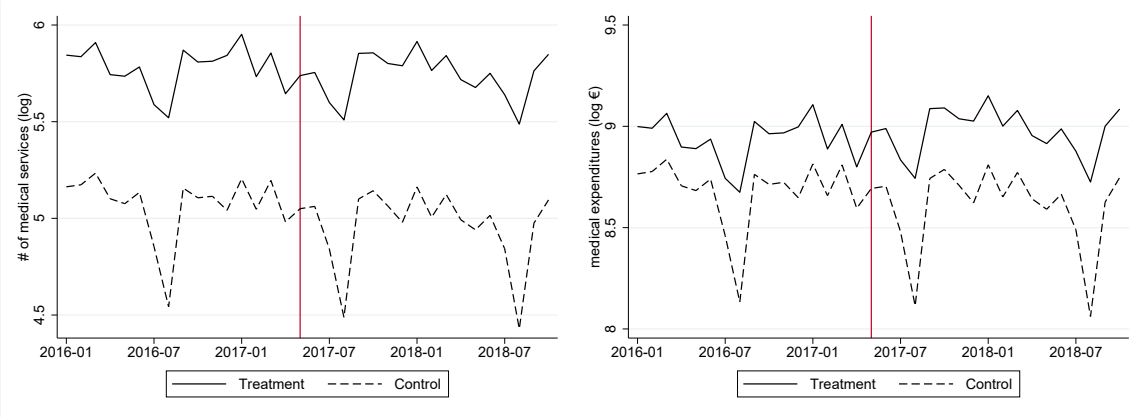


Figure 4: Evolution of GPs' activity



(a) Activity (log)

(b) Visit expenditures (log €)

Figure 5: Evolution of primary outcomes (T: Sector 1 GPs, C: Sector 2 GPs)

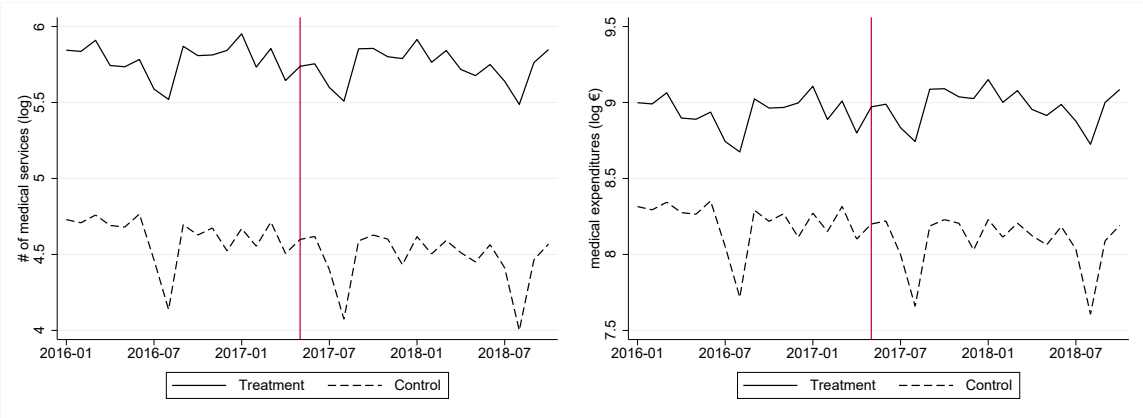
2.[7]

An objection to our main identification strategy could lie in a possible violation of SUTVA due to sector 2 GPs being affected in turn by the reform, for instance because they compete in the same market as sector 1 GPs. First, this assertion is debatable because the population of patients tends to be rather segmented in this respect, wealthier patients primarily going to sector 2 GPs. Our comprehensive data enable us to compute

[7]Salaried physicians might also be considered as a comparison group, since they were not concerned either by the regulatory change at stake. However, we found that the CTA was unlikely to hold in that case. Another identification strategy could have consisted in allowing for the whole universe of self-employed specialists to be a comparison group, but again the CTA seemed less plausible to us.

the number of patients going to a sector 2 GP among the pool of patients who visit at least a sector 1 GP. In 2016, that share does not exceed 0.8%. Second, and to empirically assess the plausibility of SUTVA in our context, we wonder in section 5.2.2 whether sector 2 GPs could have been affected by this regulatory change, relying on direct access specialists as a comparison group, and considering these sector 2 GPs as a 'treated' group. This falsification test enables us to verify that they did not seem to react to the reform, hereby comforting the empirical plausibility of SUTVA and the validity of our primary approach.

To further alleviate concerns about the comparability of groups, either sector 1 and sector 2 GPs, on the one hand, or sector 1 GPs and direct access specialists, on the other hand, and to challenge the likelihood of the CTA, we use synthetic DiD (Arkhangelsky et al., 2021) that build upon synthetic control methods (Abadie and Gardeazabal, 2003).



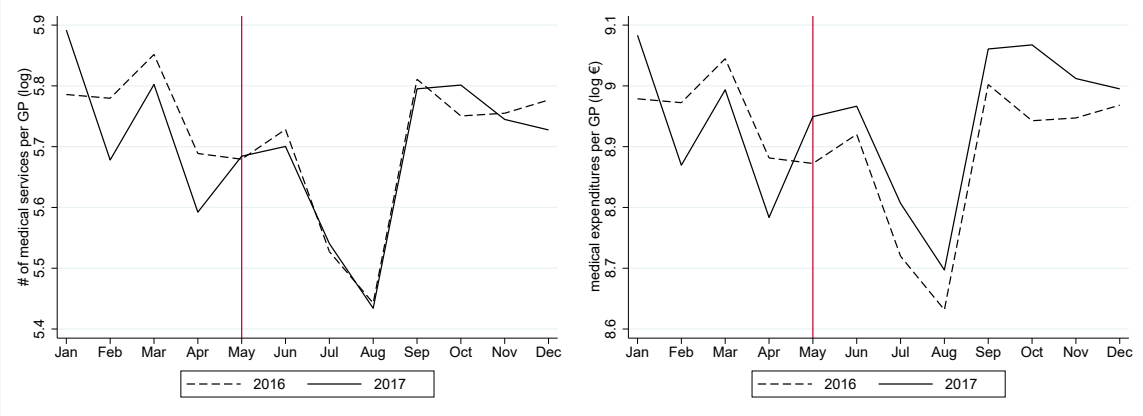
(a) Activity (log)

(b) Visit expenditures (log €)

Figure 6: Evolution of primary outcomes (T: Sector 1 GPs, C: Direct access specialists)

Last, we adopt another identification strategy in which we restrict our attention to the sole sector 1 GPs. We now view the year 2017 as a 'treated' year and the year 2016 as a 'baseline' year', so that the comparison between both years permits to account for seasonality. Such a methodology has been recently used by, e.g., Brandily et al. (2021), Andersen et al. (2022), and Bonnet et al. (2022) to evaluate the impact of the COVID-19 pandemics on various outcomes (mortality, earnings, consumption, and savings, among others). It rests on the assumption that sector 1 GPs face identical monthly seasonality in 2016 and 2017. The empirical plausibility of that hypothesis can be assessed from pre-

event data, as is the case with the CTA. Figure 7 suggests that this assumption is not verified when January is included, but that it is likely based on the sole February and March months. This Figure makes clear that sector 1 GPs have increased their activity in response to the public policy change: the level of medical services provided the months after that change is identical in 2016 and 2017, while it is lower on February and March 2016 by comparison with February and March 2017. The same diagnosis prevails for medical expenditures. A short-term anticipation effect can also be observed in April 2017 where we see a drop in the activity of sector 1 GPs, followed by some spike in May 2017, which suggests that physicians postponed some visits from April to May in order to take advantage from higher regulated fees. This adjustment behavior is all the more likely for elective medical services since doctors were clearly incentivized to postpone non-emergency encounters so that they could benefit from the extra €2 per visit. In all subsequent estimations, our preferred specification is obtained when we exclude this 2-month window from our sample, even though we show that including it or not does not dramatically alter our results. In a similar vein, we check that our estimates are not driven by the sole extensive margin, i.e. by the mere fact that some physicians might have been discouraged to exit, or incentivized to enter the healthcare market. To that aim, in some of our estimations we impose that physicians be present in both January 2016 and October 2018.



(a) Activity (log)

(b) Visit expenditures (log €)

Figure 7: Evolution of primary outcomes (T: Sector 1 GPs in 2017, C: Sector 1 GPs in 2016)

4.2 Econometric specification

Denoting various outcomes (e.g., physician activity, medical expenditures, drug prescription, etc.) by Y, and indexing doctors by j and time available at a monthly frequency by t, we consider the following linear model:

$$Y_{jt} = \beta \; T_j \times \text{Post}_t + \alpha_j + \gamma_t + \varepsilon_{jt}, \tag{1}$$

where T_j is a binary indicator for a doctor j belonging to the treatment group, Post_t is a dummy equal to 1 after May 2017, α_j is a physician fixed effect, γ_t are month-year fixed effects, and ε_{jt} is an idiosyncratic shock. Standard errors are clustered at the physician level. The exogeneity of T_j results from the change in the fee per visit being decided by the policy-maker. In this context, our parameter of interest is the coefficient β which has the usual interpretation of the average treatment effect (ATE). Since most of our dependent variables are expressed in natural logarithm, the parameter obtained corresponds to a semi-elasticity: it measures the response of the outcome considered to the treatment, in log points. To the extent that β is small, it approximately accounts for that relative response, in percentage points. When we investigate for the heterogeneity of treatment effects, we specify $\beta \equiv \beta(X_j)$ where X_j designate doctor j's covariates (gender, age, location, number of patients seen in 2016, etc.).

We also perform an event study analysis in which we allow for this coefficient to further vary over time, namely to be month-specific over the whole period:

$$Y_{jt} = \beta_t \times T_j + \alpha_j + \gamma_t + \varepsilon_{jt}, \tag{2}$$

where t_0 designates the time of the treatment, May 2017, viewed as the event, and where we normalize $\beta_{t_0-1} = 0$ if. Testing the null hypothesis $H_0 : \beta_h = 0, \; \forall h < t_0 - 1$ provides a statistical test for the absence of any differential pre-trend across treatment and comparison groups. The post-reform coefficients $\beta_h, \; \forall h \geq t_0$ provide us with dynamic (monthly) treatment effects. However, as mentioned by Coudin et al. (2015), a possibly group-specific seasonality suggests to interpret any potential month-by-month difference with caution. Note also that it helps explain why previous statistical test may be rejected in the data while not jeopardizing the core of our identification strategy, though. For this reason, we rather consider the above ATE as the relevant policy parameter.

Last, in the case when the treatment group corresponds to the year 2017 and our control group to the year 2016, our estimating equation writes as:

$$Y_{jmy} = \beta T_y \times \text{Post}_m + \nu_{jy} + \mu_m + \eta_{jmy}, \tag{3}$$

where $m = 1, \ldots, 12$ accounts for a given month of a year, Post_m is a dummy equal to 1 after May, and y is the year considered, either 2016 or 2017, so that $T_y = \mathbb{1}\{y = 2017\}$. Depending on whether one wants to retrieve static or dynamic treatment effects, one shall allow the parameter β to be month-specific or not. In a different vein, it is possible to specify either doctor ν_j or doctor-year ν_{jy} fixed effects.

5 Results

We now provide an evaluation of the previous regulatory change on various outcomes: activity (measured in medical services), its margins (activity per patient, and number of patients), visit expenditures, drug prescription and total expenditures, workload (measured in working days), and the probability of retirement (somehow the extensive margin). For each outcome considered, the empirical plausibility of the CTA can be assessed by considering the respective evolutions in both treatment and comparison groups, which we plot in Appendix Figures B1, B2, and B3. We emphasize that, within the causal inference framework exposed previously, our granular and comprehensive data permit to obtain precise estimations of the treatment effects considered.

5.1 Main results

Table 1 displays the results issued from our main identification strategy based on the comparison between sector 1 and sector 2 GPs. The estimated ATE, which can approximately be interpreted as the price semi-elasticity of physician activity, is comprised between 8 and 8.8, depending on whether the 'anticipation window' of April and May 2017 is excluded, and whether our attention is restricted to the sole physicians present in both January 2016 and October 2018. Following the 8.7% increase in fees per visit, sector 1 GPs have increased their activity, measured by the number of medical services provided, by 9,2% in our preferred specification (column 2). This empirical finding points out to a price elasticity of activity that is slightly above 1. As a consequence, visit expenditures rose by about 14%

(Table 2). In other words, when the public health insurance considers increasing physician regulated fees by 1%, it should almost forecast a 2% rise in visits' reimbursement rates. This result is certainly important from a policy viewpoint. In particular, it contrasts with the one previously obtained by Coudin et al. (2015) in France, which pointed out to a negative elasticity, even lower than -1,[8]. It is yet closer to Clemens and Gottlieb (2014) who found a medium-run price elasticity of 1.5 for US doctors in charge of Medicare patients. It should perhaps not be surprising that more attention is paid to older patients, and that the reaction is even stronger in that case, but we insist on the facts that our finding holds for the whole universe of both doctors and patients, and that our identification has a more substantial external validity than, e.g., in Coudin et al. (2015) who relied on a specific sample of compliers.

Table 1: Activity (DiD, C: Sector 2 GPs)

	(1)	(2)	(3)
Post × Treatment	0.0843***	0.0879***	0.0791***
	(0.00464)	(0.00488)	(0.00462)
Physician FE	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes
Excluding April and May 2017	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes
Observations	1849822	1740498	1489500
R^2	0.876	0.874	0.873

Note. Dependent variable: Number of medical services per GP and per month (log).
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$
Clustered standard errors at the physician level

Interestingly, the decomposition of previous effect into two margins, healthcare accessibility (proxied by the number of patients) and intensity of care (proxied by the number of services per patient) of Table 3 reveals that concerned doctors have mostly increased the number of patients seen each month (the magnitude of the effect, 8.7 log-points, being almost the same as the effect on activity), but that the number of medical services provided per patient hardly varied (the effect is positive and statistically significant at usual levels,

[8]In this article, the authors found that activity rose by 41% following a 54% price decrease.

Table 2: Visit expenditures (DiD, C: Sector 2 GPs)

	(1)	(2)	(3)
Post × Treatment	0.131***	0.134***	0.125***
	(0.00472)	(0.00496)	(0.00469)
Physician FE	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes
Excluding April and May 2017	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes
Observations	1849725	1740403	1489481
R^2	0.867	0.865	0.863

Note. Dependent variable: Medical visit expenditures per GP and per month (log €).

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

but small from an economic viewpoint, barely exceeding 0.1 log-point). The latter finding suggests that supply-induced demand mechanisms have been limited, and hence that the quality of care was not substantially altered by the regulatory change at stake. To further dig into the mechanisms that led them to admit more patients, Table 4 shows that sector 1 GPs worked more following the reform, namely 0.4 extra day per month, or 4.9 extra days per year, hence a relative increase of 2% only. It should thus be inferred from this result that the remaining part, about 6%, stems from a higher number of patients seen per workday. Since it is unlikely that doctors substantially worked longer hours, it may be the case that they have adjusted their practice through a shorter visit duration.

Table 3: Number of patients seen and intensity of care (DiD, C: Sector 2 GPs)

Dependent variable	# of patients per month (log)			# of medical services per patient (log)		
Post × Treatment	0.0821***	0.0855***	0.0771***	0.00222***	0.00239***	0.00199***
	(0.00456)	(0.00479)	(0.00456)	(0.000638)	(0.000677)	(0.000690)
Physician FE	Yes	Yes	Yes	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Excluding April and May 2017	No	Yes	Yes	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes	No	No	Yes
Observations	1849822	1740498	1489500	1849822	1740498	1489500
R^2	0.880	0.878	0.878	0.808	0.806	0.820

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

We next wonder whether physicians modified their prescription behavior following the

Table 4: Workdays (DiD, C: Sector 2 GPs)

Dependent variable	# of workdays (per month)			# of workdays (per year)		
Post × Treatment	0.411***	0.426***	0.388***	4.461***	4.695***	3.140***
	(0.0270)	(0.0284)	(0.0283)	(0.310)	(0.330)	(0.295)
Baseline	[19.7]	[19.9]	[20]	[231]	[232]	[240]
Physician FE	Yes	Yes	Yes	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Excluding April and May 2017	No	Yes	Yes	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes	No	No	Yes
Observations	1849822	1740498	1489500	1849822	1740498	1489500
R^2	0.704	0.702	0.679	0.901	0.900	0.941

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

Note. Baseline computed in treatment group before reform.

reform. Table 5 indicates that the average amount of prescribed drugs per patient has decreased by about 4.3%, which would then be consistent with a positive correlation between the length of a visit and durg prescription in our context. The reduction in drug prescription was particularly marked for antibiotics (-8.4%), but less for antidepressants (-2.3%) and opioids (-3.4%). The sum of medical visit expenditures and drug prescription has diminished by 0.7%, per patient (Table D1), suggesting that raising doctors' fee per visit is a cost-effective public policy to promote the accessibility of healthcare with limited adverse effects on the intensity of care.

Table 5: Drug prescription per patient (DiD, C: Sector 2 GPs)

Dependent variable	All drugs (log €)	Antibiotics (log €)	Antidepressants (log €)	Opioids (log €)
Post × Treatment	-0.0440***	-0.0899***	-0.0295***	-0.0367***
	(0.00514)	(0.00602)	(0.00646)	(0.00872)
Physician FE	Yes	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes	Yes
Observations	1685157	1630066	1570041	1586775
R^2	0.853	0.736	0.818	0.773

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

We further investigate whether this regulatory change induced doctors to postpone their retirement.[9] Based on a linear probability model, we find that this is actually the

[9] We measure possible retirements as the fact of providing no medical service from a given month onward during our observation period, and when aged 60 or more. By definition, the probability of a possible retirement increases over time.

case (Table 6), but that the impact is not substantial. Our estimates range from a -0.9% reduction in the probability of possible retirement to zero, depending on the identification strategy adopted, namely on the comparison group chosen. Yet our preferred specification is the one based on sector 1 GPs in 2016 as the relevant comparison group, because the CTA is more likely in that case than with the other possible comparison groups (Figures B1, B2, and B3 in Appendix).

Table 6: (cumulated) probability of retirement (over time)

Comparison group	Sector 2 GPs	Direct access specialists	Sector 1 GPs in 2016
Post × Treatment	-0.00911***	0.00913	0.000547
	(0.00215)	(0.00599)	(0.00603)
Physician FE	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes
Observations	326260	632458	632996
R^2	0.657	0.650	0.649

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

Last, we check whether the composition of patients going to a sector 1 GP was substantially altered by the reform. Preliminary evidence suggests that the share of patients aged 65 or more has slightly decreased, which thus indicates that new patients seen are younger.[10] For each doctor, we also compute the share of patients that she sees as a regular GP, i.e. as a gatekeeper.[11] Table 7 shows that, after the regulatory change at stake, this share has increased more rapidly, by 5.8%, i.e. by about +2.9 pp, among sector 1 than among sector 2 GPs. This result suggests that not only do sector 1 GPs see more patients, but also that they admit new patients as their regular GPs. It even indicates that the probability of be(com)ing the regular GP of a new patient is higher than it was for a former patient. As regards the evolution of travel time between patients and doctors, we find no conclusive evidence (Table D2). On the one hand, the average travel time has significantly increased by 0.45 minute, and the variance is also higher. On the other hand, most of the

[10] Corresponding estimates are available upon request.

[11] In the French healthcare system, visiting a specialist without being referred from one's regular GP to that specialist is (more) costly, for instance. By definition, patients visiting direct access specialists without a referral are not subject to such overcharges.

distribution of these travel times (say, between the bottom 10% and the top 10%) remained quite unchanged. All together, it suggests that the difference is mainly due to outliers in this dimension, i.e. patients who visit a doctor located far away from their home. Such visits may correspond to those happening during holidays since it is a time when patients frequently travel far away from their homes. In that sense, the accessibility of healthcare has not improved, when measured by shorter travel times. Rather, new patients seem to have now access to healthcare, while being located in similar areas as former patients.

Table 7: share of patients that the GP sees as the regular GP (DiD, C: Sector 2 GPs)

	(1)	(2)	(3)
Post × Treatment	0.0276***	0.0289***	0.0279***
	(0.000813)	(0.000862)	(0.000879)
Physician FE	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes
Excluding April and May 2017	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes
Observations	1849822	1740498	1489500
R^2	0.941	0.940	0.942

Note. Dependent variable: Share of patients that the GP sees as the regular GP.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

5.2 Threats to identification and sensitivity analysis

In this subsection, we perform a few robustness checks. We first verify that our results remain under various identifying assumptions, namely CTAs, based on alternate comparison groups. We next assess the empirical plausibility of SUTVA as regards our main identification strategy, relying on sector 2 GPs as our favorite comparison group, by wondering whether they have been affected by the regulatory change. We also simulate a fake reform as a falsification test in order to gain confidence over causality. Last, we use a synthetic DiD approach to alleviate concerns about possible violation of CTAs in each of our three identification strategies.

5.2.1 Alternate comparison groups

We here implement two other identification strategies to assess the robustness of previous empirical findings. Reassuringly, the results obtained under alternate corresponding assumptions concur with previous evidence: Tables 8 and 9 show that when comparing sector 1 GPs with direct access specialists[12] and themselves the year before, respectively, we still obtain that these GPs have increased their activity by between 7 and 9.9%, which, again, and given the +8.7% price increase, points out to an almost unitary price elasticity. As a result, doctors' visit fees rose by 16 to 18%.

Table 8: Activity and visit expenditures (DiD, C: Direct access specialists)

Dependent variable	# of medical services (log)			medical visit expenditures (log €)		
Post × Treatment	0.0903***	0.0963***	0.0879***	0.162***	0.168***	0.159***
	(0.00540)	(0.00572)	(0.00583)	(0.00554)	(0.00587)	(0.00600)
Physician FE	Yes	Yes	Yes	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes	Yes	Yes	Yes
Excluding April and May 2017	No	Yes	Yes	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes	No	No	Yes
Observations	1918543	1805033	1538620	1918376	1804867	1538569
R^2	0.886	0.884	0.885	0.875	0.873	0.872

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

Note also that the 'seasonality-based' approach, based on the comparison among sector 1 GPs between 2016 and 2017, yields much more precise estimates despite the lower sample size that results from the restriction to the sole GPs. This is due to the fact that this identification strategy exploits the within-physician variation by comparing the same doctor's outcomes over time, as opposed to the two other strategies that rely on between-physician comparisons.

5.2.2 SUTVA

Coming back to our main identification strategy based on sector 2 GPs as our comparison group, we now seek to empirically assess the plausibility of SUTVA, i.e. of the absence of response of that group to the policy change at stake. To do so, we rely on the group direct

[12] Another robustness check would consist in isolating a single specialty, namely gynecology, ophthalmology, and stomatology, as a comparison group. Though these approaches result in lower statistical power, each of these comparison groups may be viewed as more homogeneous than the whole set of direct access specialists. Corresponding estimates remain fairly unchanged, which indicates that previous findings are not driven by a single group of specialists. They are available upon request.

Table 9: Activity and visit expenditures (DiD, C: Sector 1 GPs in 2016)

Dependent variable	# of medical services (log)			medical visit expenditures (log €)		
Post × Treatment	0.0775***	0.0686***	0.0606***	0.157***	0.149***	0.141***
	(0.00140)	(0.00165)	(0.00137)	(0.00141)	(0.00165)	(0.00137)
Physician-Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Month FE	Yes	Yes	Yes	Yes	Yes	Yes
Excluding April and May 2017	No	Yes	Yes	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes	No	No	Yes
Observations	1104099	1003376	859165	1104046	1003325	859156
R^2	0.896	0.893	0.890	0.893	0.891	0.887

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

access specialists as a comparison group for which the SUTVA is perhaps more likely. We next wonder whether sector 2 GPs, here viewed as a treatment group, have modified their behavior following the reform. Table 10 indicates reassuringly that this is not the case, the effect being not significant at the 5% level. This empirical evidence hence concurs to comfort our previous approach, according to which sector 2 GPs are a valid comparison group.

Table 10: The plausibility of SUTVA (DiD, T: Sector 2 GPs, C: Direct access specialists)

	(1)	(2)	(3)
Post × Treatment	0.00524	0.00781	0.00825
	(0.00697)	(0.00736)	(0.00729)
Physician FE	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes
Excluding April and May 2017	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes
Observations	358181	336778	276922
R^2	0.875	0.873	0.874

Note. Dependent variable: Number of medical services per GP and per month (log).

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

5.2.3 Placebo experiment

We next perform a falsification test in which we simulate a fake reform occurring in May 2016, i.e. one year before the actual regulatory change. The corresponding results are

displayed by Table 11 and point out to no significant effect in that case, as expected.

Table 11: Falsification test: Fake reform in May 2016 (DiD, C: Sector 2 GPs)

	(1)	(2)	(3)
Post × Treatment	-0.00747*	0.0000575	0.000603
	(0.00389)	(0.00414)	(0.00394)
Physician FE	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes
Excluding April and May 2016	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes
Observations	1849822	1739099	1489574
R^2	0.875	0.875	0.875

Note. Dependent variable: Number of medical services per GP and per month (log).

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

5.2.4 Synthetic DiD

Last, we use a synthetic difference-in-differences approach (Arkhangelsky et al., 2021), which points out to an average treatment effect on the treated (ATT) of 5.4 log-point (see also Figure B4) when the pool of possible controls is composed of sector 2 GPs, of 6.4 log-point (see also Figure B5) when the pool of possible controls is composed of direct access specialists, and of 4.1 log-points (see also Figure B6) when the pool of possible controls is composed of sector 1 GPs in 2016. The estimates we obtain here are unsurprisingly more conservative, and it is thus interesting to already obtain a price elasticity that is not much lower than 1/2 with this method.

5.3 Heterogeneity

Finally, we investigate how the estimated price elasticity of physician activity varies depending on observed characteristics like gender, age, or location, among others. The results are displayed by Figure 12. Female GPs seem to react more strongly to a change in financial incentives (panel a) since their estimated ATE is 0.11, as opposed to 0.08 for male GPs, which corresponds to elasticities of 1.3 and 0.9, respectively. Yet the most striking

dimension of heterogeneity is age: we find markedly dispersed behavioral responses in this regard (panel b), and the effect is monotone. While doctors aged 30 to 39 increase their activity by up to 18% following that reform, the oldest GPs, still active beyond 70, hardly change their activity at all, doctors who are 60 to 69 moderately increase theirs by about 5%, and other intermediate responses are estimated for doctors aged 50 to 59 (+9%) or 40 to 49 (+12%). Interestingly, once age has been controlled for, women and men exhibit similar responses (panel g): it should thus be concluded that the apparently gender-specific reaction was mostly driven by a composition effect, whereby female doctors are simply younger than their male colleagues. Moreover, it is likely that age acts as a proxy for wealth, or permanent income. In that sense, income effects may be small at younger ages, and dominated by substitution effects, while becoming more important at older ages due to life-cycle accumulation of wealth and (almost) equalizing substitution effects in absolute. This likely explanation helps rationalize our findings and those of the literature devoted to physician labor supply, including reconciling cases when there is an apparent discrepancy. As regards location, we find no significant heterogeneity between dense and weakly dense areas (panel c), or between medical deserts and areas that are not medical deserts (panel f). Remember that doctors practicing in non dense areas are not numerous enough to produce precise estimations. Mechanically, doctors in the bottom 10% of the number of patients seen in 2016 tend to exhibit both a more substantial and significant reaction to a change in financial incentives (panel d), since they increase their activity by up to 20%. Yet this empirical finding should be interpreted with caution: part of the effect may result from a small denominator, and a remaining part could be due to the fact that these GPs have experienced career interruptions, part-time activity, etc. The effect of the reform has been gradual over time, as made clear by panel e) and by Figure C1 in Appendix, which performs the monthly event study analysis. However, as mentioned above and in Coudin et al. (2015), the monthly activity displays a high degree of seasonality, and not too much attention should be paid on fluctuations obtained at this granular level. Another noteworthy finding relates to the effect of the exposure to competition. We measure the latter through the share of 'shoppers', in IO terms, i.e. of patients who visit another GP at least once in the year.[13] Panel h) suggests that the degree of exposure to competition has

[13]For the median GP in this dimension, the 2016 share of 'shoppers' amounts to 0.11, meaning that 89% of her patients are 'loyal'. As the intuition suggests, the share of such captive patients is higher in medical deserts.

an inverted U-shaped impact on GPs' response to financial incentives. When the degree of competition is reasonably small, i.e. below some threshold that locates at about the 9th decile of the distribution of that share, a fiercer competition with surrounding doctors tends to exacerbate the response to financial incentives. Yet one cannot reject that the top 10% of doctors facing the highest level of competition react similarly as the median doctor, and even the bottom 20% of doctors in that dimension: the most exposed to competition are significantly unincentivized by comparison with their colleagues facing lower competition, namely the ones between P80 and P90 of that distribution. *Ceteris paribus*, Table 12 confirms that the two primary drivers of the reaction to financial incentives are both the age of the GP and the presence of shoppers among her patients: in particular, early-career physicians are much more likely than old-age doctors to adjust their provision of care, while later-career physicians remain largely unaffected.

6 Conclusion

This article has empirically assessed the effects of higher financial incentives provided to GPs on various outcomes related to the accessibility of healthcare, including the provision, the intensity of care, and doctors' workload, at the intensive margin, but also on retirement behavior, at the extensive margin, as well as on medical spending and drug prescription. Based on an exogenous shock and comprehensive data on the universe of both patients and doctors in France, we find a substantial response to these incentives, especially for early-career physicians, which takes the form of higher GPs' activity. Doctors see more patients, and this adjustment seems to operate through both a higher workload, measured in days, which roughly explains 1/4 of the observed changes, and a higher number of patients per day, for the remaining 3/4. Besides, the share of their patients that they see as a regular GP increases, suggesting that these doctors admit new patients and become their regular GP. Overall, providing stronger financial incentives sounds like a cost-effective policy to enhance the accessibility of healthcare, with limited adverse effects on quality: we do not detect any change in the number of visits per patient, for instance.

These empirical results are derived from a clear-cut research design due to both the reform being wide-scale and the exhaustive nature of our data, and we thus insist on their external validity being high for these reasons. We hence believe that they should be taken

into account by the decision-maker when designing public policies for deprived medical areas. For instance, they indicate that policies targeted at newly graduates from medical schools, i.e. that seek to increase the extensive margin, may be effective tools in this regard.

A natural extension of this research would consist in determining whether any change in the quality of care has been observed following that reform, which requires the availability of relevant indicators of quality like mortality, 30-day readmission, counter-guidelines prescriptions, etc. To further complement the current analysis, it would also be interesting to document substitution or displacement effects, namely redistribution of patients from hospitals to sector 1 GPs, if any. Last, empirically assessing the response to similar shocks that intervened at the end of 2023 (resp. 2024), when the fee per visit was raised, again, from €25 to €26.5 (resp. €30), is another promising area of further research.

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Table 12: Heterogeneity of treatment effects

post_treatment	-0.0233
	(0.0242)
Age (ref: 70-79)	
30-39 × post_treatment	0.157***
	(0.00801)
40-49 × post_treatment	0.103***
	(0.00734)
50-59 × post_treatment	0.0790***
	(0.00698)
60-69 × post_treatment	0.0349***
	(0.00687)
Gender (ref: men)	
women × post_treatment	-0.00277
	(0.00253)
Density of area (ref: dense)	
intermediate × post_treatment	0.000587
	(0.00334)
weakly dense × post_treatment	0.00823**
	(0.00367)
not dense × post_treatment	0.0297*
	(0.0178)
Medical desert classification (ref: not a medical desert)	
ZIP × post_treatment	0.00201
	(0.00383)
ZAC × post_treatment	0.000565
	(0.00279)
Number of patients in 2016 (ref: Group 2)	
1 × post_treatment	0.0956***
	(0.0358)
3 × post_treatment	0.00926
	(0.0231)
4 × post_treatment	0.0194
	(0.0215)
5 × post_treatment	0.0271
	(0.0213)
6 × post_treatment	0.0188
	(0.0212)
7 × post_treatment	0.0162
	(0.0212)
8 × post_treatment	0.0134
	(0.0212)
9 × post_treatment	0.0109
	(0.0212)
10 × post_treatment	0.0104
	(0.0212)
Share of shoppers (ref: Group 2)	
1 × post_treatment	-0.00516
	(0.0140)
3 × post_treatment	0.00654*
	(0.00354)
4 × post_treatment	0.00852**
	(0.00364)
5 × post_treatment	0.00754**
	(0.00376)
6 × post_treatment	0.00861**
	(0.00393)
7 × post_treatment	0.0120***
	(0.00399)
8 × post_treatment	0.0169***
	(0.00438)
9 × post_treatment	0.0172***
	(0.00495)
10 × post_treatment	-0.00696
	(0.00734)
French *département* (ref: Paris)	
Alpes-Maritimes × post_treatment	0.0382***
	(0.0139)
Gard × post_treatment	0.0378***
	(0.0146)
Landes × post_treatment	0.0538***
	(0.0145)
Marne × post_treatment	0.0421***
	(0.0138)
Pyrénées-Atlantiques × post_treatment	0.0452***
	(0.0122)
Var × post_treatment	0.0324***
	(0.0123)
Constant	5.729***
	(0.00260)
Observations	1554661
R^2	0.868

Dependent variable: Number of medical services (log).
Standard errors in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Appendix

A Working sample: Descriptive statistics

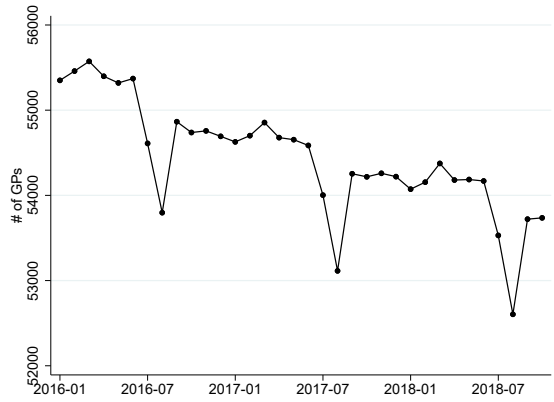


Figure A1: Number of GPs in sample

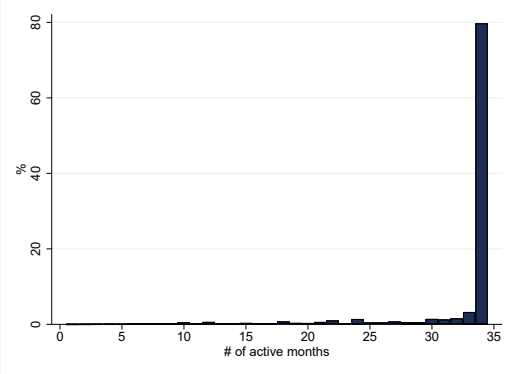


Figure A2: Number of months active in sample

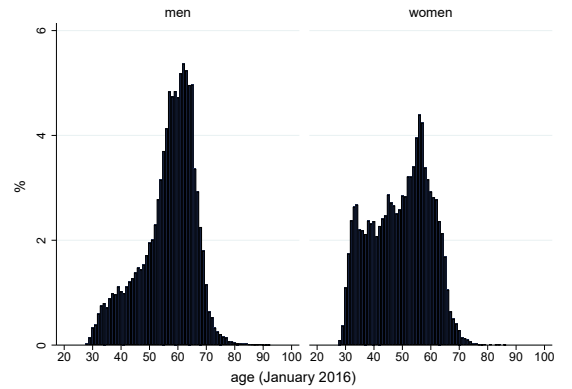
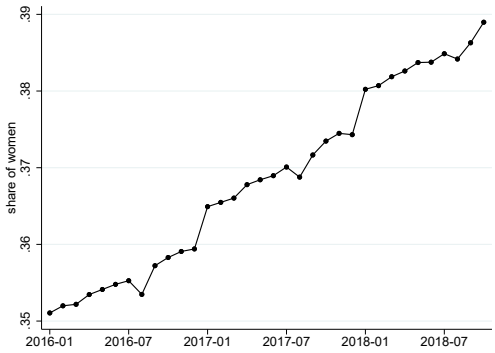
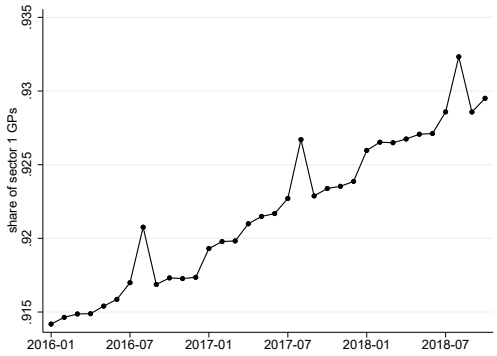


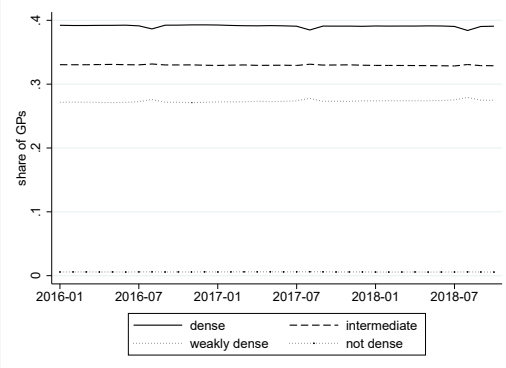
Figure A3: Age of GPs in January 2016 (before trimming)



(a) Gender



(b) Sector 1



(c) Location

Figure A4: GPs (Observed characteristics)

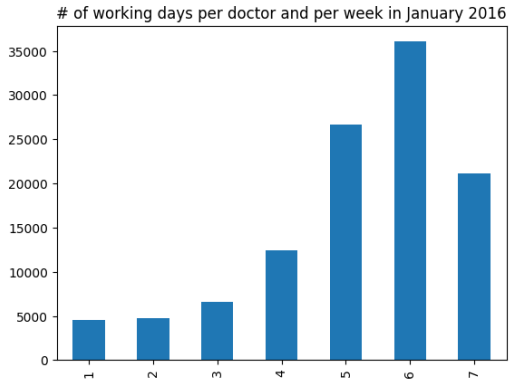


Figure A5: Number of weekly working days

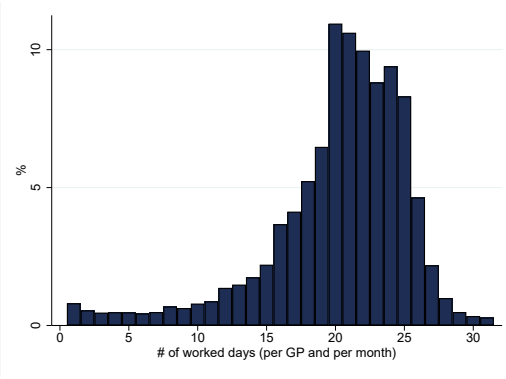
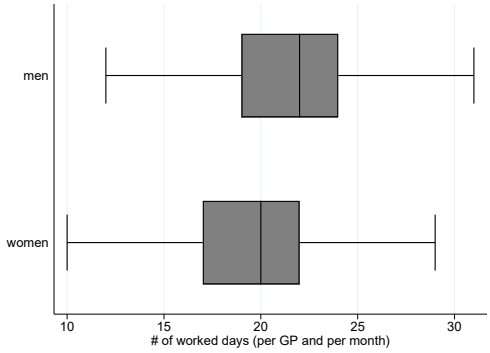
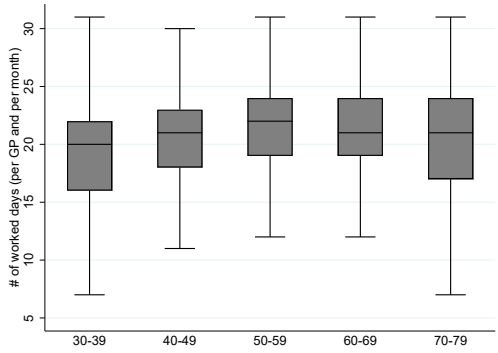


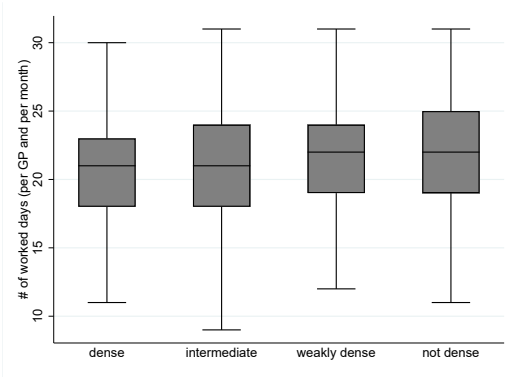
Figure A6: Number of working days (per GP and per month)



(a) Gender

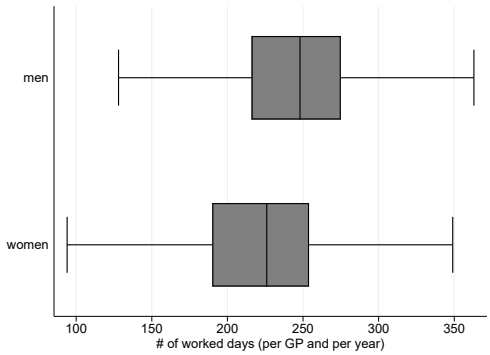


(b) Age

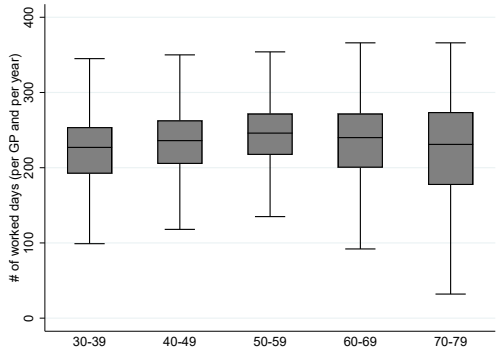


(c) Location

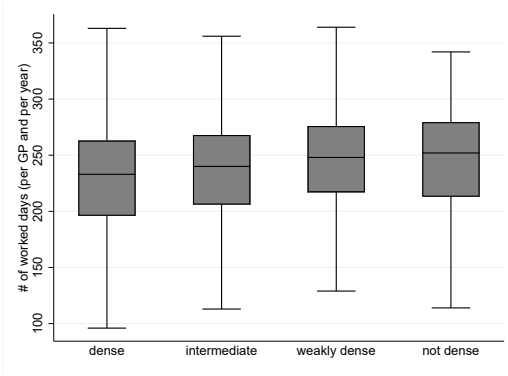
Figure A7: Number of working days (per GP and per month)



(a) Gender

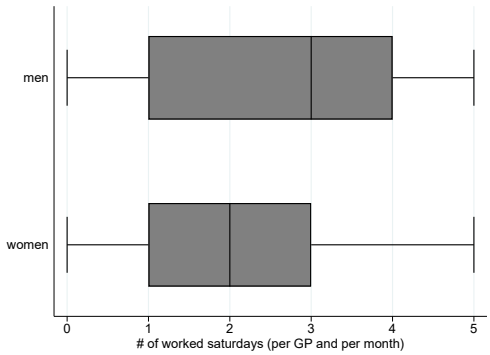


(b) Age

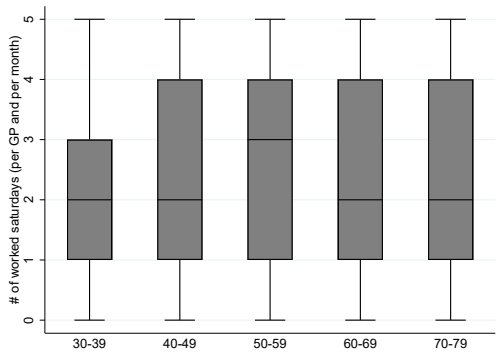


(c) Location

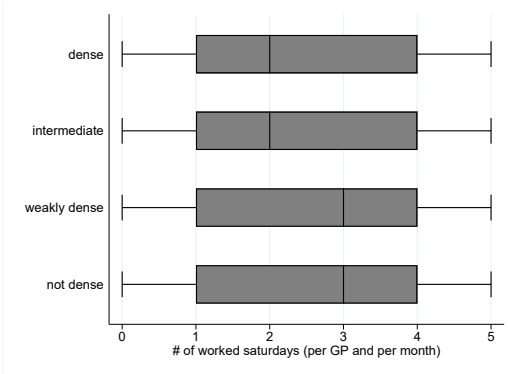
Figure A8: Number of working days (per GP and per year)



(a) Gender



(b) Age



(c) Location

Figure A9: Number of working Saturdays (per GP and per month)

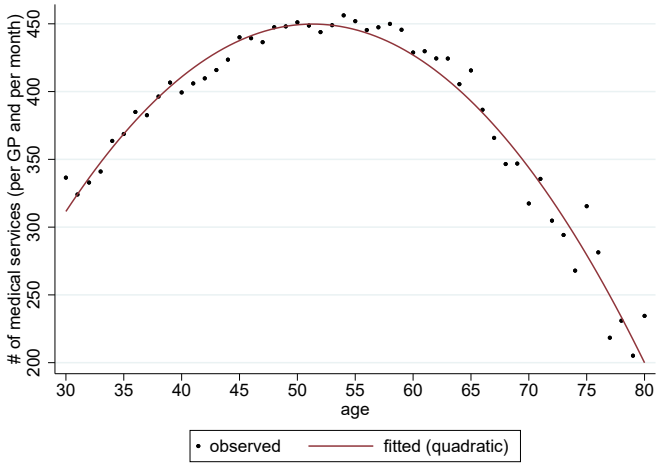


Figure A10: Number of medical services (per GP and per month), depending on age)

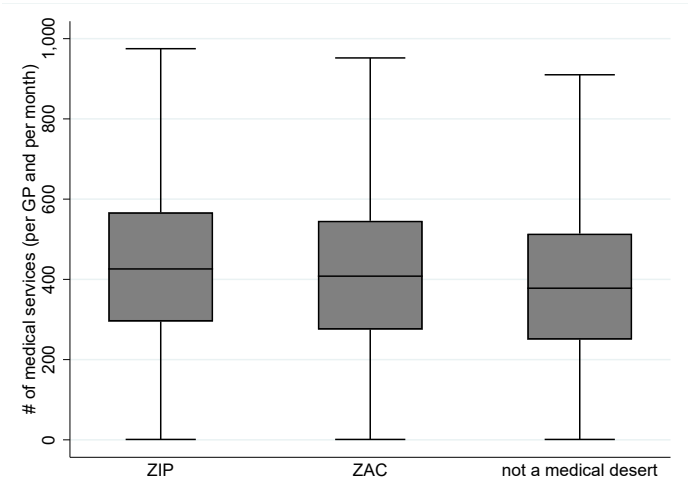


Figure A11: Number of medical services (per GP and per month) in medical deserts

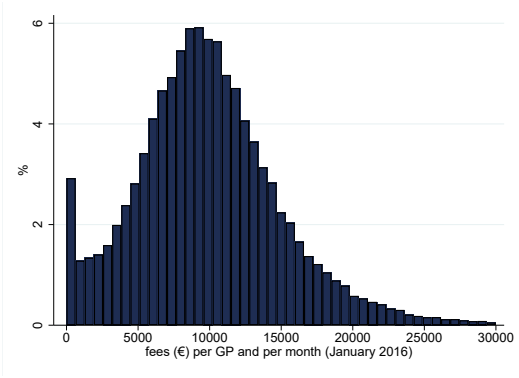
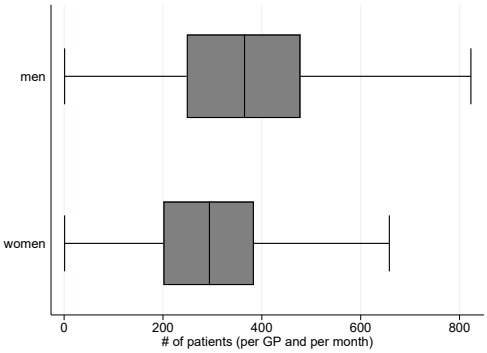
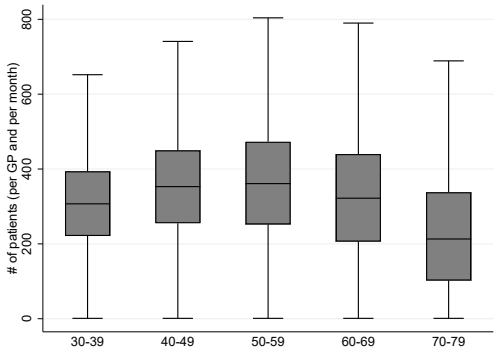


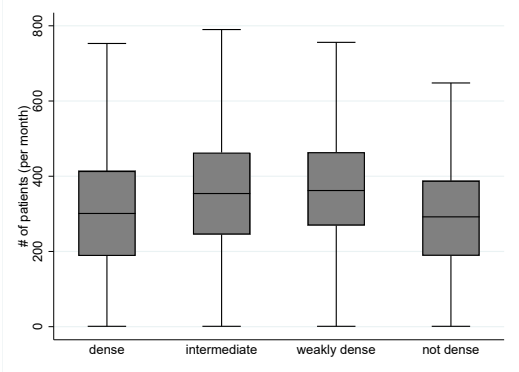
Figure A12: Medical visit expenditures (per GP and per month)



(a) Gender

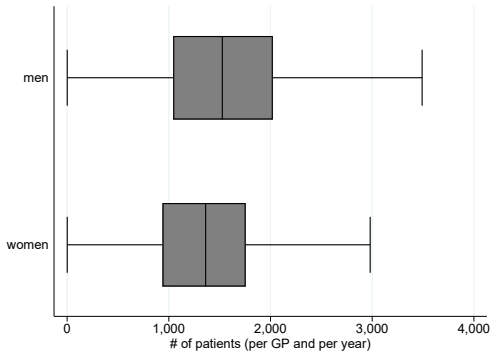


(b) Age

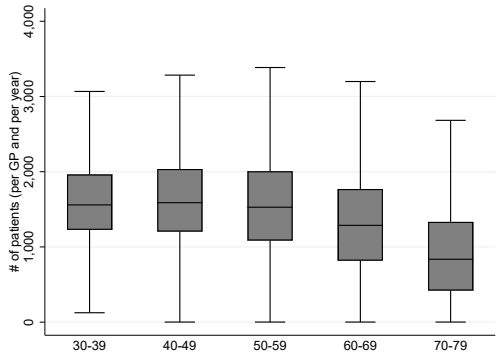


(c) Location

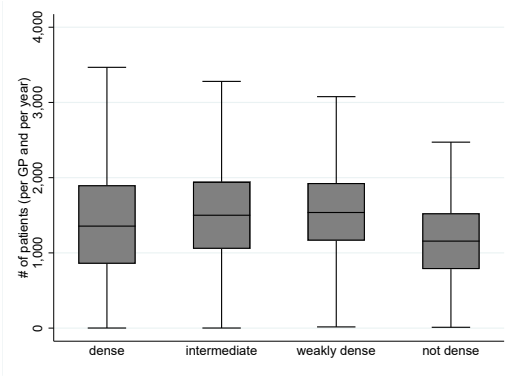
Figure A13: Number of patients (per GP and per month)



(a) Gender

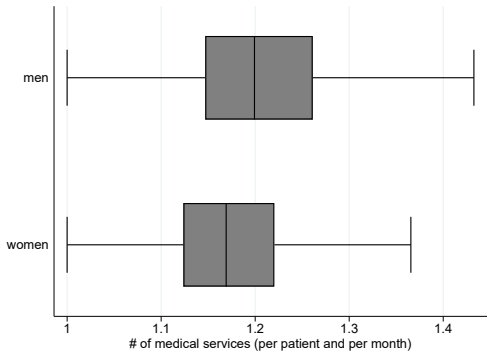


(b) Age

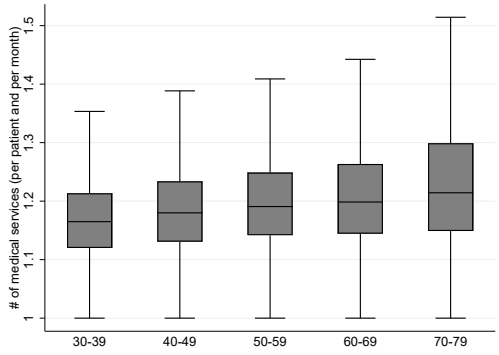


(c) Location

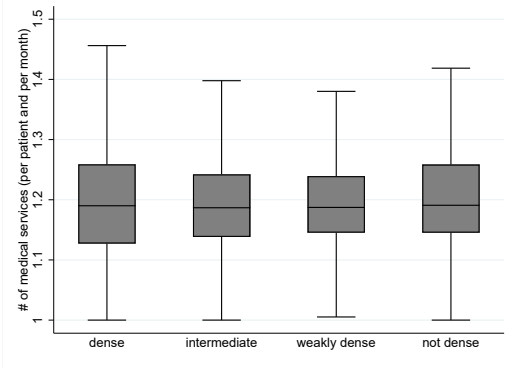
Figure A14: Number of patients (per GP and per year)



(a) Gender



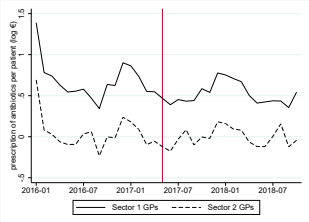
(b) Age



(c) Location

Figure A15: Number of medical services per patient

B Threats to identification



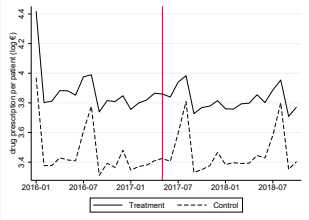
(a) Prescription of antibiotics per patient (log €)



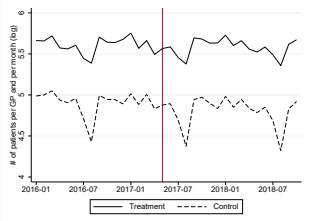
(b) Prescription of antidepressants per patient (log €)



(c) Prescription of opioids per patient (log €)



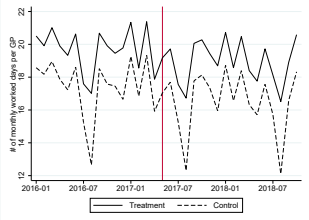
(d) Prescription of drugs per patient (log €)



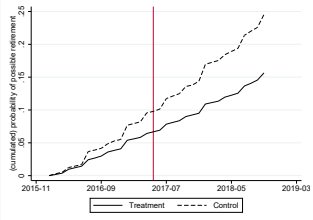
(e) Number of patients per GP and per month



(f) Number of medical services per patient

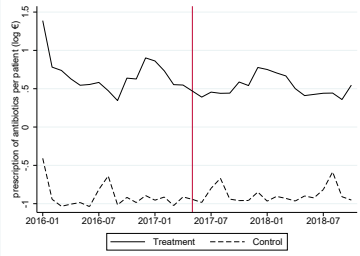


(g) Number of working days per GP and per month

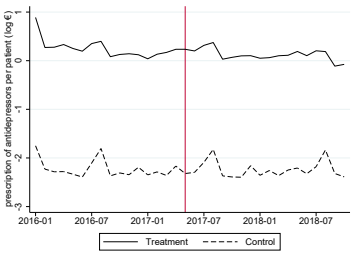


(h) Probability of (possible) retirement

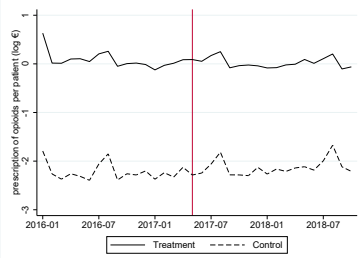
Figure B1: Evolution of outcomes (C: Sector 2 GPs)



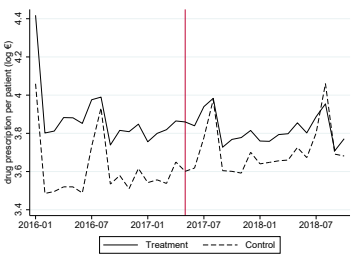
(a) Prescription of antibiotics per patient (log €)



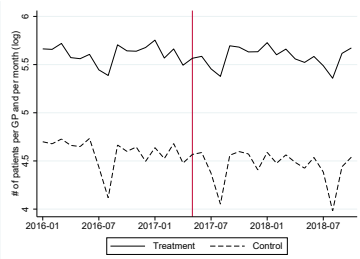
(b) Prescription of antidepressants per patient (log €)



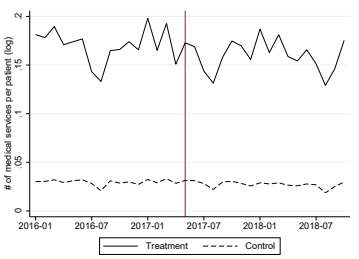
(c) Prescription of opioids per patient (log €)



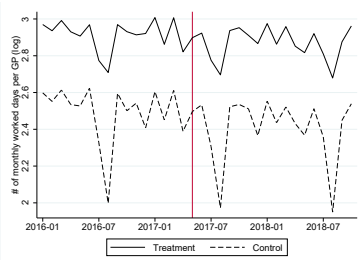
(d) Prescription of drugs per patient (log €)



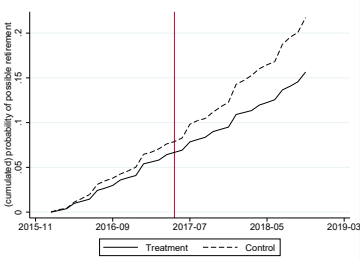
(e) Number of patients per GP and per month



(f) Number of medical services per patient

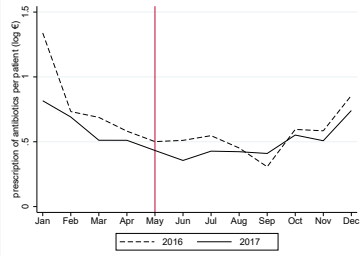


(g) Number of working days per GP and per month

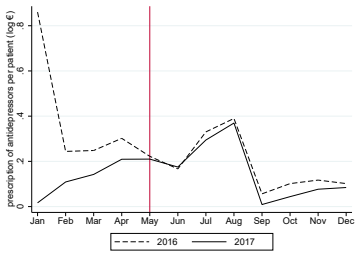


(h) Probability of (possible) retirement

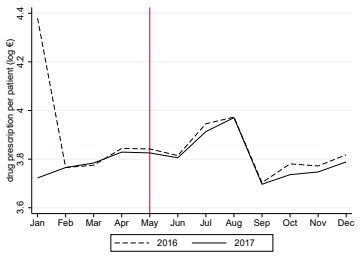
Figure B2: Evolution of outcomes (C: Direct access specialists)



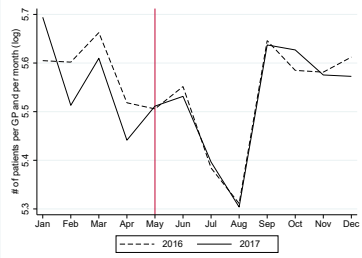
(a) Prescription of antibiotics per patient (log €)



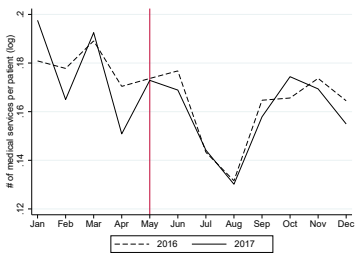
(b) Prescription of antidepressants per patient (log €)



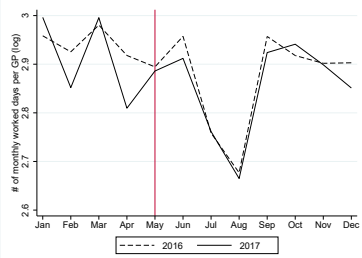
(c) Prescription of drugs per patient (log €)



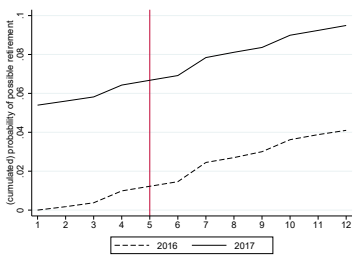
(d) Number of patients per GP and per month



(e) Number of medical services per patient



(f) Number of working days per GP and per month



(g) Probability of (possible) retirement

Figure B3: Evolution of outcomes (C: Sector 1 GPs in 2016)

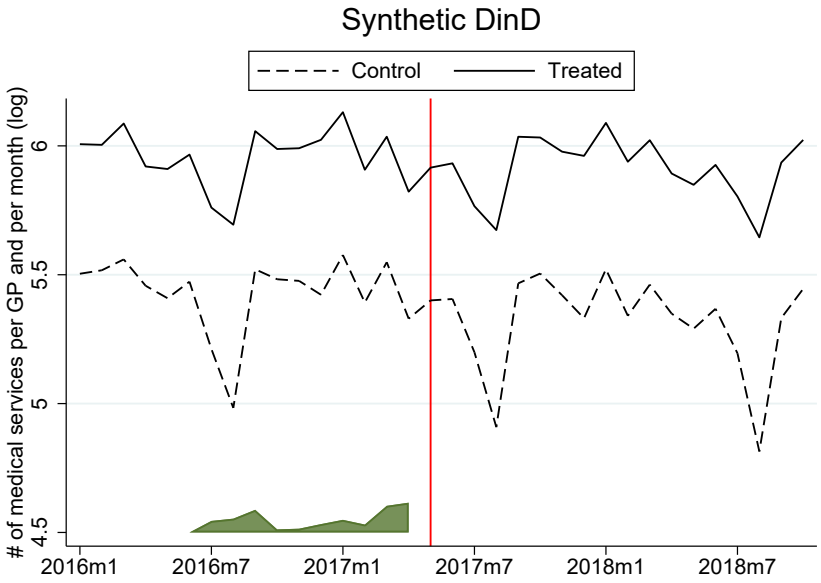


Figure B4: Synthetic DiD (T: Sector 1 GPs, C: Sector 2 GPs)

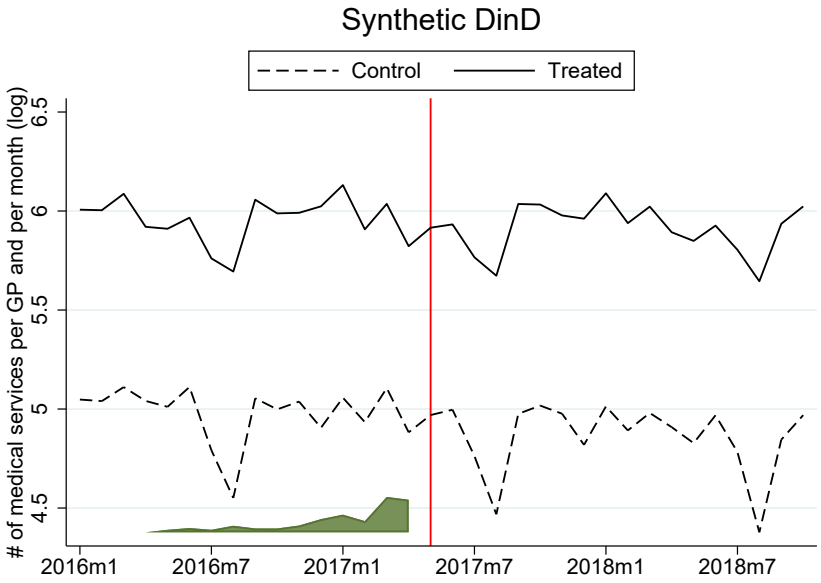


Figure B5: Synthetic DiD (T: Sector 1 GPs, C: Direct access specialists GPs)

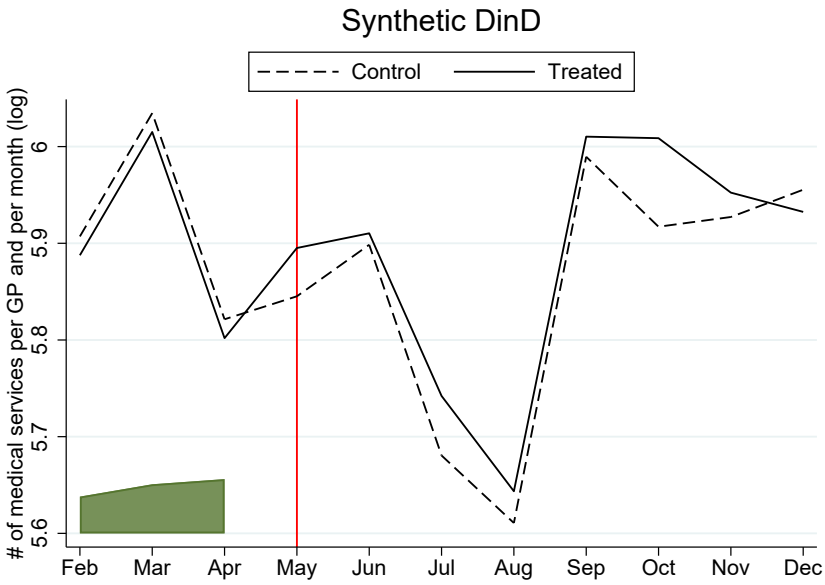


Figure B6: Synthetic DiD (T: Sector 1 GPs in 2017, C: Sector 1 GPs in 2016)

C Event study analysis

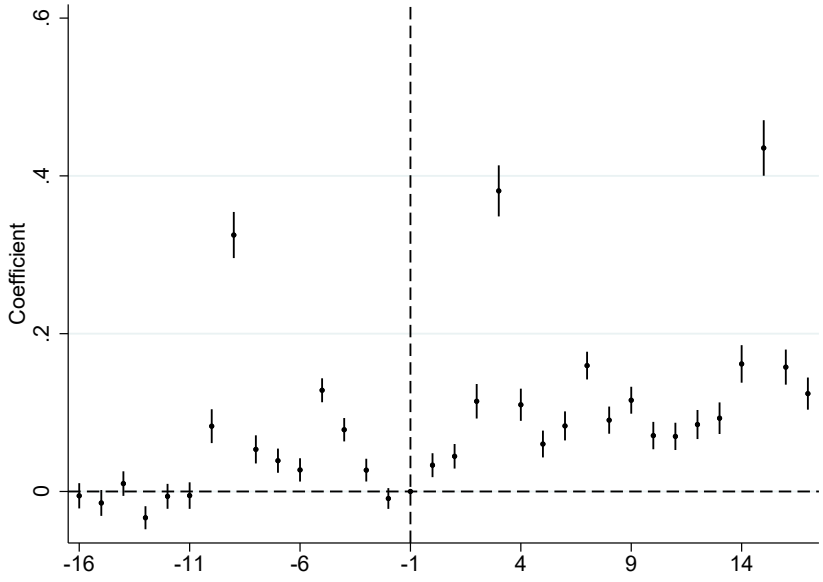


Figure C1: DiD estimates of dynamic TE (C: Sector 2 GPs)

D Other outcomes

Table D1: Total expenditures per patient (DiD, C: Sector 2 GPs)

	(1)	(2)	(3)
Post × Treatment	-0.00380	-0.00402	-0.00607**
	(0.00264)	(0.00281)	(0.00270)
Physician FE	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes
Excluding April and May 2017	No	Yes	Yes
Present in January 2016 and October 2018	No	No	Yes
Observations	1630292	1520978	1302660
R^2	0.770	0.767	0.774

Note. Dependent variable: Medical expenditures (doctor visits and drugs) per patient (log €).

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

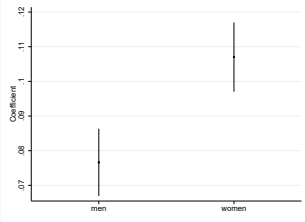
Table D2: Travel time (in minutes) between patients and GP (DiD, C: Sector 2 GPs)

Dependent variable	mean	sd	p10	p25	p50	p75	p90
Post × Treatment	0.453***	1.402***	-0.0446	0.181*	0.219	-0.0754	0.369
	(0.173)	(0.267)	(0.0833)	(0.0968)	(0.154)	(0.317)	(0.606)
Physician FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Month-year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1622173	1612357	1622173	1622173	1622173	1622173	1622173
R^2	0.760	0.678	0.503	0.592	0.593	0.639	0.702

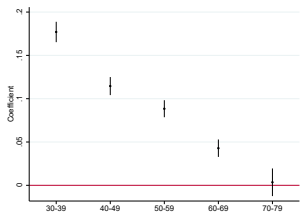
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Clustered standard errors at the physician level

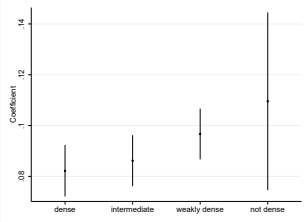
E Heterogeneity of treatment effects



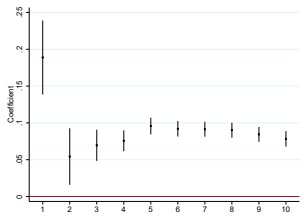
(a) Gender



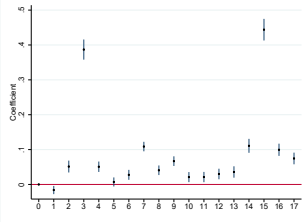
(b) Age



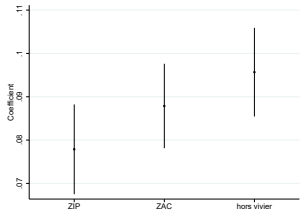
(c) Location



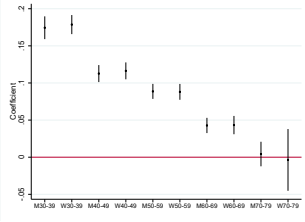
(d) # of patients (Decile group in 2016)



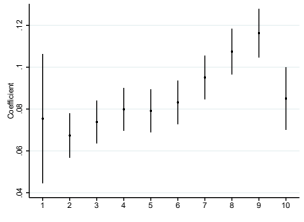
(e) Months elapsed since the reform



(f) Medical deserts



(g) Age × gender



(h) Competition (Decile group of the share of multiple-source patients)

Figure E1: Heterogeneity of treatment effects