Paid Sick Leave and Employee Absences

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Abstract. This paper studies the response of sickness absences to changes in the replacement rate for sick leave. In June 2008, a national law modified both the strength of monitoring and the monetary cost of sick leaves for public sector employees in Italy. Focusing on the National Health Service, which accounts for 21 per cent of the Italian public administration, first, I show that absences largely decreased following the reform. Second, using a difference-in-differences strategy that exploits variation in changes to the replacement rate for sick leave, I estimate that a 1 per cent point decrease in the replacement rate reduces absences by 1 per cent.

1. Introduction

Absenteeism is higher among public employees relative to their private sector counterparts (D'Amuri, 2017). The high levels of absenteeism have long been considered a plague of the public sector and one of the determinants of low productivity and high labor costs, in Italy as well as in other European countries (Lusinyan and Bonato, 2007) and in developing countries (Banerjee and Duflo, 2006). Nonetheless, the literature on absenteeism mostly focuses on the private sector.

This paper studies the effects of a change in sick leave policy that took place in 2008 in Italy and involved all public sector employees, estimated in 3.2 millions in 2012. The reform became effective in July 2008 and increased both monitoring and the monetary penalty for sickness absences.

Monitoring for sickness absences in Italy takes the form of random medical visits aiming to certify an employee's temporary inability to work. An employee on sick leave is required to be available at home for a given time window to receive the visit. The policy change increased this time window from 4 to 11 hours. However, for each sick leave episode there can only be one medical visit and the worker is discharged from the legal obligation of being at home upon receiving it. Thus, an increase in the time window can have the effect of discouraging healthy workers from taking short-term sick leaves.

As for the replacement rate, before the reform, an employee on sick leave received full compensation. The new law introduces cuts to ancillary payments and accessory

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components of the pay. Thus, the changes in the replacement rate are not homogeneous across workers but rather they vary due to across-job position differences in the compensation structure.

The focus of this study was on the National Health Service, which accounts for about 21 per cent of the total number of workers employed in the public administration. Analyzing the healthcare sector is of interest for two main reasons: (i) Doctors employed in the public sector are monitored by their own colleagues; thus, monitoring might not be as effective as in other sectors; and (ii) high absence rates in the healthcare sector have potentially very serious consequences, given the fundamental importance of promptness in providing health care.

I measure exposure to the change of the replacement rate introduced by the reform as the average realized ratio between ancillary payments and total compensation from 2001 to 2007. I limit to the pre-reform period because in the postperiod, the realized ratio is mechanically related to the absences. I then relate such measure to changes in sickness absences using a difference-in-differences strategy with variable treatment intensity. Under the assumption that monitoring has the same proportional effect on absences across workers within the same hospital, I estimate that a 1 per cent point increase in the absence penalty measure reduces absences by 1 per cent.

D'Amuri (2017) uses data from the Italian Labor Force Survey to study the effect of the same reform analyzed here on the probability of an absence. He compares public sector to private sector employees, following four subsequent monitoring changes that took place in Italy between 2008 and 2011 and concludes that both the strengthening of monitoring and the reduction in the replacement rate for sick leaves are responsible for reducing absences, with monitoring being more effective for men and monetary costs for women. This paper differs from D'Amuri (2017) along several dimensions. First, I rely on administrative data rather than on self-reported information. Second, I focus my analysis on the National Health Service, which is excluded from the analysis conducted in D'Amuri (2017).[1] Third, I use a different strategy which allows me to identify the semi-elasticity of sickness absences to the replacement rate.

De Paola *et al.* (2014) finds that the 20 per cent point average cut in the replacement rate induced by the reform reduces the incidence of sick leave claims by 8.8 per cent. This estimate is not directly comparable to the one obtained in this study because D'Amuri (2017) focuses on the impact of the reform onto the probability of claiming a sick leave in the reference week of a given quarter (irrespective of the number of days off), whereas I estimate the effect of changes in the replacement rate on the yearly average number of days off for sick leave.

De Paola *et al.* (2014) use microdata on 889 workers from a public university in Italy and identify the effect of the policy change using a regression discontinuity design. They show that the probability of an absence decreases more the bigger the reduction in the replacement rate for a given worker. However, they use a contemporaneous measure of earning losses, which is mechanically negatively related to absences in the period after the reform. Furthermore, they use data on clerical workers in a specific institution, thus raising the question as to whether their analysis extends to other public sector branches and occupations.

More in general, the economic literature on absenteeism has developed relatively recently. Brown and Sessions (1996) provide a survey of early theoretical and empirical contributions on absenteeism. Part of the literature has focused on the determinants of absenteeism (Barmby *et al.*, 1991; Ichino and Moretti, 2009; Johansson and Palme, 1996), its relation to the business cycle (Arai and Skogman Thoursie, 2005; Askildsen *et al.*, 2005; Leigh, 1985), work conditions (Ose, 2005), replacement rate (Ben Halima *et al.*, 2018), employment protection laws (Ichino and Riphahn (2005) and Olsson (2009) show that job security increases absences), group interactions (Lindbeck *et al.*, 2016), and across-countries differences (Barmby *et al.*, 2003; Lusinyan and Bonato, 2007).

This paper is more related to the branch of the literature that has studied the effects of policy changes on workers' absences: Using long time series data from Sweden, Henrekson and Persson (2004) find that less (more) generous sick leave compensation policies lead to lower (higher) levels of sick leave, and Olsson and Skogman Thoursie (2015) show that the generosity of sick leave compensation also affects the labor supply of spouses. Johansson and Palme (2005) use variation in the changes in sickness insurance across duration spells and estimate the elasticity of absence incidence to the cost of an absence to be −0.93 for men and −0.72 for women. This result is in line with the one obtained in this paper,[2] although a direct comparison cannot be made provided that the data I use do not allow me to estimate the price elasticity of the incidence of sick leave claims separately from the price elasticity of the number of days off for sick leave.

Ziebarth and Karlsson (2010) show that a 20 per cent point cut to the replacement rate for sick leaves in Germany induced a significant drop in both the incidence and the number of days off for sick leave. They estimate an elasticity of the latter to the replacement rate varying between 0.6 and 1.55, remarkably in line with the results I find. Consistently with my results and with the prediction that workers' sick leave behavior responds to monetary incentives, Dale-Olsen (2014) shows that in Norway, both male and female performance pay workers experience longer sick leaves when provided private supplementary sick pay compared with those being eligible for public sick pay only.

In designing the optimal sick leave policy governments have to take into account, the trade-off between full insurance and moral hazard concerns (Johansson and Palme, 2005). The results found in this paper suggest that cutting the replacement rate for sick leave reduces per capita sickness absences. However, the net effect of the policy on labor costs and overall welfare remains ambiguous as it depends, among other things, on whether the reduction in absences is accompanied by productivity gains, and on whether it corresponds to a reduction of moral hazard or to employees going to work sick (see for instance (Stearns and White, 2018), which show that mandated sick leave policies can reduce overall sickness absences, by limiting the extent to which workers' illness spreads to customers and coworkers).

The paper proceeds as follows: Section 2 describes the reform, Section 3 gives a brief description of the context in which this paper studies the reform, Section 4 describes the data and the sample selection procedures, Sections 5 and 6 discuss the results, and the last section concludes.

2. The change in sick leave policy: compensation and monitoring

With law n. 133/2008, the newly installed right-wing government modified the regulation on sick leave for workers employed in the public administration sector. The reform was highly publicized, as it was part of a broader campaign against inefficiencies in the public sector that received a lot of media attention and it affected about 3 million workers.

The change involved both sick leave compensation and monitoring. Before the reform, public employees were entitled to receive full compensation during sick leave.[3] The 2008 reform reduced the replacement rate by applying cuts to ancillary payments and productivity bonuses. The general provision of the law applies to all public employees, but the monetary cuts to be applied to sick leave compensation vary across workers depending on their compensation structure, as determined by national collective bargaining contracts stipulated between unions and the Agency for Bargaining Representation of the Public Administrations (ARAN). Hence, the reform induced variation in the replacement rate of sick leave across occupational categories, job positions, tenure, etc. On average, the replacement rate for sick leave dropped by 20 per cent in the public sector as a whole (RGS, 2008), and by about 5 per cent in the public health sector.

Monitoring takes the form of random medical visits. Before 2008, workers on sick leave were required to be available during a 4-hour time window, from 10 am to 12 pm and from 5 to 7 pm. Such obligation ceases when a designated physician certifies the worker's temporary incapability to work. The 2008 reform increased the time window for random inspections to 11 hours, from 8 am to 1 pm and from 2 to 8 pm. In July 2009, a new law change restored the historical 4-hour time window. Starting in February 2010, the time window for inspections covers 7 hours, from 9 am to 1 pm and from 3 to 6 pm. The law does not allow for more than one visit per sick leave episode, the rationale being that inspections only serve the purpose of certifying temporary inability to work, and cannot be used to force the worker to stay at home during sick leave. Unjustified absence of a worker during the inspection leads to loss of the replacement rate and one day of salary, and eventually to layoff.[4]

3. Data and sample selection

I use data from the Italian Annual Count, a yearly census survey, managed by the General Inspectorate for Personnel Regulations and the Analysis of Public Sector Labor Costs (IGOP), collecting information on public sector employment and labor costs. The survey is conducted by the State General Accounting Department according to the provisions set forth by Title V of Legislative Decree n. 165/2001 and covers all the institutions that are part of the public administration aggregate and fall under the provisions of the above-mentioned decree.

The information collected through the survey is the official information base for Parliament and Government decisions concerning public sector employment. The data are also used for the drafting of the annual report to the Parliament on the management of the financial resources assigned to public sector personnel.

The scope and the coverage of the survey is very broad; in fact, it targets nearly 10 thousand public administrations, accounting for about 3.4 million employees and over 134 billion Euros of annual expenditures for personnel.

The data are collected by institution and position in the occupation and include number of employees, number of absences, wages, salaries, and supplementary components of pay. I use data from 2001 to 2012 for all the local institutions of the National Health System (see the Appendix for institutional details of the Italian NHS). Over the years, some institutions merged into a bigger institution and the definition of some job positions changed over time. I aggregate the data across the different institutions/positions up to the level of aggregation for which I have data in every year. I restrict the sample to those observations for which none of the relevant variables are missing and drop the top 5 per cent of the distribution of per capita sick leaves for each occupation. I include five occupations that account for 98–99 per cent of the total Italian National Health System employees. The included occupational categories are medical doctors, healthcare executives, healthcare services providers (nurses from now on), admins, and technicians. Finally, I balance the panel by dropping all the job position-by-institution cells that I do not observe in every year. The resulting data set consists of 34,800 observations, 2,900 cells per 12 years. The final sample includes 206 institutions and five occupational categories, further divided into 27 job positions. If data for a given job position at a given institution are missing in any given year, that cell is dropped from the sample. This means that an institution is in the sample if data about at least one job position are available for every year from 2001 to 2012. In order to ensure that the sample selection does not drive the results, I repeat the analysis using different sample selection rules for both the balanced and the unbalanced panels.

Table 1 reports summary statistics. The average job position-by-institution cell has about 138 employees but the median cell has 57 employees. Given that the cells have different sizes, it is not surprising that there is huge variation in the number of days off for sick leave. Average per capita sickness absences are about 11 days in a year, and the average ratio between non-base compensation and total compensation is about 0.05, meaning that the average cut to the replacement rate for sick leave amounted to 5 per cent in the public health sector.

3.1. The evolution of sick leave absences over time

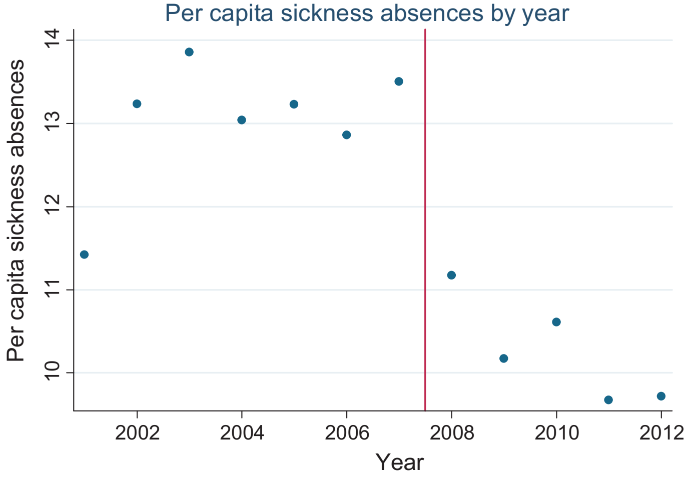
This section shows the evolution of sick leave absences over time in the Italian National Health System. Figure 1 shows the number of days off per employee in each year due to sick leave absences. Per employee sickness absences increase between 2001 and 2003 and oscillate between 2003 and 2007 around an average level of about 13 days per year. There is a sharp decrease in 2008, which continues in 2009. They are slightly higher in 2010, but they go down again in 2011 and 2012. Figure 2 shows the evolution of per employee sickness absences by occupation. Doctors have the lowest level of absences, with an average number of per capita sick leave absences of 7.3 between 2001 and 2007. The same statistic takes value 10.6 for non-medical healthcare executives. Nurses take an average of 13.7 days off per year before 2008, admins 14.4, and technicians show the highest absences level,

Table 1. Summary statistics

	Mean	**sd**	**p25**	**p50**	**p75**
Number of employees	137.83	274.84	27	57	123
Days off for sick leave	1640.58	3672.96	184	579	1467.5
Per capita sickness absences	10.94	6.66	5.87	9.90	14.78
Non-base to total compensation	0.05	0.04	0.02	0.04	0.07
N	34800				

The table reports mean, standard deviation, 25th, 50th, and 75th percentiles for the distribution of number of employees, number of days off in a year for sick leave, and per capita sickness absences across job positions by institutions cells for the years 2001–2012. The last row reports summary statistics for the ratio between non-base compensation and total compensation averaged across years 2001–2007.

Figure 1. Per capita number of days off due to sick leave by year. The vertical line indicates 2007. The change in sick leave policy happens in June 2008



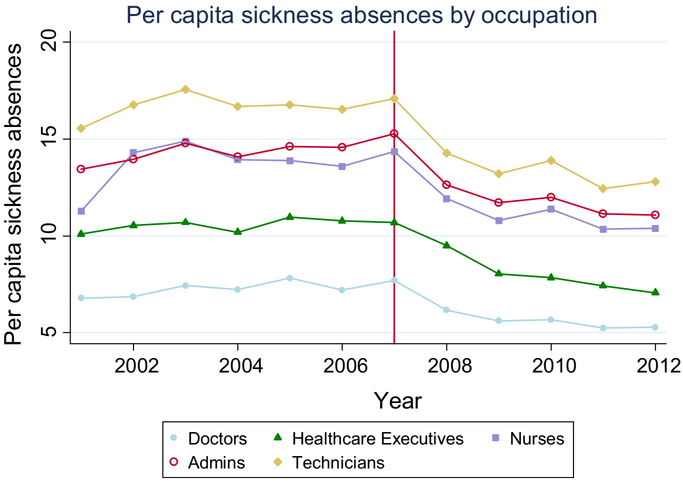
with an average number of days off for sick leave of 16.7. Starting in 2008, there is a decrease in per capita sick leaves for all the occupations. The average number of sickness absences goes down by about 3 days for all occupations, except for doctors, whose per capita absences go down by less than 2 days.

Subsequent monitoring changes took place in July 2009 and February 2010. In July 2009, the time window for medical inspection goes from 11 to 4 hours, and in February 2010, it goes up again from 4 to 7 hours. There is no evidence of an increase in absences in 2009; however, the less restrictive monitoring was in force only for 6 months, so the yearly structure of the data does not allow to draw clear conclusions. Figure 2 shows a small increase in the level of absences for non-managerial positions in 2010 relative to 2009. This increase coincides with the strengthening in monitoring that takes place in February 2010. However, it is also consistent with a delay in the effect of monitoring. Thus, the increase registered in 2010 could be a residual effect of the weakening of monitoring introduced in 2009. The fact that the increase is only registered for non-managerial positions is consistent with the idea that non-managerial positions respond more to monitoring changes. D’Amuri (2017) suggests that this phenomenon can be explained by wage compression in the public sector and the existence of rents for low job positions. The dynamics shown in Figure 2 are also consistent with the opposite result: Doctors and healthcare executives respond to the 3-hour increase in monitoring more than other occupational categories. The monitoring changes applied to all public workers, so it is not possible to identify an untreated group that would allow to estimate the effect of monitoring *per se*.

4. Increasing the monetary penalty for an absence: strategies and results

As explained above, the policy changed both the monitoring and the monetary penalty of a sickness absence. Before the reform, workers’ compensation included all bonuses and

Figure 2. Per capita number of days off due to sick leave by occupation and year. The vertical line indicates 2007. The change in sick leave policy happens in June 2008



accessory payments for the first six months of sickness absence. The reform reduced the replacement rate by applying cuts to ancillary payments and productivity bonuses. The compensation structure of Italian public sector employees is mostly determined by national collective bargaining agreements. Labor relations in the National Health System are governed by three National Collective Bargaining Agreements. Each of them applies to a broad category of workers: doctors and vets, non-doctors' health executives, and non-executives. The last category includes nurses, admins, and technicians. The wage structure varies to a great extent across occupations and within occupations across job positions and tenure. There is thus variation in the reduction of the replacement rate across workers.

In order to separately identify the effect of the changes in the monetary costs from the changes in the monitoring introduced with the 2008 reform, I use a difference-in-differences strategy exploiting variation in the monetary cost of an absence across job position-by-institution cells.

The reform at study cuts the replacement rate for the first 10 days of sick leave absences. More specifically, wage cuts are to be applied to ancillary payments and productivity bonuses. Thus, the change in the monetary penalty for an absence varies across workers depending on the structure of their compensation. Using administrative data on workers' compensation from 2001 to 2007, I build the average percentage incidence of ancillary payments on total compensation for each job position and institution. Figure 3 shows the overall variation in this absence penalty measure, whereas Figure 4 shows the variation across job position-by-hospital cells for each occupation.

The main estimating equation is.

$$\log(\text{p.c.sickdays})_{\text{jht}} = \alpha_{\text{jh}} + \alpha_t + \beta \cdot Z_{\text{jh}} \cdot \text{Post}_t + \epsilon_{\text{jht}} \quad [1]$$

in which $\log(\text{p.c.sickdays})_{\text{jht}}$ is the natural logarithm of per capita sickness absences in job position j, institution h, and year t, α_{jh} is a full set of job position-by-institution fixed

Figure 3. Overall variation in the non-base compensation relative to total compensation. This measure is obtained by computing the average ratio across years 2001 to 2007 between non-base compensation and total compensation at the job position-by-institution level

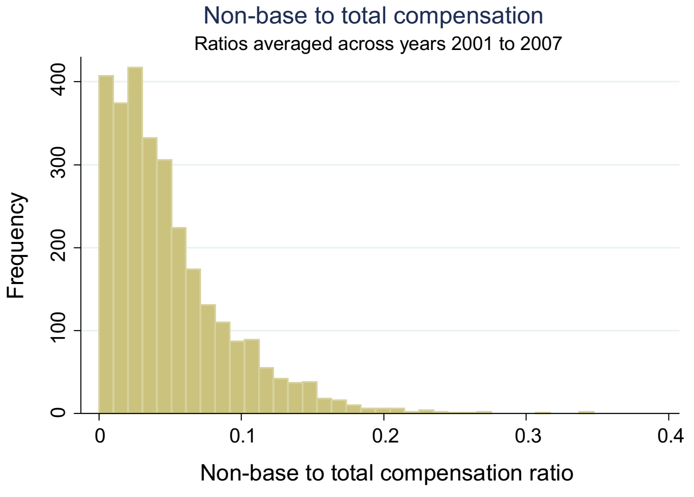
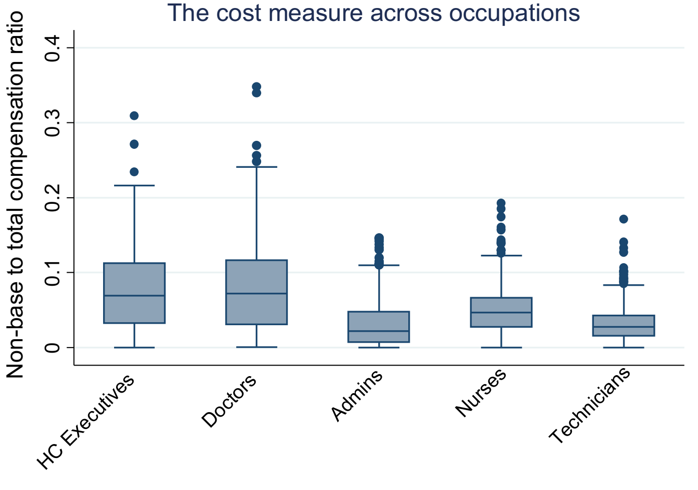


Figure 4. Box plots of non-base to total compensation ratios, averaged across years 2001 to 2007, by occupation. The within-occupation variation is a combination of variation across job positions and across institutions



effects, α_t indicate year fixed effects, Z_{jh} is the absence penalty measure described above, and $Post_t$ is a dummy variable that takes value one from 2008 onward.

The specification in equation [1] allows for different absences levels for each job position-by-institution cell – i.e., the same job position is allowed to have different absences levels across institutions, and the same institution is allowed to have different absences levels across job positions. Thus, identification of the causal effect relies on the assumption that any difference in trends across job position-by-institution cells is independent of Z_{jh}. Under the assumption that the change in monitoring has the same percentage effect on absences across cells, we can interpret the coefficient β in equation [1] as the effect of decreasing the replacement rate for sick leave absences. Figure 5 shows that there is a negative correlation between sick leave levels in the pre-period and the cost measure Z_{jh}. In order to mitigate the potential concern of heterogeneous monitoring effects across absences levels, one may consider to include the average level of per capita absences in the pre-period interacted with $Post_t$ in equation [1], but OLS gives inconsistent estimates in the presence of serial correlation. I thus choose to parameterize the dependence of the monitoring effects on the levels of per capita absences using a log specification and assume that monitoring has a homogeneous proportional effect across cells.

Table 2 reports the estimates for the coefficient of interest for different models. Standard errors are clustered at the institution level in every specification. Each column includes a different set of fixed effects. Column (1) in Table 2 reports the estimates for the interaction between the absence penalty measure Z and the Post dummy in a specification that includes institution fixed effects, job position fixed effects, and year fixed effects. A 1 percentage point increase in the absence penalty measure corresponds to a 1.24 per cent decrease in per capita sickness absences. Column (2) in Table 2 reports the estimates from equation [1]. A one percentage point increase in Z reduces per capita sickness absences by about 0.65 per cent. Column (3) in Table 2 includes institution-by-year fixed effects to account for potential differential impacts of monitoring across institutions. The point estimate from this model suggests that a one percentage point increase in Z reduces absences by about 1 per cent. The decrease in the estimated coefficient from column (1) to column (2) reflects the fact that lower levels of per capita absences are correlated with higher Z within hospitals across job positions. Thus, the additive structure for job position and hospital fixed effects is rejected. The decrease in the estimated coefficient from column (2) to

Figure 5. Cross-sectional relation between per capita sickness absences in 2005 and the absence penalty measure Z. Per capita sickness absences are averaged over 50 quantiles of the absence penalty measure

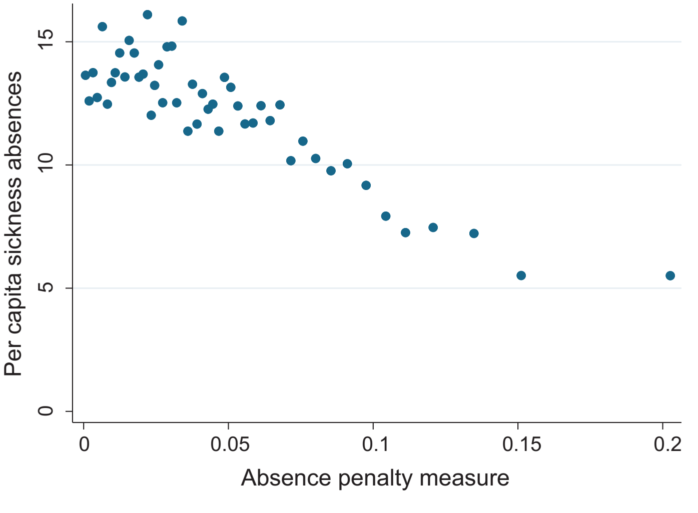


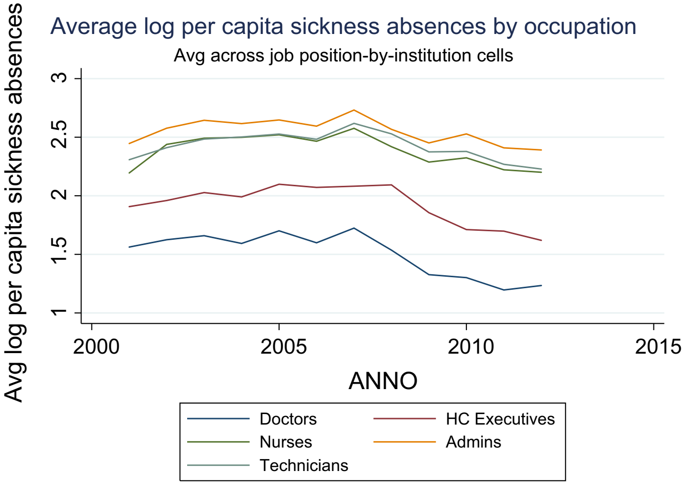
Table 2. Sick leave response to changes in the absence penalty measure

	Log(Per capita sickness absences)				
	(1)	**(2)**	**(3)**	**(4)**	**(5)**
Z_Post	−1.24****	−0.65***	−1.02****	0.06	−0.19
	(0.24)	(0.25)	(0.21)	(0.23)	(0.26)
Institution FEs	Yes	Yes	Yes	Yes	Yes
Job position FEs	Yes	Yes	Yes	Yes	Yes
Institution-by-job position FEs	No	Yes	Yes	Yes	Yes
Institution-by-year FEs	No	No	Yes	No	Yes
Occupation-by-year FEs	No	No	No	Yes	Yes
N	34,800	34,800	34,800	34,800	34,800

Column (1) includes institution fixed effects and job position fixed effects, and column (2) includes a full set of institution-by-job position fixed effects. Column (3) adds institution-by-year fixed effects to account for heterogeneous monitoring effects across institutions. Column (4) includes a full set of institution-by-job position fixed effects and occupation-by-year fixed effects. Column (5) includes institution-by-job position, institution-by-year, and occupation-by-year fixed effects. Standard errors in parentheses are clustered at the institution level to allow for arbitrary serial correlation within institution. The sample includes 206 institutions, 27 job positions, and five occupational categories: doctors, healthcare executives, admins, nurses, and technicians. The table only reports the coefficient of interest, namely the interaction between the absence penalty measure Z and a dummy for the postreform.
*$P < 0.10$,
**$P < 0.05$,
***$P < 0.01$,
****$P < 0.001$.

column (3) in Table 2 is consistent with the idea that institutions with higher levels of absenteeism are more affected by the change in monitoring. Given that higher levels of Z correspond to lower levels of per capita absences in the pre-reform period, it can be that institutions with higher Z experience smaller reductions in absences for the part that concerns monitoring. Thus, the estimated effect of the absence penalty measure is biased toward zero when not controlling for institution-specific time effects. The estimates in column (4) and (5) in Table 2 are obtained, respectively, including occupation-by-year fixed effects alone and with institution-by-year fixed effects. These two specifications, respectively, allow for the monitoring change to have different impacts across occupations and both across occupations and across institutions. One can argue that random medical inspections have different impacts on doctors than technicians, e.g., because of within-profession favors. The point estimates are close to zero and not significant. This result might be a combination of the institutional setting that determines public workers' compensation structure in Italy and measurement error. The compensation structure of workers in the same occupation is determined by the same collective bargaining agreements. Within occupation, the compensation structure can vary based on productivity, workers' composition, and availability of funds. The standard errors in column (4), however, do not increase considerably, suggesting that there is enough variation within occupations. One possibility is that the measurement error in the absence penalty measure due to misreporting is more severe within occupations than across occupations. Thus, the inclusion of occupation-by-year fixed effects might worsen the downward bias and even revert the sign of the estimates. In other words, the variation across institutions and job positions within the same occupation might be misleading. If measurement error has a systematic component at the

Figure 6. Across cells average log per capita sickness absences by occupation. The figure plots the average computed across job position-by-institution cells within each occupation over time. The different occupations appear to be on parallel trends before the introduction of the sick leave policy change



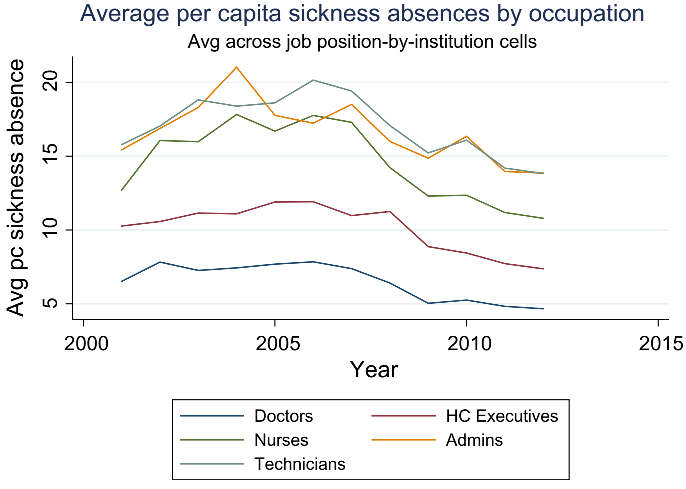
institution level, then the measurement error issue becomes less serious when I include institution by year fixed effects. This observation might explain why the point estimate of interest is slightly more negative for the specification in which both institution-by-year and occupation-by-year fixed effects are included.

Columns (2) and (3) in Table 2 suggest that the change in the absence penalty measure has a sizable negative impact on absences. However, this result is consistent with the hypothesis that most of the reduction in sick leave absences following the 2008 reform is driven by the strengthening of the monitoring, as the actual changes in the absence penalty measure were fairly small (Z has mean 0.05).

Given that columns (2) and (3) rely on across occupations comparisons over time, it is important to test whether the dynamics of per capita sickness absences are comparable across occupations. Figure 6 shows the trends across occupations before and after the introduction of the policy and provides support to the assumption of parallel trends across occupations. Under the assumption of homogeneous proportional monitoring effects across occupations, the estimates in column (3) provide the causal effect of a change in the absence penalty measure on per capita sickness absences. Figure 7 shows that the levels (rather than the logarithm) of per capita sickness absences are not on parallel trends across occupations, thus providing additional support for the log specification.

Table 3 shows that the results are not driven by the exclusion of cells that do not appear in all years. A comparison between panel A and panel B in Table 3 confirms that the results are very similar in the balanced and the unbalanced panels. Table 4 shows the estimates using different sample selection rules for the unbalanced panel. Panel A uses the full sample, panel B excludes the top 5 per cent of the per capita absences, panel C excludes the top 3 per cent, and panel D the top 1 per cent. The results are very similar across all sample selection rules.

Figure 7. Across cells average per capita sickness absences by occupation. The figure plots the average computed across job position-by-institution cells within each occupation over time. Absences levels do not appear to be on parallel trends across occupations before the introduction of the sick leave policy change, thus suggesting that a level specification is not supported by the data



5. Event study analysis

Identification in the difference-in-differences setting relies on the assumption that differences in trends across institution-by-job position cells are uncorrelated with treatment status. This section presents the event study analysis for the changes in the absence penalty measure.

Figure 8 shows the estimates from the following model:

$$\log(\text{p.c.sickdays})_{\text{jht}} = \alpha_{\text{jh}} + \alpha_t + \sum_{\tau=2001}^{2006} \beta_\tau \cdot Z_{\text{jh}} \cdot 1(\text{year} = \tau)_t + \sum_{\tau=2008}^{2012} \beta_\tau \cdot Z_{\text{jh}} \cdot 1(\text{year} = \tau)_t + \epsilon_{\text{jht}} \quad [2]$$

in which $\log(\text{p.c.sickdays})_{\text{jht}}$ is the natural logarithm of per capita sickness absences in job position j, institution h, and year t, α_{jh} is a full set of job position-by-institution fixed effects, α_{t} is a full set of year fixed effects, Z_{jh} is the cost measure described above standardized to have mean zero and standard deviation one, and β_τ are time-varying coefficients on the cost measure. Equation 2 includes interactions between the cost measure and year dummies for every year excluded 2007. Under the assumption of parallel trends, $\beta_\tau = 0$ for $\tau < 2007$. Figure 8 reports the point estimates for β_τ in Equation 2 and 95 per cent confidence intervals. There is no evidence of pre-trends, as the point estimates in the pre-period are close to zero and not statistically different from zero.

The point estimates of β_τ for $\tau > 2007$ show the dynamics of the effect of the policy. There is no significant effect in 2008 and in 2009, whereas the effect becomes negative and significant in 2010 and fades away gradually in 2011 and 2012. The dynamics are somewhat surprising as the policy change took place between June and August 2008 and again suggest that the change in the monetary cost had only a marginal impact on absences in the public health sector. On the other hand, the estimated effect might be small because of

Table 3. Sick leave response to changes in the absence penalty measure: comparison between balanced and unbalanced sample

	Log(Per capita sickness absences)				
	(1)	**(2)**	**(3)**	**(4)**	**(5)**
			Panel A: balanced panel		
Z_Post	−1.24****	−0.65***	−1.02****	0.06	−0.19
	(0.24)	(0.25)	(0.21)	(0.23)	(0.26)
N	34,800	34,800	34,800	34,800	34,800
			Panel B: unbalanced panel		
Z_Post	−1.29****	−0.56**	−0.86****	0.15	−0.0739
	(0.24)	(0.25)	(0.21)	(0.24)	(0.23)
N	59,918	59,918	59,918	59,918	59,918
Institution FEs	Yes	Yes	Yes	Yes	Yes
Job position FEs	Yes	Yes	Yes	Yes	Yes
Institution-by-job position FEs	No	Yes	Yes	Yes	Yes
Institution-by-year FEs	No	No	Yes	No	Yes
Occupation-by-year FEs	No	No	No	Yes	Yes

Column (1) includes institution fixed effects and job position fixed effects, and column (2) includes a full set of institution-by-job position fixed effects. Column (3) adds institution-by-year fixed effects to account for heterogeneous monitoring effects across institutions. Column (4) includes a full set of institution-by-job position fixed effects and occupation-by-year fixed effects. Column (5) includes institution-by-job position, institution-by-year, and occupation-by-year fixed effects. Standard errors in parentheses are clustered at the institution level to allow for arbitrary serial correlation within institution. The sample includes 206 institutions, 27 job positions, and five occupational categories: doctors, healthcare executives, admins, nurses, and technicians. The table only reports the coefficient of interest, namely the interaction between the absence penalty measure Z and a dummy for the postreform. Panel A restricts the sample to the job position-by-institution cells that appear in every year, whereas panel B uses all the data.
*$P < 0.10$,
**$P < 0.05$,
***$P < 0.01$,
****$P < 0.001$.

omitted variable bias: The cost measure Z is negatively correlated with absences levels. If the effect of monitoring is higher, the higher the absences level, then the estimates reported in Figure 8 are downward biased. In line with this hypothesis, allowing for heterogeneous monitoring effects across institutions increases the point estimate for the change in cost as reported in column (3) of Table 2 and shown in Figure 9. Such observation motivates my choice of specifying the outcome variable as the logarithm of per capita absences rather than using the levels.

As additional evidence supporting the log specification, I estimate model 2 in levels. Figure 10 reports the estimated coefficients and 95 per cent confidence intervals for the following model:

$$\text{p.c.sickdays}_{\text{jht}} = \alpha_{\text{jh}} + \alpha_t + \sum_{\tau=2001}^{2006} \beta_\tau \cdot Z_{\text{jh}} \cdot 1(\text{year} = \tau)_t + \sum_{\tau=2008}^{2012} \beta_\tau \cdot Z_{\text{jh}} \cdot 1(\text{year} = \tau)_t + \epsilon_{\text{jht}} \quad [3]$$

in which all variables are as defined above, whereas the outcome variable is the number of days off due to sick leave divided by the number of workers in job position j, institution h, and year t. The presence of pre-trends in levels is clear. Before 2007, absences were going

Table 4. Sick leave response to changes in the absence penalty measure unbalanced panel: comparison between different sample selection rules

	Log(Per capita sickness absences)			
	(1)	**(2)**	**(3)**	**(4)**
	Panel A: Full sample			
Z_Post	−1.429***	−0.725**	−1.094****	0.268
	(0.21)	(0.24)	(0.23)	(0.32)
N	62,997	62,997	62,997	62,997
	Panel B: Cut top 5 per cent			
Z_Post	−1.314****	−0.603***	−0.907****	0.156
	(0.193)	(0.228)	(0.213)	(0.238)
N	59,848	59,848	59,848	59,848
	Panel C: Cut top 3 per cent			
Z_Post	−1.353****	−0.616***	−0.927****	0.162
	(0.193)	(0.227)	(0.214)	(0.236)
N	61,111	61,111	61,111	61,111
	Panel D: Cut top 1 per cent			
Z_Post	−1.394****	−0.637***	−0.959****	0.193
	(0.203)	(0.236)	(0.222)	(0.245)
N	62,368	62,368	62,368	62,368
Institution FEs	Yes	Yes	Yes	Yes
Job position FEs	Yes	Yes	Yes	Yes
Institution-by-job position FEs	No	Yes	Yes	Yes
Institution-by-year FEs	No	No	Yes	No
Occupation-by-year FEs	No	No	No	Yes

Column (1) includes institution fixed effects and job position fixed effects, and column (2) includes a full set of institution-by-job position fixed effects. Column (3) adds institution-by-year fixed effects to account for heterogeneous monitoring effects across institutions. Column (4) includes occupation-by-year fixed effects. Standard errors in parentheses are clustered at the institution level. The sample includes 206 institutions, 27 job positions, and 5 occupational categories. Panel A uses the full sample, panel B excludes the top 5 per cent of the per capita sickness absences, panel C excludes the top 3 per cent, and panel D excludes the top 1 per cent.
*$P < 0.10$,
**$P < 0.05$,
***$P < 0.01$,
****$P < 0.001$.

down in cells characterized by higher average non-base to total compensation ratios, whereas the trend turns after the policy change. The observed pattern in the pre-period is consistent with the averages shown in Figure 2: Occupations with higher levels of absences are on steeper trends. The pattern in the period after the reform is consistent with heterogeneity of monitoring effects across levels of absences: It is plausible to assume that the higher the initial level of absences, the higher the probability that workers are shirking by claiming sick leave without needing it; monitoring changes in principle should only affect absenteeism and not induce employees to go to work sick; hence, monitoring strengthening should induce a larger reduction in absences the higher their initial level.

6. Conclusions

This paper studies a reform that took place in Italy in 2008 affecting all public sector employees characterized by both an improvement in monitoring and a decrease in the

Figure 8. Point estimates and 95 per cent confidence intervals for β_τ from equation [2]. The specification includes a full set of institution-by-job position fixed effects and the cost measure interacted with year dummies, excluding 2007. The outcome variable is defined as the logarithm of the number of days off due to sick leave divided by the number of workers in job position j, institution h, and year t

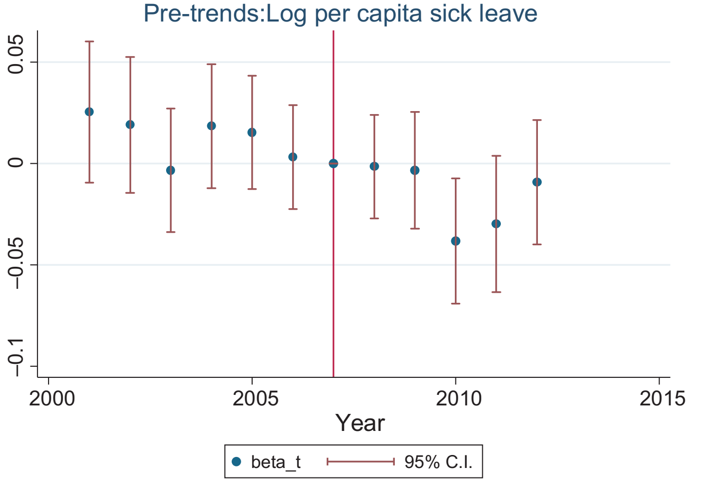
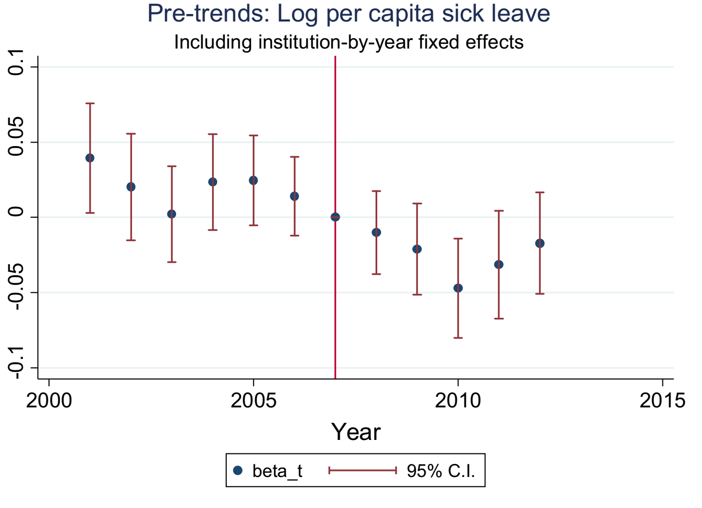


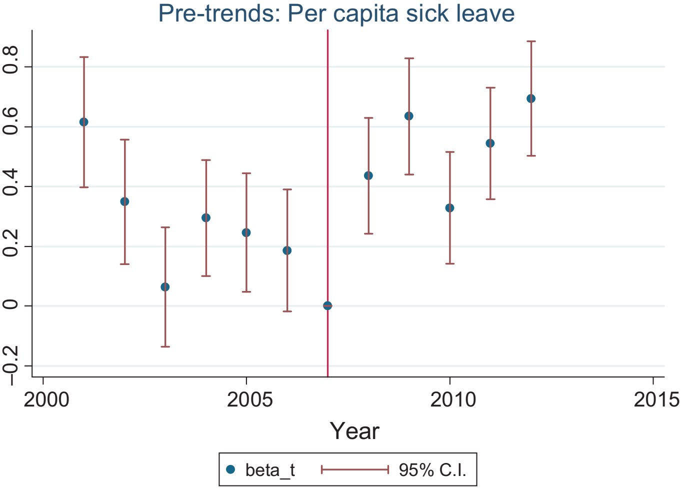
Figure 9. Point estimates and 95 per cent confidence intervals for β_τ from equation [2] adding institution-by-year fixed effects. The specification includes a full set of institution-by-job position fixed effects, institution-by-year fixed effects, and the cost measure interacted with year dummies, excluding 2007. The outcome variable is defined as the logarithm of the number of days off due to sick leave divided by the number of workers in job position j, institution h, and year t



replacement rate. Using a difference-in-differences strategy, I relate the changes in per capita sickness absences with the changes in the replacement rate. Under some identifying assumptions spelled above, I estimate that a 1 percentage point increase in the absence penalty measure reduces the per capita number of days off for sick leave by 1 per cent.

This result is not in contrast with D'Amuri (2017) which analyses the effects of the 2008 reform and subsequent monitoring changes on the public sector workers as a whole (excluding the National Health Service) and finds evidence that most reduction in sick leave absences is due to monitoring changes rather than changes in the replacement rate. In fact, the changes in the replacement rate were fairly small (5 per cent points on average). Thus, the estimate in this paper together with D'Amuri (2017) implies that both a strengthening of monitoring and a decrease in the replacement rate are effective in reducing absences. Increasing monitoring, however, can lead to high costs, which the government would need to finance potentially distorting other sectors of the economy. Reducing the replacement rate reduces costs for the government for equal levels of absences. The net effect on labor costs from the policy is ambiguous, and it depends, among other things, on the productivity gains induced by the reform. Ziebarth and Karlsson (2010) study the effect of statutory pay for sick leave absences in Germany both on absenteeism and on labor costs. They show that the direct impact of reducing the replacement rate is higher than the indirect impact on labor costs through a reduction in absences, with a 6.7 per cent estimated reduction in labor costs. Thus, reducing the replacement rate can be a source of funding to increase monitoring. If the increase in monitoring has a first-order impact on absenteeism, whereas small decreases in the replacement rate correspond to small distortions from the first best full insurance level and a first-order decrease in labor costs, then the policy mix implemented in 2008 in the Italian public sector can be a way to approach optimal sick leave policy. For the monitoring system to work, it is necessary that workers assign a high

Figure 10. Point estimates and 95 per cent confidence intervals for β_τ from equation [3]. The specification includes a full set of institution-by-job position fixed effects and the cost measure standardized to have mean zero and standard deviation one interacted with year dummies, excluding 2007. The outcome variable is defined as the number of days off due to sick leave divided by the number of workers in job position j, institution h, and year t



enough probability to the event of receiving an inspection. If the fear of inspections is enough to deter absenteeism, inspections are ex post inefficient. In 2012, there were about 1.2 millions visits for a total cost of about 50 millions of euros, and only 8.2 per cent of such visits resulted in a reduction of the sick leave. In 2013, the National Institute of Social Security, responsible for the inspections, announced a reduction in the number of visits to limit the cost. It would be interesting to evaluate whether such announcement had an impact on sick leave absences. This would allow to shed light on the extent to which costly inspections can be reduced while keeping unaffected their deterrence effect.

Notes

[1] D'Amuri (2017) excludes the healthcare sector and the education sector because the data he uses do not allow him to identify whether workers are employed in the public or the private sector.

[2] Given that they analyze a cut in the replacement rate from 90% to 65%, their estimate implies that the incidence of sick leave drops by about 0.7 per cent for a one percentage point cut in the replacement rate.

[3] The replacement rate during sick leave of different lengths was regulated by National Collective Bargaining Contract and varied slightly across contractual categories.

[4] If more than 3 unjustified absences are recorder in 2 subsequent years or more than 7 in the last 10 years.

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Appendix

The Italian National Health Service

The reform described above involved all public administrations. This paper focuses on the National Health Service, which accounts for about 21 per cent of the public sector workers. This section provides background information about the Italian National Health System.

The Italian health market is characterized by the presence of universal compulsory public insurance – with a small fraction of the population complementing public insurance with private insurance – and a system of public and private providers. The law guarantees consumers of healthcare services freedom of choice between public and private providers – public insurance covers services supplied by 'licensed' private healthcare providers.

The National Health Service, established in 1978, is a complex system of institutions and services aiming to 'the promotion, maintenance and recovery of physical and psychological health of the entire population...' (art. 1 Law 833/1,978). In 1999, it became a system of Regional Health Systems organized on two levels of political governance – national and regional. At the local level, there are two different types of institutions: Local Health Authorities (LHAs) and Independent Hospitals (IHs). Local Health Authorities are in charge of distributing healthcare services in a given area, both directly and by negotiating agreements with other public providers and licensed private providers. Direct production is carried out by hospitals and specialistic clinics directly managed by the Local Health Authorities. Other public providers include Independent Hospitals, big public hospitals, independent of the Local Health Authorities, which produce healthcare services, based on contractual agreements with Local Health Authorities.

The number of Local Health Authorities has been varying over time with a general tendency to make each LHA coincide with a province. However, major cities have several LHAs. Currently, there are 101 Local Health Authorities (145 in 2012) and 102 Independent Hospitals.