DISCUSSION PAPER SERIES

DP15142

SLOW DOWN BEFORE YOU STOP: THE EFFECT OF THE 2010 FRENCH PENSION REFORM ON OLDER TEACHERS' SICK LEAVES

Hippolyte d'Albis, Denis Fougere and Pierre Gouëdard

LABOUR ECONOMICS

ISSN 0265-8003

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Discussion Paper DP15142
Published 07 August 2020
Submitted 01 August 2020

Centre for Economic Policy Research
33 Great Sutton Street, London EC1V 0DX, UK
Tel: +44 (0)20 7183 8801
www.cepr.org

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Abstract

This paper proposes an evaluation of the pre-retirement consequences of a reform of the French pension system that increased the minimum legal retirement age. Our empirical strategy relies on the comparison of two groups of cohorts. The control group consists of cohorts not affected by the increase in the minimum legal retirement age while the treatment group consists of cohorts born later. Using a sample of 38,652 high-school teachers, we identify the effect of increasing the minimum retirement age on short sick leaves (i.e., of less than three months) by comparing probabilities to take at least one sick leave during a schooling year before retirement across the two groups. Estimates of panel data models show that teachers affected by the reform have an increased probability to take short sick leaves before retirement. This is mainly due to teachers who decide to retire at the minimum legal retirement age, while those who continue to work above the minimum retirement age do not increase the frequency of their short sick leaves before retirement. This last result is predicted by a theoretical model that analyzes the optimal retirement choice over the life-cycle, and it is confirmed by using an empirical strategy that distinguishes teachers according to their retirement age.

JEL Classification: J26, I12

Keywords: Retirement age, pension reform, sick leaves, teachers' absenteeism

Hippolyte d'Albis - hippolyte.dalbis@psemail.eu
CNRS and Paris School of Economics

Denis Fougere - denis.fougere@sciencespo.fr
CNRS, OSC and LIEPP (Sciences Po Paris), IZA (Bonn) and CEPR

Pierre Gouëdard - pierre.gouedard@oecd.org
OECD and LIEPP (Sciences Po Paris)

Acknowledgements
We thank Cedric Afsa and Pierrette Briant (DEPP, French Ministry of Education) for having helped us to access the administrative datasets that we use in this study. We also thank Antoine Bozio, Cecilia Garcia Penalosa, Elise Huillery and Etienne Wasmer for helpful comments. This work is supported by two public grants overseen by the French National Research Agency (ANR): the first is part of the "Investissements d'Avenir" program LIEPP (ANR-11-LABX-0091, ANR-11-IDEX-0005-02) while the second is within the JPI MYBL framework (ANR-16-MYBL-0001-02.

Slow Down Before You Stop: The Effect of the 2010 French Pension Reform on Older Teachers' Sick Leaves*

Hippolyte d'Albis†, Denis Fougère‡, and Pierre Gouëdard§

August 5, 2020

Abstract This article proposes an evaluation of the pre-retirement consequences of a reform of the French pension system that increased the minimum legal retirement age. Our empirical strategy relies on the comparison of two groups of cohorts. The control group consists of cohorts not affected by the increase in the minimum legal retirement age while the treatment group consists of cohorts born later. Using a sample of 38,652 high-school teachers, we identify the effect of increasing the minimum retirement age on short sick leaves (i.e., of less than three months) by comparing probabilities to take at least one sick leave during a schooling year before retirement across the two groups. Estimates of panel data models show that teachers affected by the reform have an increased probability to take short sick leaves before retirement. This is mainly due to teachers who decide to retire at the minimum legal retirement age, while those who continue to work above the minimum retirement age do not increase the frequency of their short sick leaves before retirement. This last result is predicted by a theoretical model that analyzes the optimal retirement choice over the life-cycle, and it is confirmed by using an empirical strategy that distinguishes teachers according to their retirement age.
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†Paris School of Economics, CNRS

‡CNRS, OSC - LIEPP (Sciences Po Paris), CEPR and IZA

§OECD and LIEPP (Sciences Po Paris)

1 Introduction

Many developed countries are currently encouraging their workers to postpone retirement. In a context of increased life expectancy, fiscal stability of the pension system is jeopardized if no additional contributions compensate for pensions benefits that have to be paid over a potentially longer period for each retiree. This is especially true in France, where the system is of the pay-as-you-go kind (i.e., a contributory pension scheme), which implies that the financial balance directly depends on the worker-retiree ratio. As baby boomers massively retire since the end of the 1990s, the French Government has launched several reforms aiming at keeping people at work in order to ensure the fiscal sustainability of the system.

An abundant literature has been dedicated to study the effect of delayed retirement on individual well-being and health. Main pitfalls consist in selection bias and endogeneity of the retirement decision. In fact, if people in bad health retire first, in other words if they self-select out, then the impact on health of postponing retirement could be misestimated. A spurious positive correlation between work and health might then appear (see, for instance, Kalwij and Vermeulen (2008)). Health might affect work, and vice versa. If it is obvious that bad health might induce early retirement, it is also true that staying employed helps maintaining a certain level of health. Some articles have indeed shown that the retirement decision is traumatic, and can trigger off a quick decline in the health status. Overall, many unobserved factors may blur the relationship between health and retirement and make it difficult to infer the causal effect of retirement on health.

Most of the studies examining the health-retirement link are based upon a comparison between employees and retirees. On the one hand, work is supposed to induce a stress detrimental to health, as shown by Ekerdt *et al.* (1983). For instance, in the US, Coe and

Lindeboom (2008) find a temporary increase in the self-reported measure of health after retirement, their results being confirmed by Coe and Zamarro (2011) who use European data. Similarly, Neuman (2008) provides evidence that retirement maintains health according to subjective measures, but finds no effect on objective health variables. Bound and Waidmann (2007) show that retirement has a temporary positive effect on health of men (only) in the UK. On the other hand, retirement might be considered as a stressful event, after which retired people lack the physical and mental activity that is associated with work. As work increases activity, income and social networking, some other studies present alternative results. Bonsang *et al.* (2012), but also Rohwedder and Willis (2010), show how retirement induces a negative effect on memory. Dave *et al.* (2008) document how earlier retirement is associated with poor physical and mental health after retirement.

In this study, we depict the recourse to the health care system by workers close to the statutory retirement age by exploiting an exogenous reform of the French pension system implemented in 2010. We use information on teachers' individual sick leaves provided by the *Département de l'Évaluation, de la Prospective et de la Performance* (DEPP, hereafter), the Statistics Department of the French Ministry of Education. The focus is on secondary-school teachers who, because of the 2010 reform, face different minimum statutory retirement ages depending on their birth cohort. Prior to this reform, secondary-school teachers had the possibility to retire as soon as they reached 60 years old. After the reform, younger cohorts (namely those who reached 60 in 2011 or after) faced a progressive increase in the minimum legal retirement age, from 60 to 61 and 2 months, between 2011 and 2015. We analyze differences in the frequencies of sick leaves taken before 60 years old by teachers who are subject to different minimum legal retirement ages (since they belong to different birth cohorts) and interpret them as "voluntary" sick leaves, provided that teachers from these cohorts are indeed similar.

Consequently, our study lies at the interface between two streams of the academic literature, namely the health-work relationship and the monitoring of workers' absenteeism. A large literature has been devoted to the impact of sick-leave policy or disability insurance on workers' absenteeism (see, e.g., Winkler (1980), Meyer *et al.* (1995), Gruber (2000), Neuhauser and Raphael (2004) and Ziebarth and Karlsson (2010)). Unsurprisingly, the more generous the system, the more absenteeism. Other studies enumerate procedures which may reduce absenteeism, by aligning workers' and employers' objectives (see, e.g., Bowers and McIver (2000), Bowers (2001), and Clotfelter *et al.* (2009)).

Exploiting the increase in the minimum legal retirement age which was implemented since 2011, we show how teachers in cohorts affected by this reform took more often "short" sick leaves (i.e., of less than three months) than similar teachers of pre-reform cohorts. The reform thus had an adverse effect in the sense that after its implementation, teachers were more likely to take short sick leaves, probably to compensate for the increase in the total disutility of work associated to a longer working lifetime. This increase in absenteeism propensity, which was hardly predictable by the legislator, had nevertheless a non negligible impact on public finances, and reduced the fiscal gains of the reform. Over and above, once this kind of behavior is observed, some monitoring could be implemented to deter it, because, as Clotfelter *et al.* (2009) wrote, "whatever the importance of strong training, classroom experience, or advanced pedagogical methods for the scholastic development of students, these factors can have scant effect on a day a teacher is away from school."

Several studies have assessed the effects of reforms increasing the minimum retirement age on individual health after retirement (see, for instance, recent papers by Eibich (2015), Shai (2018), and Hagen (2018)). Our paper is one of the very few studies that consider the effects of a change in the minimum retirement age or in the pension system rules on the pre-retirement labor supply or health status of older workers. Among these very few studies, we can mention the article by De Grip *et al.* (2011) who assess the impact of a reform of the Dutch pension system on the mental health of workers nearing retirement age, the article by Fitzpatrick (2014) who provides evidence about how public sector retiree health insurance affects the labor supply of public employees, and the study by McCarthy and Wang (2019) who show that pension rules (including overtime and other payments used to determine pensions) give a strong incentive to allocate overtime to senior workers within work teams.

To rationalize our findings, we develop a theoretical model elucidating the mechanisms through which workers react to an increase in the minimum statutory retirement age before retiring. This model allows workers to choose their retirement age, which may be either equal to the minimum retirement age (in which case workers are said to be "constrained") or greater than this minimum age (in which case they are said to be "unconstrained"). This model has three main predictions: (*i*) among workers who are not constrained, there are no significant differences in the probability to take a sick leave between those in the treatment group (i.e., those cohorts affected by the reform) and those in the control group (i.e., those cohorts not affected by the reform), (*ii*) among workers who are constrained, the probability to take a sick leave is higher for those in the treatment group than for those in the control group, and (*iii*) the proportion of constrained workers within a given cohort is higher in the treatment group. These theoretical predictions are then tested. By splitting the sample between teachers who leave at the statutory retirement age and the others, we show that only the former respond to the reform by an increase in the frequency of short sick leaves.

The next section presents the institutional framework of the 2010 French pension reform. Section 3 presents our data and some descriptive analysis. Section 4 presents econometric estimates that confirm descriptive statistics presented in Section 3. Section

5 presents our theoretical model and proposes some empirical claims that are tested in Section 6. Section 7 concludes.

2 Institutional framework

Starting from 1993, a series of pension reforms have extended the duration of contributions required to obtain a full pension rate in France. The 1993 reform, which was implemented under the government of the then Prime Minister Edouard Balladur, concerned only the private sector. It increased the number of years required for getting full pension. This number was previously equal to 37.5 years. After the reform, it increased by one quarter each year until 2003, in order to reach 40 years in 2003. Moreover, the 1993 reform changed the number of years on which individual past wages are averaged in order to compute the reference wage used for calculating the pension rate. From 1993 on, this number has increased from 10 to 25 years. Finally, before 1993, past wages and pensions were discounted according to GDP growth, while since the 1993 reform, they are only indexed on inflation.

Then, the 2003 pension reform, which was launched under the authority of the then Ministry of Social Affairs, François Fillon, was an extension of the 1993 reform to the French public sector. The primary objective was to ensure a convergence of conditions for access to a full pension rate in the private and public sectors. It introduced four main changes. First, it introduced a gradual increase in the number of quarters required to get the full pension in the public sector, which were raised from 37.5 to 40 years between 2004 and 2008. Second, it introduced two incentive mechanisms: a penalty (*decote*) for those who retire before getting the required quarters of contribution, and a premium for those who postpone retirement (*surcote*). The penalty was implemented from 2006 onwards and increased since then. For instance, it reduced the pension by 0.125 percentage point (pp, hereafter) per missing quarter in 2006, and by 1.250 pp in 2015. The premium was implemented from 2004 onwards and also increased since then. For instance, the pension was increased by 0.75 pp per extra quarter in 2004, and by 1.250 pp in 2015.[1]

In France, employees of some government-owned corporations (e.g. military, the police, energy companies, public transport workers, opera workers, members of parliament) enjoy a special retirement plan: they have lower retirement ages and require fewer working years for getting the full pension rate. These special pension regimes have much more alarming support ratios than the private sector schemes, and thus require significant taxpayer funding. These special pension schemes were reformed in 2007. The objective was to align the duration of employee contribution of the public transportation and energy sectors with those of the private and the public sectors.

The 2010 reform, which was implemented under the authority of the then Ministry of Labor, Eric Woerth, was designed to solve the financing problem that was exacerbated by the 2008 crisis. The 2010 reform increased gradually the minimum statutory retirement age (i.e. the age at which pension benefits become available, albeit not in full) from 60 to 62 years old for those born after 1951.[2] More precisely, the minimum statutory retirement age was set to 60 years and 4 months for employees born in 1951, to 60 years and 9 months for those born in 1952, and to 61 years and 2 months for persons born in 1953 and after. Moreover, the maximum retirement age was maintained at 65 for employees born before 1951 and increased to 67 for those born after.

Thanks to our collaboration with the statistical Department of the French Ministry

[1] Using a model *à la* Stock and Wise (1990), Fougère and Gouëdard (2019) detail how the parameters of the pension calculation have been modified and how these financial incentives have affected the retirement decision of teachers.

[2] Note that people who started working at age 14 could retire at age 58 and those with harmful jobs could still access a full pension at age 60, but the teachers are not concerned.

of Education (DEPP), we have had access to individual data concerning the effective retirement age and other characteristics (including the occurrence of at least one sick leave each year before retirement) for teachers belonging to cohorts 1948-1953. Two groups of cohorts are therefore concerned differently by the reform: teachers born before 1951, who can still retire at 60, and the others who have to retire later. It has to be noted that all of these cohorts are similarly impacted by the financial incentives introduced by the previous 2003 reform.[3] Only the minimum legal retirement age varies across cohorts. Thus, in our study, teachers belonging to 1948-1950 cohorts, who can still retire at age 60, are called the control group while the younger cohorts (1951-1953) constitute the treatment group.

Due to technical and administrative constraints, it was not possible for us to observe sick-leave periods of teachers after the end of year 2013.[4] Despite this, we know the effective retirement age of each teacher, regardless of her year of birth. Figure 1 shows the impact of the reform on the retirement behavior of the control and treated cohorts. We observe the sharp drop in the retirement probability at age 60 and a shift in the distribution towards the age of 61. The increase in the minimum legal retirement age has clearly delayed retirement. Note that the peak at age 65 corresponds to the maximum age of retirement for teachers born in 1951 or after.

Figure 1 about here.

Figure 2 represents the average probability to take at least one sick leave each year between 51 and 59 years old separately for the control and and treated cohorts. We distinguish here between what we call "short" sick leaves (of less than three months) and "long" sick leaves (of 3 months and more).[5] First, we must notice that there is no significant

[3]Public school teachers are civil servants. As such, they were affected by the 2003 reform.

[4]However, this allows us to observe the sick leaves of teachers born in 1953 until 2013, i.e., when they reach their 60th birthday

[5]This is the only information available to us in the data files. In particular, for reasons of medical secrecy, we have no information about the exact duration of each sick leave, its cause, and the number of sick leaves experienced by a teacher during a given school year.

difference between the frequencies of long sick leaves, which are often related to serious diseases, in the control and treatment groups (in this case, we have grouped on one side all the control cohorts, and on the other side all the impacted cohorts). Conversely, a discrepancy appears across cohorts with respect to the occurrence of at least one short sick leave during a year. Here we still group the control cohorts, but we distinguish the treated ones (1951, 1952, 1953). For teachers belonging to the control cohorts, the probability to take a short sick leave is constant with respect to age. This probability increases significantly at age 58 for the impacted cohort born in 1951, so during the year 2009. For the impacted cohort born in 1952, this probability increases significantly at age 57, also during 2009. Finally, for the impacted cohort born in 1953, the probability to take at least one short sick leave increases significantly at age 56, so once again in 2009. Just before age 60, the maximum values of these probabilities are remarkably similar for the three impacted cohorts (between 48 and 49%), while this frequency stays approximately equal to 42% for the control cohorts born in 1948, 1949 or 1950.

Figure 2 about here.

One important remark should be made at this stage. Figure 2 shows that, on average, teachers belonging to cohorts affected by the reform increased the frequency of their short sick leaves one or two years before the adoption of the reform, which was submitted to the French Parliament on September 7, 2010, and promulgated by the Constitutional Council on November 10, 2010. This is not surprising: it means that the reform was not unexpected. Let us recall some historical facts. On October 29, 2007, the French Pension Guarantee Board, *Commission de garantie des retraites*, issues an opinion advocating the gradual extension of the contribution period for all employees, civil servants, the liberal professions and the self-employed by 40 to 41 years. Three weeks after, on November

22, 2007, in its 5th Update Report, the Pension Policy Board, *Conseil d'Orientation des Retraites*, (COR, hereafter) reveals that the implementation of the 2003 pension reform did not cause a sufficient increase in the average retirement age and that the measures aimed at prolonging the activity of older workers did not make it possible to influence the behavior of employers and employees. An undated summary note, which was commissioned by the COR General Secretariat in order to prepare its 5th Update Report, mentions that in the most probable scenario, the insurance period required for retiring at the full pension rate should increase from 40 years for cohort born in 1948 (and which reaches 60 years old in 2008) to 41 years for cohort 1952 (age 60 in 2012).[6] On April 28, 2008, the French Government declares that the insurance period required to be entitled to a full pension rate will be increased by one quarter per year to reach 41 years on January 1, 2012. These historical elements show that, about one or two years before enacting the 2010 reform, workers born after 1950 could expect with a relative high probability that their minimum statutory retirement age would be increased; they consequently modified their behavior one year before the enactment of the reform.

The main hypothesis that we would like to test is the following: as long as older teachers are only "incited" to delay retirement, there is no reason for them to modify their retirement behavior, other things being equal. In particular, this is the case when they can always retire at age 60, eventually with reduced pension benefits. But as soon as they are forced to postpone their retirement date, they could want to compensate the disutility associated with a longer working period by more frequent short sick leaves before retiring. One could also consider that this reaction represented some form of opposition to the 2010 reform, which was not popular.[7] The fact that the probability to take at least one long

[6] Source: http://www.cor-retraites.fr/IMG/pdf/doc-927.pdf

[7] The announcement of the reform caused "a series of general strikes and demonstrations which occurred in France throughout September and October 2010. They involved union members from both the private and public sectors protesting in cities, including Bordeaux, Lille, Lyon, Marseille, Paris, Toulouse,

sick leave during a year does not seem to be modified (cf. Figure 2) is in line with this opportunistic behavior, since long sick leaves are far less manipulable than short ones. This approach was already proposed by Winkler (1980), who studied absenteeism and focused on short-term absences, assuming they are less likely to be due to illness than long-term absences.

3 Data

The administrative data that we use are hosted at the DEPP. We have constructed an original database by merging two types of statistical information. First, information concerning teachers' sick leaves comes from annual administrative files which are pooled to follow teachers' sickness history over time. From these administrative files, which are covered by medical secrecy, we were authorized to collect information about the fact that a teacher has taken at least either one short sick leave or one long sick leave during a school year, for all school years between 2000-2001 and 2012-2013. However, we have no information about the exact duration of each sick leave, its cause, and the number of sick leaves experienced by a teacher during a given school year. Then, by using teachers' identifiers, we have merged these files with administrative databases providing information on some individual characteristics of teachers, including their birth date, their gender, their marital status, their spouse's birth date, their number of children, and the identifiers of the high schools to which they are assigned. After merging we finally get a panel data set including all public high school teachers born between January 1, 1948, and December 31, 1953.

Montpellier and Strasbourg [...] The strikes have led to a reduction in public transport services, motorway blockages by lorry drivers and disruption to oil deliveries to refineries leading to a national fuel shortage. French students also joined the workers in the protests with barricades being built at around 400 high schools across the country in order to try to prevent other pupils attending classes." Source: https://en.wikipedia.org/wiki/2010 French pension reform strikes.

3.1 Descriptive statistics

Our sample contains 38,652 teachers: 22,320 were born between January 1, 1948, and December 31, 1950 (this is our control group), 16,332 were born between January 1, 1951, and December 31, 1953 (this is our treatment group). As we have information about sick leaves for school years 2000-2001 to 2012-2013, the age of teachers in the control group runs from 51 to 65 years old, whereas the age of those in the treatment group goes from 48 to 62 years old.

Table 2 presents some descriptive statistics on teachers who are 53 years old, which is the youngest age available in all considered cohorts. Seniority measures the total number of years spent in the public sector. Besides, the salary grade informs on the teacher's career achievement on a scale from 1 to 12. As it directly determines the teacher's wage level, the higher the salary grade, the higher the wage. For instance, a teacher with a grade of 9 would earn approximately 2,100 euros (after taxes) per month, whereas the highest grade corresponds to a monthly net wage of 2,800 euros (in 2015). 90% of the teachers are employed full time at age 53. The average teaching load per week is 18 hours, which is also the mode of the distribution of the number of teaching hours.

Table 2 about here.

Other covariates reported in Table 2 describe some features of the working environment. The average class size is calculated as the ratio of the total number of students over the total number of teachers assigned to the school. However, as some teachers can split their working time between several schools, the total number of teachers in a school may be overestimated, and the ratio may be artificially low: in France, teachers in high schools are usually facing no less than 25 students in a classroom. This measure is nonetheless a proxy of the effective number of students a teacher is facing.

The average salary grade and the proportion of teachers above 40 in the school are proxies for the quality and the experience of the educational staff of the school. Besides, the “priority education zone” variable indicates if a teacher is working in a high school that benefits from additional means from the government, in order to cope with social and educational issues. In this type of schools, working conditions are said to be more difficult, and one can expect that teachers working there are more likely to take sick leaves.

Except for full time status, we observe some discrepancies between the control and treatment groups. It is interesting to consider that at the same age, the treatment group has a higher seniority, but a lower salary grade. In any event, this difference represents only a value of 0.05 on a scale varying from 1 to 12, which roughly corresponds to 10 euros a month. On top of that, since the average number of working hours per week in the treatment group is lower, while the proportion of full time teachers is the same, one can expect that teachers from the treatment group work less intensively than those in the control group (at age 53).

Whereas the proportion of married male teachers is comparable in the two groups, the proportion of married female teachers is slightly higher in the control group. However, this information must be carefully considered. In fact, the marital status is declared by the teacher to the administration: by default, it is set at single, and teachers have to produce administrative documents to update this information. Usually, they do it if the change in their marital status confers them some advantages as, for instance, some tax deductions, or possibilities to move to another school that is close to that of ones spouse. The discrepancy between these two ratios might as well reflect an administrative change, and does not necessarily jeopardize our identification strategy.

Since Table 2 presents descriptive statistics at age 53, it only reflects a snapshot of the reality. Appendix A depicts the evolution of mean differences from age 55 up to age 59.

One can observe that for treated cohorts, the average salary grade increases slower than for control cohorts. Moreover, teachers in treatment cohorts are more likely to work full time, and their average number of working hours per week is higher. They also face larger class sizes. Conversely, both groups work equally in priority education zones.

3.2 Data reliability

Research on health depends crucially on data quality. On the one hand, there is some controversy about the best indicators intended to measure health status: objective or subjective variables, weighted means of different indicators, etc. On the other hand, authors wonder how to get unbiased estimates, when data seem either "unreliable" (which is often the case for subjective indicators) or plagued with errors (like in administrative datasets). For instance, Bowers (2001) underlines how the greater or lesser degree of rigor in the treatment of information ultimately affects the quality of the statistical analysis.

In France, when a teacher is absent, the school reports it to the administration that is in charge of the schooling district (there are 26 different districts on the territory). By doing so, the school is directly asking either for a replacement teacher or for a financial compensation if another teacher in the school has to take the additional workload. There is a direct incentive for the school to notify the absence because some assistance will be set. At the end of the year, once the board has collected all the data concerning its district, it transfers them to the department in charge of statistics at the Ministry of Education, the DEPP. This decentralized process works if schools are indeed reporting efficiently absenteeism, which should be the case given their incentives. Conversely, one cannot really control whether the encoding process is fully accurate (the true reasons of sick leaves, the right number of days,...).

Another point that should be highlighted is the reform of the waiting period, *délai de* *carence*, that took place on January 1st, 2012. Before that reform, French teachers were paid from the first day they were absent from their school. The reform introduced a waiting period before compensation. In other words, the first day of absence was unpaid. This reform was finally canceled two years after (on January 1, 2014), when the government noticed that, when sick, teachers were taking more than one day off, or were not taking any day off until their health had so worsened that they had to take a long sick leave. Because the waiting period was fiscally not advantageous and politically unpopular, the government went back soon thereafter.

Therefore, one could imagine that during this period, the probability to take a short sick leave would drop. The reform was indeed intended to discourage this behavior. Consequently, the estimates of the impact of the 2010 reform on absenteeism due to short sick leaves could be underestimated. However, it only concerns the two last years of the panel, which corresponds only to 13% of the total number of teachers in our sample. [8]

4 Econometric analysis: Part 1

4.1 The econometric model

In order to confirm the descriptive statistics reported in Figure 2 that shows that the frequency of short sick leaves increases for the impacted cohort before their retirement,

[8]We introduced a 2012 year dummy in our regressions to take into account the possible effect of the 2012 reform of the waiting period (i.e., *délai de carence*). The associated coefficient was never statistically significant.

we propose to estimate the following linear probability model:

$$
\begin{aligned}
sick_{it} = \alpha + \gamma' X_{it} + \textstyle\sum_{s=55}^{59} \delta_s \mathbf{1}(age = s) + \textstyle\sum_{s=58}^{59} \delta_{s\_1951} \mathbf{1}(age = s, birth = 1951) \\
+ \textstyle\sum_{s=57}^{59} \delta_{s\_1952} \mathbf{1}(age = s, birth = 1952) + \textstyle\sum_{s=56}^{59} \delta_{s\_1953} \mathbf{1}(age = s, birth = 1953) \\
+ \nu_i + \varepsilon_{it}
\end{aligned}
\tag{1}
$$

where $sick_{it}$ is a binary variable taking the value 1 if teacher i takes at least one short sick leave during the schooling year t (0 otherwise), X_{it} is a vector of individual or school covariates presented in Table 2 (seniority, salary grade, number of hours per week, marital status, gender, etc.), ν_i is an individual fixed effect and ε_{it} is an error term potentially correlated over time. The set of binary coefficients δ_s captures the effect of age on the probability to take a short sick leave. These coefficients should increase with age: the older a teacher, the more likely she takes a sick leave. Coefficients of main interest are δ_{58_1951}, δ_{59_1951}, δ_{57_1952}, δ_{58_1952}, δ_{59_1952}, δ_{56_1953}, δ_{57_1953}, δ_{58_1953} and δ_{59_1953}, which are associated with interaction terms between the treatment (namely, being affected by the reform) and relevant age dummies for a given cohort. The three cohorts potentially impacted by the reform are teachers born in 1951 who had 58 years old in 2009 and 59 years old in 2010, teachers born in 1952 who had 57 years old in 2009, 58 years old in 2010 and 59 years old in 2011, and teachers born in 1953 who had 56 years old in 2009, 57 years old in 2010, 58 years old in 2011 and 59 years old in 2012. If these latter coefficients are positive and statistically significant, it means that postponing the minimum legal retirement age after 2010 onwards increased the probability of taking at least one short sick leave during a year before being allowed to retire. The coefficients δ_{58_1951}, δ_{57_1952} and δ_{56_1953} correspond to the behavior with respect to short sick leaves of those teachers affected by the reform one year before its implementation in 2010. According to our previous discussion (see section 2), these coefficients should be positive since those teachers expect in 2009 to be affected from 2010 onwards.

4.2 A linear probability model

The estimation of a linear probability model (LPM, hereafter) raises several well known issues: if the distribution of errors is not uniform, estimated outcomes might either exceed 1 or be lower than 0, which means they are biased and inconsistent. However, if the distribution of predicted outcomes is not extreme, namely if estimated values are essentially in the [0.3, 0.7] range, the conditional expectation of the binary outcome may be assumed to be linear, and a LPM may be used. In addition, Horrace and Oaxaca (2006) show that if only a few (or none) predicted probabilities fall outside the [0,1] interval, then the LPM estimates may be considered to be unbiased and consistent. Figure 3 presents the estimated density function of the conditional predicted probabilities (bars) given teachers' covariates. The short dashed line is the kernel estimator of this density. Our estimation shows that only 0.85% (i.e., less than 1%) of our observations do no meet the [0, 1] boundary constraint. The LPM is therefore a reliable benchmark.

Figure 3 about here.

4.3 Estimates

Beyond simple cross-section OLS estimation of the LPM model, some individual fixed effects, potentially correlated with the regressors, may be added to this model within a panel framework in order to exploit the longitudinal feature of our data. The first two columns of Table 3 report parameter estimates of the linear probability model (1), without (OLS) and with (FE) fixed individual effects. In order to circumvent drawbacks of the LPM model, a conditional logit model (CLM, hereafter) with fixed effects is also estimated.

Estimates of its average marginal effects are reported in the last column of Table 3.[9] In all tables reporting estimated coefficients (Tables 3 to 9), the number of observations is equal to the product of the number of teachers by the number of years during which they are observed (i. e., $N \times T$).

Table 3 reports estimates of the main parameters of these three models (LPM-OLS, LPM-FE, CLM) when applied to the individual probability to take at least one short sick leave during a school year. Table 3 shows that the probability to take at least one short sick leave during the school year is slightly increasing with age, from age 55 to age 59. According to parameter estimates of the CLM-FE model, the increase is rather smooth, going from 1 pp at age 55 to approximately 5 pp at age 59. For a given age, it is also possible to compare teachers affected by the reform with those belonging to the pre-reform cohorts. For instance, according to the CLM-FE model, at age 59, the probability to take at least one short sick leave during the school year increases by 2.6 pp for teachers born in 1951, by 5 pp for teachers born in 1952, and by 4.8 pp for teachers born in 1953. At age 58, this probability is increased by approximately 6 pp for teachers born in 1952 or 1953. Estimates obtained with the LPM-FE model are slightly lower. Thus, for the three cohorts impacted by the reform, the probability to take at least one short sick leave during the school year is increased after the reform, whatever the estimation method.

Teachers who are affected from 2010 onwards took more frequently short sick leaves in the year 2009 preceding the reform implementation: almost 6 pp more for those born in 1951, 4.3 pp more for those born in 1952, and 6 pp more for those born in 1953 (according

[9]Due to the incidental parameter issue, the fixed effect probit cannot be estimated. In fact, we must estimate N different fixed effects (with N tending to infinity) but we can only use a limited number T of periods for estimating each fixed effect. Estimates of these fixed effects are therefore inconsistent and contaminate other parameter estimates. One way to deal with the incidental parameter is to estimate a conditional logit model. Contrary to linear models, where demeaning is enough to get rid off individual fixed effects, non linear models do not offer such a shortcut. At the price of reducing the sample to individuals whose health status changes during the time span, one can use the conditional maximum likelihood estimator introduced by Chamberlin (1980) and get consistent estimates of parameters associated with time varying covariates.

to the estimates of the CLM model). Thus the anticipation effect is quite significant, and comparable to the effects estimated for the years after the actual reform implementation.

Estimated coefficients associated with covariates are not reported in Table 3, but are available upon request. The set of covariates includes the teacher's seniority in the current school year, her salary grade in the current school year, interaction terms between gender and marital status (married vs single), the proportion of teachers over 40 years old in the school, the average salary grade of teachers in the school, the average class size in the school, and a binary variable indicating whether the school is located in a priority education zone. Interaction terms between gender and marital status and the binary variable indicating whether the school is located in a priority education zone are fixed over the observation period. Consequently, they are omitted from the estimation of the two models with fixed effects.

Estimated coefficients associated with individual seniority and salary grade are statistically significant, but their sign is reversed when individual fixed effects are introduced. When fixed effects are incorporated, the probability to take at least one short sick leave during a school year increases with seniority but decreases with the salary grade. This could be due to an omitted variable bias. The positive (respectively, negative) bias affecting the salary grade (respectively, seniority) coefficient in the OLS model advocates for a positive (respectively, negative) correlation between unobservable characteristics and salary grade (respectively, seniority). One can think that teachers with the highest salary grade are those who worked the hardest during their career, committing themselves into administrative tasks and developing projects within the school on top of their own teaching activity. Once corrected for individual fixed effects, which may capture the unobserved individual level of effort, one observes that an increase of one scale point in the salary grade decreases by 0.15 pp the probability of taking at least one short sick leave during the school year. The probability to take at least one short sick leave during the school year increases by 1 pp when the number of worked hours increases by one unit: this estimate is statistically significant at the 0.1 percent level.

Coefficients associated with the interaction term between gender and marital status (married vs single) cannot be estimated for models with individual fixed effects because marital status is essentially time-invariant at these ages (55-60 years old). If we consider the linear probability model estimated by OLS (without fixed effects), we find that, compared to married males, single males are more likely to experience at least one short sick leave during the school year: their probability is 3 pp higher. For single (respectively, married) females, it is 16.3 (respectively, 11.4) pp higher. Thus, women tend to take more often short sick leaves than men. One can suppose that women either take more often care of their health status, or, as their spouse or parents get older, take more often care of their elderly relatives. Notwithstanding the gender, married teachers take less often short sick leaves, which is in line with the literature stating that married people are happier, less often sick, and less likely to indulge in risky behavior (see, e.g., Verbrugge (1979), Umberson (1987) and Goldman *et al.* (1995)).

The probability to take at least one short sick leave during the school year decreases significantly with the proportion of teachers above 40, but slightly (and not significantly) with the average salary grade of the teachers in the same high school. One can assume that the more homogeneous the educational staff, the more bonded: in that case, peer pressure could lift the threshold over which a teacher decides to take a short sick leave. Moreover, teachers whose working conditions are more demanding are more likely to be sick. The class size matters: when the teacher faces a larger class, her working conditions are more difficult because she has to maintain discipline, to speak louder, to face a higher level of stress, etc. For similar reasons, according to the LPM-OLS model, it seems that working in a priority education area also affects health status: in that situation, the probability to take a short sick leave during the year is increased by around 4 pp.

Table 3 about here.

In order to verify that the reform affects only the probability to take sick leaves during the school year, we reestimate the previous three models by using as outcome variable a dummy variable indicating whether the teacher takes at least one long sick leave during the school year. Table 4 shows that this probability is still increasing with age (between 55 and 59 years old), but to a much lesser extent than previously. Moreover, the age effect is statistically insignificant when using the conditional logit model with individual fixed effects. However, the main result is here that the estimated coefficients associated with interaction terms between age and variables indicating that the teacher belongs to a cohort affected by the reform (namely, teachers born in 1951, 1952 or 1953) are in general statistically insignificant (except in two cases out of 27) and very small. Thus estimates reported in Table 4 validate our main result.

Table 4 about here.

Another sensitivity test consists in choosing a placebo treatment group. In that exercise, the control group is composed of teachers born in 1945, 1946, and 1947, while the fake treatment group comprises teachers born in 1948, 1949, and 1950. These six cohorts (1945-1950) are not affected by the 2010 reform. For all the teachers born between 1945 and 1950, the statutory retirement age is 60 years old. The estimated model is the following:

$$
\begin{aligned}
sick_{it} = \alpha + \gamma' X_{it} + \textstyle\sum_{s=55}^{59} \delta_s \mathbf{1}(age = s) + \sum_{s=56}^{59} \delta_{s\_1948} \mathbf{1}(age = s, birth = 1948) \\
+ \textstyle\sum_{s=56}^{59} \delta_{s\_1949} \mathbf{1}(age = s, birth = 1949) + \sum_{s=56}^{59} \delta_{s\_1950} \mathbf{1}(age = s, birth = 1950) \\
+ \nu_i + \varepsilon_{it}
\end{aligned}
\tag{2}
$$

where, as previously, $sick_{it}$ is a binary variable taking the value 1 if teacher i takes at least one short sick leave during the schooling year t (0 otherwise), X_{it} is a vector of individual or school covariates presented in Table 2 (seniority, salary grade, number of hours per week, marital status, gender, etc.), ν_i is an individual fixed effect and ε_{it} is an error term potentially correlated over time. Let us remind that the set of binary coefficients δ_s captures the effect of age on the probability to take a short sick leave. Coefficients of main interest are δ_{56_1948} to δ_{59_1948}, δ_{56_1949} to δ_{59_1949}, and δ_{56_1950} to δ_{59_1950}, which are associated with interaction terms between the fake treatment and relevant age dummies for a given cohort. If these latter coefficients were positive and statistically significant, it would imply that the 2010 reform would have also increase the probability of taking short sick leaves before being allowed to retire for the 1948-1950 cohorts that are not directly affected by the reform. In this case, it would not be possible to conclude that the reform is the cause of the observed increase in the frequency of short-term sick leaves before retirement. Table 5 reports estimates of this model for the two panel specifications which are the most realistic ones (because they incorporate individual fixed effects). As expected, the age effect is statistically significant in both models. But the main result is that estimated coefficients associated with interaction terms between age and variables indicating that the teacher belongs to the 1948, 1949 or 1950 cohort, i.e., to the fake treatment group, are in general statistically insignificant (except in 4 cases out of 24) and very small. They are statistically significant and positive for teachers who are born in 1949 or 1950, and who are 59 years old (namely, one year before their statutory retirement age). This could be the consequence of the announcement of a probable increase in the statutory retirement age two years before the effective implementation of the reform. Despite this last result, the placebo test provides evidence that the increase in the frequency of short-term sick leaves in cohorts directly affected by the 2010 reform (i.e., for teachers born in 1951 and after) is mainly caused by the effective implementation of the reform.

Table 5 about here.

In the next section a background theoretical model is proposed. It explains how modifying the minimum retirement age increases the global labor disutility of a teacher and might induce her to take more frequent short sick leaves.

5 A background model

We develop a life-cycle model in which workers may choose endogenously their retirement age. This model, whose objective is to analyze the effects of an increase in the minimal age at retirement on this individual choice, is in the spirit of Chang (1991), Hazan (2009) or Kalemli-Ozcan and Weil (2010). Our model hinges on the basic idea that the share of teachers who take a sick leave not only increases with age, but also decreases with the individual's welfare. We thus study the consequences of an increase in the minimal age at retirement on welfare, and interpret a decrease in welfare as an increase in the probability to take a sick leave. Our theoretical results are presented below in a sequence of Lemmas whereas the testable implications of our model are given in three main claims.

Let $R \geq R_{\min}$ be the retirement age that is bounded by a legal constraint imposing a minimal age at retirement, denoted $R_{\min}$.[10] The instantaneous utility at date t of an

[10] It is easy to extend our framework by introducing a maximal age at retirement, but our main results would be unchanged.

agent of type i who is born at date b, depends on whether this agent is working or retired. It is given by the following functions:

$$\begin{cases} u\left(c\left(t,b\right)\right)-\phi_{i}\left(t-b\right) \text{ if } t-b\leq R, \\ \\ u\left(c\left(t,b\right)\right) \text{ if } t-b>R. \end{cases} \tag{3}$$

The function u is increasing and concave in $c\left(t,b\right)$, which stands for consumption at age $t-b$. The agent's specific function $\phi_{i}\left(t-b\right)$ is positive and increasing with age. We see from definitions (3) that, except age and cohort, the only difference across agents stems from the disutility of labor. The agent's welfare is evaluated with her intertemporal utility computed as the discounted flow of future instantaneous utilities. As it seems reasonable to assume that there is no welfare loss induced by the reform before it is announced by the Government, we compute below the welfare from the date of the announcement onward. Note that this intuition is reflected in the empirical part of our study in the distinction between sick leaves at a given age that are observed before the announcement from those that are observed after. Let a be the date at which the reform is announced (which can precede the date of implementation) and let us consider that the duration of life is deterministic and equal to $\omega>0$. Therefore, the welfare at date $a\leq b+R_{\min}$ of the agent i born at date b is given by:

$$\int_{a}^{b+\omega} e^{-\rho(t-a)}u\left(c\left(t,b\right)\right)dt-\int_{a}^{b+R} e^{-\rho(t-a)}\phi_{i}\left(t-b\right)dt \tag{4}$$

where ρ is the discount rate. We notice that consumption brings utility till the death of the agent that occurs at date $b+\omega$ while the disutility of work lasts till retirement that takes place at date $b+R$.

Non financial income depends on whether the agent is working or retired. It satisfies:

$$\begin{cases} w(t-b) \text{ if } t-b \leq R, \\ \\ \theta w(t-b) \text{ if } t-b > R, \end{cases} \tag{5}$$

were $w(t-b)$ is the wage received during the working period, which is assumed to be a non decreasing and concave function of age, i.e., $w'(t-b) \geq 0$ and $w''(t-b) \leq 0$. Moreover, $\theta w(R)$ is the pension received during retirement, which is assumed to be proportional, although lower, to the wage at retirement: $\theta \in (0,1)$. All those assumptions reproduce the wage and pension system of the population studied in the empirical part of our study. The intertemporal budget constraint equalizes the discounted flow of future consumption and the financial and human wealths. We assume that the interest rate equals the discount rate. At date a, the intertemporal budget constraint of an agent born at date b is thus given by:

$$\begin{aligned} & \int_a^{b+\omega} e^{-\rho(t-a)} c(t,b)\, dt \\ = \quad & x(a,b) + \int_a^{b+R} e^{-\rho(t-a)} w(t-b)\, dt + \theta w(R) \int_{b+R}^{b+\omega} e^{-\rho(t-a)} dt \end{aligned} \tag{6}$$

where $x(a,b)$ is the financial wealth, which is given.

The problem of the agent is to choose the age at retirement R and the consumption flow $c(t,b)$ for all $t \in [a, b+\omega]$ that maximize the welfare function (4) subject to the constraint (6), to the inequality $R \geq R_{\min}$ and for a given initial wealth $x(a,b)$. We are aware that our framework could be extended in many directions by including e.g. a risk of dying and possibly annuity markets, leisure decisions at the intensive margin and, ultimately, sick leave and health care decisions. We nevertheless think that this parsimonious model is sufficient to rationalize our empirical findings. Let us start with a technical statement:

Lemma 1. *There exists a unique solution to the optimization problem if* $\phi_i(a-b)=\varepsilon$, $\phi_i(\omega)=1/\varepsilon$, *with* ε *being small enough, and* $\theta > 1/1+e^{-(\omega-a)\rho}$.

Proof. See Appendix B.

The first two conditions ensure the existence of a solution by simply imposing limit conditions to the disutility function. Disutility should be small enough at date a to exclude the possibility for the worker to retire at the date at which the reform is announced (and be then consistent with our empirical analysis that consider employed individuals) and, symmetrically, be large enough at date $b+\omega$ to be sure that agents retire before they die. The third condition, which implies that the replacement rate should be large enough, is sufficient to guarantee the concavity of the optimization problem, and ultimately the uniqueness of the solution. It is simple to check whether this condition is satisfied in reality: if we consider an individual with a life expectancy at age $b+a$ that is equal to 40 years and a discount rate of 1%, the condition implies that θ should be larger than 0.6. Below, it is assumed that conditions stated in Lemma 1 are satisfied.

If it exists, the optimal solution can be constrained (with $R^* = R_{\min}$) or not (with $R^* > R_{\min}$). We can then establish a second technical result.

Lemma 2. *For date* a *and cohort* b*, there exists a positive and finite threshold, denoted* $\bar{\phi}(a,b,R_{\min})$*, such that the optimal retirement age equals the minimal retirement age if* $\phi_i(R_{\min}) \geq \bar{\phi}(a,b,R_{\min})$ *and is larger than this minimal age otherwise.*

Proof. See Appendix B.

Lemma 2 suggests that, within a cohort, heterogeneity of labor disutility allows to distinguish those who are constrained from those who are not, the former having the larger disutility. Differences in disutility can be explained by many factors, including health conditions, which are assumed to be exogenous to the model. A first message of our theoretical model is that only the teachers who are constrained could see their welfare being impacted by a reform consisting in an increase in $R_{\min}$. Conversely, teachers who are not constrained have a level of welfare that is independent from $R_{\min}$ and are, by definition, not affected by the reform that we consider. Therefore, a first test of the empirical validity of the model can be made by analyzing the differential effects of the reform on constrained agents (those who retire either during or at the end of the academic year during which they reach the minimal retirement age) and on unconstrained agents (those who retire during one of the following years). Let us start with a first theoretical claim which concerns the unconstrained agents. This claim can be tested empirically:

Claim 1. *Among individuals who are not constrained, there are no significant differences in the probability to take a sick leave between those in the treatment group and those in the control group.*

Let us now consider agents that are constrained and analyze the effect of a change in the minimal retirement age. Let us remark that our theoretical model considers a continuous increase in $R_{\min}$ whereas, in reality, the change is discrete. This implies that it may exist agents who were not constrained before the reform and which are constrained after. It would be simple to show that our results also apply to such a change.

Lemma 3. *An increase in $R_{\min}$ reduces the welfare of constrained agents.*

Proof. See Appendix B.

Agents that are constrained incur losses in welfare when the reform is announced. This comes from the fact that at age $R_{\min}$, they already experience a marginal disutility of working which is larger than the marginal gain in terms of consumption that is associated with a postponement of the retirement age. This translates into a testable proposition:

for them, increasing the retirement age should be associated with a higher probability to take a sick leave. This result can be summarized by the following testable claim.

Claim 2. *Among individuals who are constrained, the probability to take a sick leave is higher for those who are affected by an increase in the minimum statutory retirement age* $R_{\min}$.

We now investigate the effect of the reform on the share of individuals who are affected. We obtain this second result:

Lemma 4. *An increase in* $R_{\min}$ *implies an increase in the number of constrained agents within cohort* b.

Proof. See Appendix B.

An increase in the minimal retirement age is increasing the share of constrained individuals, i. e., those for which the reform is inducing a welfare loss. This result comes from the fact that the threshold defined in Lemma 2 is decreasing with the minimal retirement age. We summarize this in the claim below.

Claim 3. *The proportion of constrained individuals within a given cohort is higher in the treatment group.*

This last claim is validated by Figure 4 that reports nonparametric kernel estimates of the distribution of actual retirement ages for the different cohorts. The proportion of teachers actually retiring at the minimum statutory retirement age increases significantly for post-reform cohorts (teachers born in 1951, 1952, or 1953). The mode of each distribution is exactly positioned on the legal minimum retirement age (60 and 4 months for those born in 1951, 60 and 9 months for those born in 1952, 61 and 2 months for those born in 1953). The proportions of teachers from the different cohorts who retire during the school year when they reach the statutory minimum age for retirement are the following ones: 46% on average for control cohorts (1948-1950), 49% for teachers born in 1951, 53% for teachers born in 1952, and 60% for those born in 1953. These proportions constitute an additional empirical validation of Claim 3.

Figure 4 about here.

Claims 1 and 2, which were directly deduced from our theoretical model, are validated by the complementary econometric analysis which is reported in the next section.

6 Econometric analysis: Part 2

To check the empirical validity of Claims 1 and 2 of our theoretical model, we split the sample in two subsamples. Because most teachers often either prefer or have to finish the school year during which they reach the minimum retirement age [11], we have decided to differentiate substantially the subsamples of constrained and unconstrained teachers, and in particular not to consider as constrained teachers those retiring exactly at this statutory age. This would force us to consider only teachers leaving at the minimum age, which is measured in years and months for those born in 1951 and after: the subsamples thus obtained would be very small and we could not have the desired statistical power. Our idea was therefore to include in the group of constrained teachers those who retire, not exactly at the minimum age, but at the end of the school year during which they reach the minimum retirement age, either because they decide to do so or because they are required to do so by the school principal. Here are three examples: 1) a teacher born in December 1950 was 60 years old in December 2010 but she finished the 2010-2011 school year in June 2011, at the age of 60 years and 6 months; 2) a teacher born in September 1952 is 60

[11] Our data do not allow us to know whether teachers either decide or must finish the shooling year during which they reach the statutory retirement age.

years old in September 2012, and 60 years and 9 months in June 2013, which is the date on which she can retire; 3) a teacher born in December 1952 is 60 in December 2012, and 60 years and 9 months in September 2013: it is likely that her school principal will allow her to retire in September 2013 so as not to force her to work an extra year.

In order to address these difficulties as much as possible, without over-complicating the identification of constrained teachers, we have chosen to consider as constrained teachers those who retire before age 61 (to within one month). This group may not include teachers born at the beginning or at the end of 1953 (who could be obliged to retire at 61 years and 5 or 6 months for reasons of continuity of teaching). For the opposite reason, we have chosen to consider as unconstrained teachers those retiring after having 62 years old. This choice excludes teachers born between 1948 and 1950, for whom the statutory retirement age was 60, but who would have decided to work an additional year, up to 61 and a half for example. Our choice is therefore too "conservative", in the sense that the two subsamples that we consider could have had higher sizes. It is certainly imperfect, but it leads to a simple rule which nevertheless provides sufficient sizes.

With the first subsample, which contains the 17,863 teachers who retire before age 61, we test for the validity of Claim 2, which states that, among teachers who are constrained, the probability to take a short sick leave is higher for those who are affected by an increase in the statutory retirement age. This means that, for this first subsample, coefficients associated with interaction terms between the treatment (being affected by the reform, namely for those born in 1951, 1952 or 1953) and relevant age dummies should be significantly positive. Estimates reported in Table 6 confirm that claim: all corresponding estimates (denoted δ_{58_1951}, δ_{59_1951}, δ_{57_1952}, δ_{58_1952}, δ_{59_1952}, δ_{56_1953}, δ_{57_1953}, δ_{58_1953} and δ_{59_1953} in equation 1) are highly statistically significant and positive, varying from 3 to 6 pp according to the cohort, the year and the model.

Table 6 about here.

The second subsample is composed of the 3,122 teachers who retire after having 62 years old. As we said above, the choice of this age limit could appear to be drastic, but it assures us to keep in this subsample teachers born before 1951 who continue to work well after their statutory retirement age. This subsample allows us to test for the validity of Claim 1, which states that, among teachers who are not constrained, there are no significant differences in the probability to take a short sick leave between those in the treatment group and those in the control group. This claim would imply that all estimated coefficients that are denoted δ_{58_1951}, δ_{59_1951}, δ_{57_1952}, δ_{58_1952}, and δ_{59_1952} in equation (1), should be zero. In fact, their estimates reported in Table 7 are not statistically different from zero (while very imprecise), except 2 out of 15. Claims 1 and 2 of our theoretical model are then empirically confirmed, which reinforces its validity.

Table 7 about here.

To corroborate the causal interpretation of these results, we verify that the reform has not affected the probability of taking at least one long sick leave during a schooling year before retirement for these two groups of teachers (i.e., the so-called constrained and unconstrained teachers). For that purpose, we reestimate separately for each of these two groups the same models as before by using as outcome variable a dummy variable indicating whether the teacher takes at least one long sick leave during the school year. Estimates of the interest parameters are reported in Tables 8 and 9. Since the linear probability model with individual fixed effects (the second column in these two tables) is the most realistic model and the one concerning a greater number of observations, we concentrate our comments on the estimated coefficients of this model. For both groups of teachers, the age effect is statistically significant and positive, especially for the teachers retiring before age 61, which is consistent with the well-known relationship between health status and early retirement. However, the main result is here that the estimated coefficients associated with interaction terms between age and variables indicating that the teacher belongs to a cohort affected by the reform (namely, teachers born in 1951, 1952 or 1953) are in general statistically insignificant (except in two cases out of 18, but only at the 10 percent level of statistical significance) and very small. Consequently, it appears that the 2010 reform did not affect the probability of taking at least one long sick leave during a schooling year before retirement both for teachers retiring at the statutory retirement age and those retiring after that age. The only effect is the one concerning short sick leaves.

Tables 8 and 9 about here.

7 Conclusion

France, as many Western countries, has continuously reformed its public pension system over the last thirty years to cope with the aging of its population. Among possible institutional reforms, the increase in the statutory retirement age is the one that meet fiercest oppositions among the concerned workers. This might be due to the fact that, contrarily to other technical sustainability measures, the age at retirement is a parameter that is well understood by anyone. Demonstrations, strikes or political sparring are well known consequences of the announcements of pensions reforms. In this article, we provide a new one by showing that they may also reduce the labor intensity of older workers. Using a unique database of 38,652 French high school teachers that includes information on their sick leaves over the period 2000 to 2013, we were able to identify the effect of the announcement of a reform that postponed the statutory retirement age. The reform, which was enacted in 2010, concerned all cohorts born as of 1951 without modifying the rules for those who were born before.

Our results suggest that teachers of impacted cohorts were more likely to experience at least one short sick leave (of less than three months) than those of control cohorts. It is unlikely that this increase could be explained by some deterioration of teachers' health status. First, we show that, among impacted cohorts, only teachers that actually leave at the statutory retirement age increase the frequency of their short sick leaves before retirement. Second, we find no significant difference in the proportion of long sick leaves (of more than three months) between control and impacted cohorts. The increase in the frequency of short sick leaves for teachers impacted by the reform can therefore be interpreted as an increase in voluntary absenteeism caused by a decrease in utility, as described by a standard life-cycle model. The financial cost for the Government of this increase in short sick leaves (of less than three months) is hard to evaluate as we do not observe the exact number of days of absence during the school years. We know that the increase in the number of short sick leaves is associated with an increase in their average duration but, in any case, this additional cost should remain low in comparison with the financial gain induced by the reform.

This accounting is nevertheless only one dimension of the issue. Our study, which is one of the rare studies that evaluate the consequences of a change in the pension system on older teachers' behavior, can be extended in various directions. Absenteeism, especially if it reveals some demotivation, is quite detrimental for the quality of teachers' pedagogy and may therefore crucially affect students' achievement. This topic may be the subject of further research. This would shed some light on the link between teachers' working conditions and human capital acquisition.

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Appendix A: Additional Table

Table 1: Mean differences between control and treatment cohorts, by age

Age	55	56	57	58	59
Seniority (in years)	0.87***	0.76***	0.74***	0.72***	0.71***
Salary grade	-0.06***	-0.14***	-0.32***	-0.47***	-0.45***
Full-time work	0.01*	0.05***	0.06***	0.06***	0.06***
Number of teaching hours per week	-0.05*	0.69***	0.95***	0.91***	0.87***
Married male	-0.01	-0.01	-0.01	-0.01	-0.01*
Single male	0.01**	0.01***	0.01***	0.01***	0.01***
Married female	-0.03***	-0.03***	-0.03***	-0.03***	-0.02***
Single female	0.02***	0.03***	0.02***	0.02***	0.02***
Average class size	0.15***	0.15***	0.24***	0.41***	0.54***
Proportion of teachers over 40 in the school	0.00***	0.00***	0.00*	0.00***	0.01***
Salary grade of teachers in the school	0.13***	0.10***	0.01	-0.04***	-0.08***
Priority education zone	0.02***	0.00*	0.00	0.00	0.00

Notes: Standard errors in parentheses. Statistical significance levels are indicated as follows: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Appendix B: Proofs

Proof of Lemma 1. This proof follows the one developed in d'Albis *et al.* (2012). As the discount rate equals the interest rate, the optimal consumption is constant for all $t \in [a, b+\omega]$. Let us denote by $c(a,b,R)$ this optimal consumption for a given R. Using (6), it is defined by:

$$c(a,b,R) = \frac{x(a,b)\, e^{-\rho(a-b)} + \Psi(a-b,R)}{\int_{a-b}^{\omega} e^{-\rho z} dz} \tag{7}$$

where $e^{\rho(a-b)}\Psi(a-b,R)$ is human wealth at age $a-b$ for a given R and:

$$\Psi(a-b,R) = \int_{a-b}^{R} e^{-\rho z} w(z)\, dz + \theta w(R) \int_{R}^{\omega} e^{-\rho z} dz. \tag{8}$$

Interestingly, due to the additive separability imposed on the instantaneous utility function, the function $c(a,b,R)$ does not depend on i. We also notice that the impact of a change in R on the optimal consumption is proportional to the partial derivative of function Ψ with respect to R, which is denoted below Ψ'_R and which is given by:

$$\Psi'_R = e^{-\rho R}(1-\theta)\, w(R) + \theta w'(R) \int_{R}^{\omega} e^{-\rho z} dz. \tag{9}$$

Given that $w'(R) \geq 0$, Ψ'_R is positive and does not depend on $a-b$. Moreover, it is useful to compute the following expression:

$$\Psi''_{RR} + \rho\Psi'_R = \theta w''(R) \int_{R}^{\omega} e^{-\rho z} dz + \left[e^{-\rho R}(1-\theta) - \theta e^{-\omega\rho}\right] w'(R). \tag{10}$$

Given that $w''(R) \leq 0$, $\Psi''_{RR} + \rho\Psi'_R$ is negative for all $R \geq a$ if $\theta > 1/1 + e^{-(\omega-a)\rho}$.

Since consumption is constant, the intertemporal utility at date $a < b + R_{\min}$ can be written as a function denoted $v_i(a, b, R)$ that satisfies:

$$v_i(a, b, R) = u(c(a, b, R))\, e^{\rho(a-b)} \int_{a-b}^{\omega} e^{-\rho z} dz - e^{\rho(a-b)} \int_{a-b}^{R} e^{-\rho z} \phi_i(z)\, dz \tag{11}$$

where $c(a, b, R)$ is given by equation (7). The optimal retirement age, denoted R^*, is the solution for R that maximizes (11) subject to (7), (8) and $R \geq R_{\min}$. First, let us abstract from the latter constraints and derive the conditions for the existence of a unique solution that belongs to (a, ω). The first-order derivative of the objective function(11) with respect to R is given by:

$$\frac{\partial v_i(a, b, R)}{\partial R} = u'(c(a, b, R)) \frac{\Psi'_R}{\int_{a-b}^{\omega} e^{-\rho z} dz} - e^{-\rho[R-(a-b)]} \phi_i(R), \tag{12}$$

where Ψ'_R, defined by equation (9), is positive. Sufficient conditions for the existence of a solution that belongs to (a, ω) would then be: $\phi_i(a-b) = 0$ and $\phi_i(\omega) = +\infty$ (or, equivalently, as stated in the Lemma). The second-order derivative of the objective function (11) with respect to R is:

$$\begin{aligned} \frac{\partial^2 v_i(a, b, R)}{\partial R^2} &= u''(c(a, b, R)) \left[\frac{\Psi'_R}{\int_{a-b}^{\omega} e^{-\rho z} dz} \right]^2 + u'(c(a, b, R)) \frac{\Psi''_{RR}}{\int_{a-b}^{\omega} e^{-\rho z} dz} \\ &\quad + \rho e^{-\rho[R-(a-b)]} \phi_i(R) - e^{-\rho[R-(a-b)]} \phi'_i(R) \end{aligned} \tag{13}$$

which, using (12), can be written for $\partial v_i(a, b, R) / \partial R = 0$ as:

$$\begin{aligned} \left. \frac{\partial^2 v_i(a, b, R)}{\partial R^2} \right|_{\frac{\partial v_i(a,b,R)}{\partial R}=0} &= u''(c(a, b, R)) \left[\frac{\Psi'_2}{\int_{a-b}^{\omega} e^{-\rho z} dz} \right]^2 \\ &\quad + u'(c(a, b, R)) \frac{[\Psi''_{22} + \rho \Psi'_2]}{\int_{a-b}^{\omega} e^{-\rho z} dz} - e^{-\rho[R-(a-b)]} \phi'_i(R) \end{aligned} \tag{14}$$

which is negative provided that expression (10) is negative. $\square$

Proof of Lemma 2. As long as the objective function is concave with respect to R, the optimal solution R^* satisfies:

$$R^* = R_{\min} \text{ if } \left.\frac{\partial v_i(a,b,R)}{\partial R}\right|_{R=R_{\min}} \leq 0 \tag{15}$$

and $R^* > R_{\min}$ otherwise. We have:

$$\left.\frac{\partial v_i(a,b,R)}{\partial R}\right|_{R=R_{\min}} = u'\left(c\left(a,b,R_{\min}\right)\right) \frac{\Psi'_R\big|_{R=R_{\min}}}{\int_{a-b}^{\omega} e^{-\rho z} dz} - e^{-\rho[R_{\min}-(a-b)]}\phi_i\left(R_{\min}\right). \tag{16}$$

Hence, $\bar{\phi}\left(R_{\min}\right)$ is defined as:

$$\bar{\phi}\left(a,b,R_{\min}\right) = u'\left(c\left(a,b,R_{\min}\right)\right) \frac{\Psi'_R\big|_{R=R_{\min}}}{e^{-\rho[R_{\min}-(a-b)]}\int_{a-b}^{\omega} e^{-\rho z} dz}. \ \square \tag{17}$$

Proof of Lemma 3. The utility of a constrained agent can be written as:

$$v_i\left(a,b,R_{\min}\right) = u\left(c\left(a,b,R_{\min}\right)\right) e^{\rho(a-b)} \int_{a-b}^{\omega} e^{-\rho z} dz - e^{\rho(a-b)} \int_{a-b}^{R_{\min}} e^{-\rho z}\phi_i(z)\, dz. \tag{18}$$

Its first-order derivative with respect to $R_{\min}$, which is given in (16), is by definition negative (see equation (15)). $\square$

Proof of Lemma 4. To prove Lemma 4, we must show that $\bar{\phi}\left(a,b,R_{\min}\right)$ decreases with $R_{\min}$. Using equation (17), we have:

$$\frac{\frac{\partial\bar{\phi}(a,b,R_{\min})}{\partial R_{\min}}}{\bar{\phi}\left(a,b,R_{\min}\right)} = \frac{u''\left(c\left(a,b,R_{\min}\right)\right)}{u'\left(c\left(a,b,R_{\min}\right)\right)} \frac{\partial c\left(a,b,R_{\min}\right)}{\partial R_{\min}} + \frac{\frac{\partial \Psi'_{RR}\big|_{R=R_{\min}}}{\partial R_{\min}}}{\Psi'_R\big|_{R=R_{\min}}} + \rho. \tag{19}$$

We then use (9) and (10) to conclude. $\square$

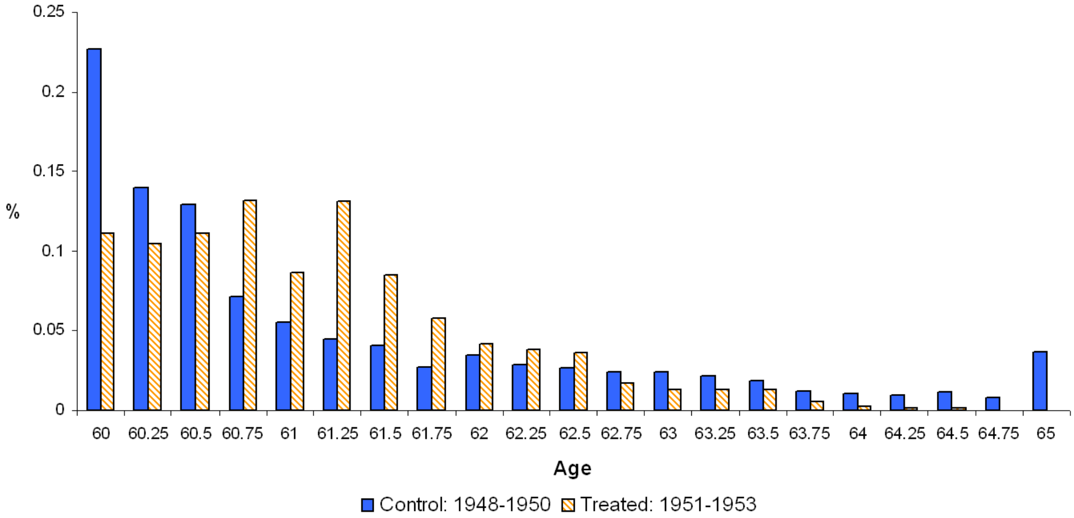
Tables and Figures

Table 2: Summary statistics for public high-school teachers aged 53

	Average value for 1948-1950 cohorts (Control)	Average value for 1951-1953 cohorts (Treatment)	Difference
Seniority	14.89	15.63	0.74***
	(0.070)	(0.070)	
Salary grade	9.02	8.97	-0.05***
	(0.012)	(0.014)	
Full-time work	0.90	0.90	0.00
	(0.002)	(0.002)	
Teaching hours per week	18.02	17.77	-0.25***
	(0.016)	(0.02)	
Married male	0.27	0.26	-0.01**
	(0.003)	(0.003)	
Single male	0.09	0.10	0.01***
	(0.002)	(0.002)	
Married female	0.42	0.40	-0.02***
	(0.003)	(0.004)	
Single female	0.22	0.23	0.01***
	(0.003)	(0.003)	
Average class size	12.43	12.27	-0.18***
(in the school)	(0.020)	(0.020)	
Proportion of teachers over 40	0.60	0.58	-0.02***
(in the school)	(0.001)	(0.001)	
Average salary grade of teachers	7.12	7.02	-0.10***
(in the school)	(0.004)	(0.005)	
Priority education zone	0.05	0.07	0.02***
	(0.001)	(0.002)	
Number of observations	22,320	16,332	

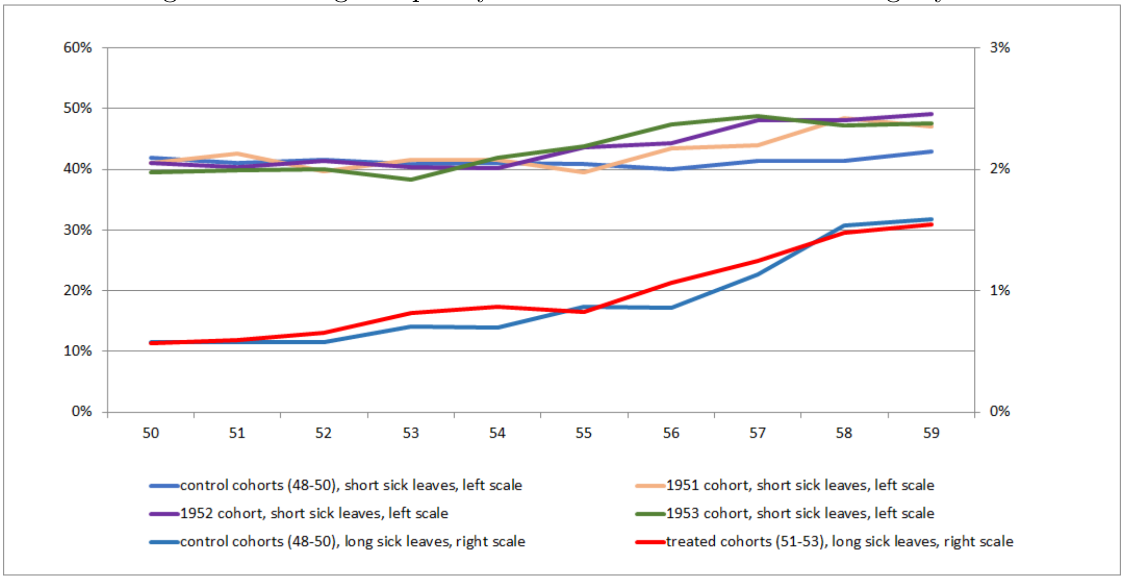
Notes: Standard errors in parentheses. Statistical significance levels are indicated as follows: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Figure 1: Distribution of the effective retirement age



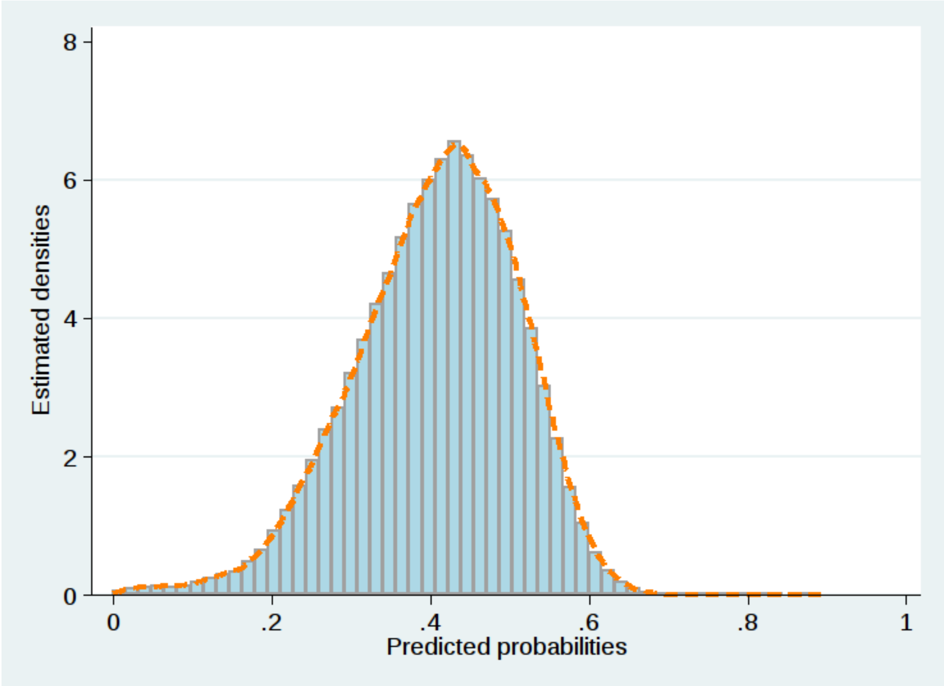
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Figure 2: Average frequency of at least one sick leave during a year



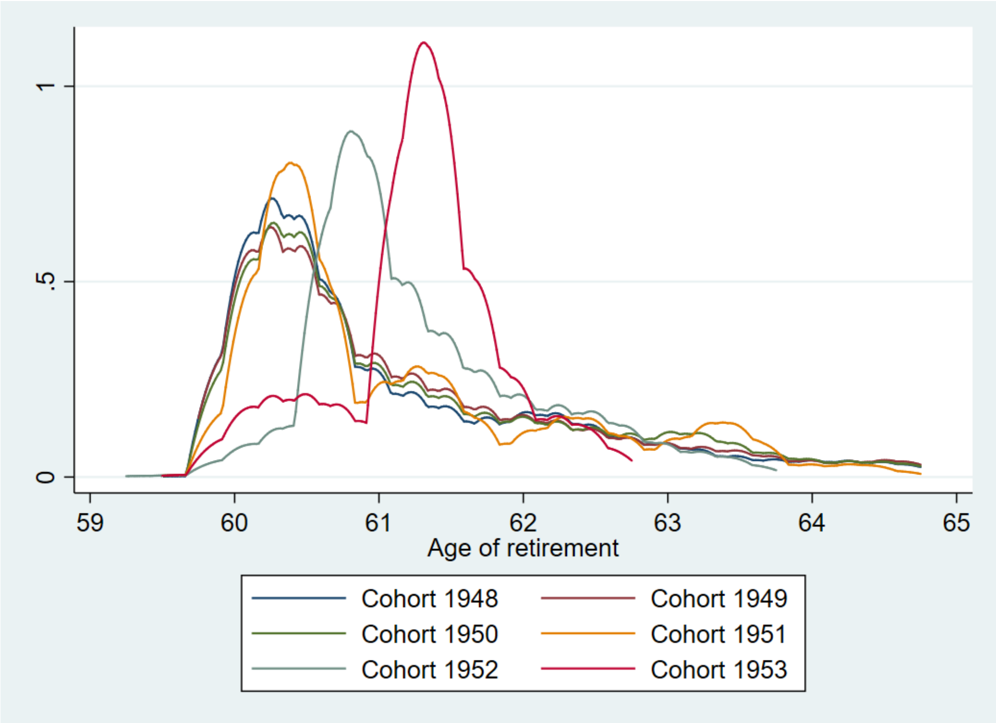
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Figure 3: Density of probabilities estimated with a linear probability model



Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Figure 4: Distribution of the actual retirement age, according to the birth cohort



Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Table 3: Main estimated parameters of the probability to take at least one short sick leave (of less than three months) during a school year

	LPM (Cross-sectional OLS)	LPM (Panel data)	CLM (Panel data)
Age=55	0.0189***	0.00982***	0.0115***
	(0.00311)	(0.00313)	(0.00383)
Age=56	0.0214***	0.0114***	0.0134***
	(0.00332)	(0.00380)	(0.00457)
Age=57	0.0317***	0.0207***	0.0242***
	(0.00366)	(0.00452)	(0.00567)
Age=58	0.0308***	0.0205***	0.0242***
	(0.00402)	(0.00528)	(0.00647)
Age=59	0.0528***	0.0414***	0.0479***
	(0.00410)	(0.00591)	(0.00797)
Cohort=1951, age=58	0.0663***	0.0536***	0.0589***
	(0.00875)	(0.00802)	(0.00975)
Cohort=1951, age=59	0.0365***	0.0241***	0.0263***
	(0.00883)	(0.00838)	(0.00907)
Cohort=1952, age=57	0.0482***	0.0388***	0.0434***
	(0.00769)	(0.00711)	(0.00850)
Cohort=1952, age=58	0.0649***	0.0535***	0.0599***
	(0.00788)	(0.00736)	(0.00912)
Cohort=1952, age=59	0.0544***	0.0459***	0.0506***
	(0.00836)	(0.00785)	(0.00926)
Cohort=1953, age=56	0.0543***	0.0541***	0.0617***
	(0.00791)	(0.00744)	(0.00930)
Cohort=1953, age=57	0.0574***	0.0536***	0.0609***
	(0.00805)	(0.00750)	(0.00940)
Cohort=1953, age=58	0.0601***	0.0570***	0.0638***
	(0.00868)	(0.00819)	(0.0100)
Cohort=1953, age=59	0.0390***	0.0439**	0.0478**
	(0.0149)	(0.0135)	(0.0156)
Intercept	0.213***	0.247***	
	(0.0194)	(0.0246)	
Covariates	Yes	Yes	Yes
Individual fixed effects	No	Yes	Yes
Akaike information criterion	368,411.2	245,808.5	175,716
Number of observations	266,831	266,831	214,685

Notes: LPM: Linear Probability Model, CPM: Conditional Logit Model. Robust standard errors in parentheses. Statistical significance levels are indicated as follows: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The set of covariates, whose estimated coefficients are not reported here, includes the teacher's seniority in the current school year, her salary grade in the current school year, interaction terms between gender and marital status (married vs single), the proportion of teachers over 40 years old in the school, the average salary grade of teachers in the school, the average class size in the school, and a binary variable indicating whether the school is located in a priority education zone. Interaction terms between gender and marital status and the binary variable indicating whether the school is located in a priority education zone are fixed over the observation period. Consequently, they are omitted from the estimation of the two models with fixed effects. The number of observations is equal to the product of the number of teachers by the number of years during which they are observed.
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Table 4: Main estimated parameters of the probability to take at least one long sick leave (of more than three months) during a school year

	LPM (Cross-sectional OLS)	LPM (Panel data)	CLM (Panel data)
Age=55	0.00182***	0.000784	0.00152
	(0.000533)	(0.000607)	(0.00236)
Age=56	0.00222***	0.00133*	0.00216
	(0.000589)	(0.000747)	(0.00301)
Age=57	0.00432***	0.00373***	0.00750
	(0.000719)	(0.000933)	(0.00572)
Age=58	0.00667***	0.00659***	0.0131
	(0.000874)	(0.00113)	(0.00893)
Age=59	0.00687***	0.00717***	0.0143
	(0.000885)	(0.00123)	(0.00990)
Cohort=1951, age=58	-0.000329	-0.000762	-0.00312
	(0.00154)	(0.00157)	(0.00406)
Cohort=1951, age=59	-0.000343	-0.00175	-0.00529
	(0.00191)	(0.00195)	(0.00499)
Cohort=1952, age=57	0.000634	-0.000884	-0.00304
	(0.00196)	(0.00197)	(0.00419)
Cohort=1952, age=58	0.000593	-0.000665	-0.00335
	(0.00176)	(0.00182)	(0.00403)
Cohort=1952, age=59	0.00199	0.000927	-0.00103
	(0.00195)	(0.00200)	(0.00379)
Cohort=1953, age=56	0.00352*	0.00247	0.00419
	(0.00164)	(0.00168)	(0.00458)
Cohort=1953, age=57	0.00180	0.000293	-0.00111
	(0.00170)	(0.00172)	(0.00387)
Cohort=1953, age=58	-0.00179	-0.00345*	-0.00988
	(0.00175)	(0.00180)	(0.00685)
Cohort=1953, age=59	0.00380	0.00209	-0.000377
	(0.00320)	(0.00337)	(0.00676)
Intercept	0.00728*	-0.0220***	
	(0.00354)	(0.00522)	
Covariates	Yes	Yes	Yes
Individual fixed effects	No	Yes	Yes
Akaike information criterion	-535,219.4	-587,990.7	7,591
Number of observations	266,831	266,831	14,701

Notes: Same notes as those of Table 3.
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Table 5: Main estimated parameters of the probability to take at least one short sick leave (of less than three months) during a school year: placebo test with a fake treatment group

	LPM (Panel data)	CLM (Panel data)
Age=55	0.00616	0.00780
	(0.00436)	(0.00548)
Age=56	0.01030	0.01300
	(0.00719)	(0.00893)
Age=57	0.02320***	0.02930***
	(0.00768)	(0.00982)
Age=58	0.02430***	0.03070***
	(0.00833)	(0.0106)
Age=59	-0.02020**	-0.02640**
	(0.00920)	(0.01070)
Cohort=1948, age=56	-0.00415	-0.00475
	(0.00991)	(0.01250)
Cohort=1948, age=57	0.00631	0.00767
	(0.01010)	(0.01250)
Cohort=1948, age=58	-0.00042	-0.00045
	(0.01020)	(0.01250)
Cohort=1948, age=59	0.01250	0.01760
	(0.01050)	(0.01290)
Cohort=1949, age=56	-0.00574	-0.00680
	(0.00959)	(0.01190)
Cohort=1949, age=57	-0.00018	-0.00025
	(0.00973)	(0.01190)
Cohort=1949, age=58	0.00937	0.01170
	(0.00979)	(0.01210)
Cohort=1949, age=59	0.02890***	0.03810***
	(0.01010)	(0.01290)
Cohort=1950, age=56	-0.00437	-0.00539
	(0.00971)	(0.01200)
Cohort=1950, age=57	0.00638	0.00786
	(0.00983)	(0.0122)
Cohort=1950, age=58	0.01500	0.01780
	(0.00997)	(0.01250)
Cohort=1950, age=59	0.04510***	0.05780***
	(0.01040)	(0.01400)
Intercept	0.25900***	
	(0.03920)	
Covariates	Yes	Yes
Individual fixed effects	Yes	Yes
Akaike information criterion	124,258.948	85,844.1
Number of observations	138,984	106,793

Notes: Same notes as those of Table 3.
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Table 6: Main estimated parameters of the probability to take at least one short sick leave (of less than three months) during a school year for teachers who retire before age 61

	LPM (Cross-sectional OLS)	LPM (Panel data)	CLM (Panel data)
Age=55	0.0226***	0.00976**	0.00897**
	(0.00418)	(0.00425)	(0.00417)
Age=56	0.0258***	0.0126**	0.0115**
	(0.00434)	(0.00508)	(0.00501)
Age=57	0.0428***	0.0285***	0.0261***
	(0.00470)	(0.00602)	(0.00677)
Age=58	0.0399***	0.0255***	0.0236***
	(0.00511)	(0.00702)	(0.00743)
Age=59	0.0538***	0.0380***	0.0349***
	(0.00520)	(0.00793)	(0.00899)
Cohort=1951, age=58	0.0668***	0.0522***	0.0451***
	(0.0114)	(0.0105)	(0.0107)
Cohort=1951, age=59	0.0376***	0.0229**	0.0193**
	(0.0115)	(0.0109)	(0.00958)
Cohort=1952, age=57	0.0570***	0.0429***	0.0372***
	(0.0110)	(0.0102)	(0.0101)
Cohort=1952, age=58	0.0638***	0.0485***	0.0421***
	(0.0111)	(0.0104)	(0.0105)
Cohort=1952, age=59	0.0695***	0.0591***	0.0508***
	(0.0119)	(0.0112)	(0.0115)
Cohort=1953, age=56	0.0445***	0.0409***	0.0380***
	(0.0133)	(0.0125)	(0.0121)
Cohort=1953, age=57	0.0375***	0.0296**	0.0281**
	(0.0135)	(0.0123)	(0.0118)
Cohort=1953, age=58	0.0628***	0.0611***	0.0554***
	(0.0142)	(0.0133)	(0.0136)
Cohort=1953, age=59	0.0546**	0.0712***	0.0641***
	(0.0247)	(0.0217)	(0.0214)
Intercept	0.0993***	0.178***	
	(0.0256)	(0.0351)	
Covariates	Yes	Yes	Yes
Individual fixed effects	No	Yes	Yes
Akaike information criterion	204,199.4	138,511.6	
Number of observations	147,605	147,605	120,454

Notes: Same notes as those of Table 3.
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Table 7: Main estimated parameters of the probability to take at least one short sick leave (of less than three months) during a school year for teachers who retire after age 62

	LPM (Cross-sectional OLS)	LPM (Panel data)	CLM (Panel data)
Age=55	-0.00228	-0.00927	-0.0143
	(0.00997)	(0.00963)	(0.0155)
Age=56	0.00555	-0.00434	-0.00691
	(0.00992)	(0.0110)	(0.0177)
Age=57	0.0129	-0.000637	-0.000934
	(0.0103)	(0.0128)	(0.0203)
Age=58	0.0250**	0.0134	0.0212
	(0.0113)	(0.0151)	(0.0234)
Age=59	0.0501***	0.0357**	0.0551**
	(0.0116)	(0.0171)	(0.0256)
Cohort=1951, age=58	0.0512**	0.0303	0.0446
	(0.0240)	(0.0216)	(0.0325)
Cohort=1951, age=59	0.0235	0.00400	0.00462
	(0.0240)	(0.0233)	(0.0328)
Cohort=1952, age=57	0.0380	0.0368	0.0559
	(0.0305)	(0.0289)	(0.0419)
Cohort=1952, age=58	0.0449	0.0433	0.0644
	(0.0321)	(0.0281)	(0.0423)
Cohort=1952, age=59	0.0541	0.0536	0.0770*
	(0.0352)	(0.0340)	(0.0463)
Cohort=1953, age=56	-0.0879	-0.0394	-0.0476
	(0.1330)	(0.1570)	(0.1690)
Cohort=1953, age=57	0.0519	0.1010	0.1330
	(0.1330)	(0.1550)	(0.1630)
Cohort=1953, age=58	0.1850	0.2310*	0.2850*
	(0.1250)	(0.1210)	(0.1630)
Cohort=1953, age=59	-0.0184	0.0421	0.0549
	(0.1430)	(0.1450)	(0.1940)
Intercept	0.506***	0.445***	
	(0.0643)	(0.0786)	
Covariates	Yes	Yes	Yes
Individual fixed effects	No	Yes	Yes
Akaike information criterion	34,330.5	22,123.6	
Number of observations	25,224	25,224	19,431

Notes: Same notes as those of Table 3.
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Table 8: Main estimated parameters of the probability to take at least one long sick leave (of more than three months) during a school year for teachers who retire before age 61

	LPM (Cross-sectional OLS)	LPM (Panel data)	CLM (Panel data)
Age=55	0.00241***	0.00154*	0.00345
	(0.00075)	(0.00087)	(0.00346)
Age=56	0.00238***	0.00209**	0.00397
	(0.00078)	(0.00106)	(0.00415)
Age=57	0.00489***	0.00530***	0.00962
	(0.00094)	(0.00132)	(0.00847)
Age=58	0.00753***	0.00872***	0.01470
	(0.00114)	(0.00160)	(0.01250)
Age=59	0.00694***	0.00866***	0.01510
	(0.00111)	(0.00176)	(0.01320)
Cohort=1951, age=58	0.00051	-0.00172	-0.00378
	(0.00261)	(0.00261)	(0.00437)
Cohort=1951, age=59	0.00092	-0.00086	-0.00389
	(0.00258)	(0.00261)	(0.00460)
Cohort=1952, age=57	0.00017	-0.00186	-0.00528
	(0.00223)	(0.00230)	(0.00525)
Cohort=1952, age=58	0.00495	0.00159	-0.00120
	(0.00324)	(0.00327)	(0.00392)
Cohort=1952, age=59	0.00694**	0.00562*	0.00189
	(0.00318)	(0.00322)	(0.00371)
Cohort=1953, age=56	0.00581*	0.00298	0.00214
	(0.00304)	(0.00309)	(0.00432)
Cohort=1953, age=57	0.00495	0.00159	-0.00120
	(0.00324)	(0.00327)	(0.00392)
Cohort=1953, age=58	0.00292	0.00039	-0.00414
	(0.00337)	(0.00339)	(0.00515)
Cohort=1953, age=59	0.00784	0.00686	0.00204
	(0.00579)	(0.00612)	(0.00711)
Intercept	0.00429	-0.0291***	
	(0.00500)	(0.00798)	
Covariates	Yes	Yes	Yes
Individual fixed effects	No	Yes	Yes
Akaike information criterion	-285,827.2	-315,881.1	
Number of observations	147,605	147,605	8,509

Notes: Same notes as those of Table 3.
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)

Table 9: Main estimated parameters of the probability to take at least one long sick leave (of more than three months) during a school year for teachers who retire after age 62

	LPM (Cross-sectional OLS)	LPM (Panel data)	CLM (Panel data)
Age=55	0.00687***	0.00683***	0.224
	(0.00196)	(0.00198)	(0.175)
Age=56	0.00448**	0.00484**	0.163
	(0.00177)	(0.00193)	(0.169)
Age=57	0.00403**	0.00487**	0.161
	(0.00177)	(0.00217)	(0.193)
Age=58	0.00286	0.00417**	0.152
	(0.00178)	(0.00211)	(0.212)
Age=59	0.01240***	0.01330***	0.361
	(0.00277)	(0.00314)	(0.335)
Cohort=1951, age=58	0.00528	0.00371	0.039
	(0.00512)	(0.00510)	(0.125)
Cohort=1951, age=59	-0.00562	-0.00629	-0.166
	(0.00524)	(0.00546)	(0.149)
Cohort=1952, age=57	0.01480*	0.01540*	0.329
	(0.00891)	(0.00874)	(0.236)
Cohort=1952, age=58	-0.00181	-0.00292	-0.102
	(0.00403)	(0.00377)	(0.249)
Cohort=1952, age=59	0.01720	0.01800	0.209
	(0.01210)	(0.01230)	(0.181)
Cohort=1953, age=56	-0.00993***	-0.01550	-2.819
	(0.00327)	(0.01210)	(317.3)
Cohort=1953, age=57	-0.00880***	-0.01490	-2.847
	(0.00318)	(0.01220)	(300.2)
Cohort=1953, age=58	-0.00746**	-0.01460	-2.860
	(0.00310)	(0.01210)	(348.9)
Cohort=1953, age=59	-0.00510	-0.01070	-5.197
	(0.02230)	(0.02150)	(252.9)
Intercept	0.00414	-0.00515	
	(0.01070)	(0.01520)	
Covariates	Yes	Yes	Yes
Individual fixed effects	No	Yes	Yes
Akaike information criterion	-52,414.0	-57,496.2	
Number of observations	25,224	25,224	1,265

Notes: Same notes as those of Table 3.
Source: Calculations by the authors from administrative files provided by DEPP (French Ministry of Education)