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Social insurance reform and absenteeism in Vietnam

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There has been very little research into sickness absence in Vietnam. This article reveals how the labor market might adjust to more generous illness benefits offered by the Vietnam Social Insurance Law. In the study, we combined the difference-in-differences and matching techniques and used data from four waves of the Vietnam Labor Force Survey. The results indicate that the generosity of the benefits positively links to both the incidence and the duration of sick leave. The article also sheds light on the mechanism underlying the relationship by disaggregating the effects on gender, income, and job tenure groups. The results support current government efforts to enhance the public social insurance system.

Key Practitioner Message*: • Sickness insurance affects sick leave incidence and duration; • Low-income workers and female workers are sensitive to sickness benefits reform; • There is a link between job tenure and sick leave*.

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Key words: social insurance, sickness absence, gender, job tenure, income, Vietnam

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Social security is considered to be a key component of social welfare in most nations worldwide. In 2013, Vietnam's Prime Minister approved the Strategy of Development of the Social Security Sector through 2020, with the overall objectives being modernizing, strengthening capacity, and improving the sector's quality and efficiency. To realize the strategy, the government has reformed the social insurance system, replacing the 2006 social insurance framework with new legislation enacted in 2014. Recently, the 7th Plenum of the 12th Ruling Party Central Committee on May 23, 2018, also considered further reforming the social insurance system as one of three main focuses of the nation (in addition to building a contingent of senior officials at the strategic level and reforming wage policies). A goal of the reform was to turn the system into a key pillar of the national social protection system. The government is currently developing a Master Plan on Social Insurance to increase insurance coverage and enhance social protection (Cunha, 2018).

A major change of the 2014 Social Insurance Law compared with the 2006 version was the increase in sickness benefits for compulsory social insurance card holders. This study evaluated the influence of the change in sickness benefits on labor force participation. Two different approaches were combined in the study: difference-in-differences (DID) and propensity score matching (PSM). The study concentrated on short-term sickness due to information availability of the data sets used – Vietnam Labor Force Surveys. The findings indicate that the legislation was associated with a rise in both the likelihood and the duration of sickness absence. Moreover, low-income workers and female workers responded well to the law reform. The findings also suggest that the relationship between sick leave and the reform was driven by low-income workers with 1 year of job tenure or longer. For robustness checks, attrition samples were used.

The study contributes to the literature in labor economics and health policy fields in several ways. *First*, it is the first attempt to investigate the link between a social insurance law change and labor market participation using a Vietnamese data set. Very few studies have explored this connection using an Asian context. A considerable body of literature has quantified the influence of health insurance in Vietnam on many outcomes (e.g., health status, health-related spending, health-care service utilization, and household consumption), but very little is known about the labor supply. The social insurance system in Vietnam was mentioned in some past works, but none of them concentrated on sickness benefits under a social insurance law reform, to our knowledge. *Second*, the anatomy of the link by gender, income, and tenure groups is another innovation of our research compared with other similar works. It has allowed us to discover the mechanisms underlying the working behaviors of the treatment group. *Third*, the combination of DID and PSM techniques to reduce bias is also an advantage of this study. The combination

has been significantly used in the literature on health insurance reform, but we found very little evidence for its use in sickness insurance reform. Our study's findings provide insights into recent changes in social insurance policies, which help to improve their qualification. It is also a contribution toward realizing the Resolution No. 28-NQ/TW on Reforming Social Insurance Policies adopted by the Ruling Party.

The next section presents the social insurance system and sickness benefits in Vietnam. The third section provides a short review of the literature. The fourth section presents the empirical methods and data. The fifth section presents the results and the last section discusses the findings and draws conclusions.

Social insurance system and sickness benefits in Vietnam

In Vietnam, the Social Security Agency is responsible for implementing public policies and managing the national social security fund. Three mandatory social security systems have been implemented in this country: social insurance, health insurance, and unemployment insurance. The primary purpose of the national insurance policies is to help people to minimize the risks of illness, occupational accidents, and aging. Participants in the national social insurance system can recover income losses that arise from work absence due to sickness or accidents. There are two types of social insurance: (i) Compulsory insurance, which covers illnesses, maternity leave, industrial accidents and occupational diseases, retirement, and death allowance; and (ii) Voluntary insurance, which covers retirement and death allowance. Workers usually join the social security system and have a compulsory insurance card after signing a labor contract for 3 months or longer. The salaries written in the labor contracts include these three types of insurance. During the period investigated in this study, mandated social insurance contributions were 8% for employees and 18% for employers. The contributions were computed on the base salary plus other fixed allowances. For workers with a labor contract term of less than 3 months, the insurance premium is included in their salary. However, they are responsible for paying their own social insurance or other insurance.

The first sickness benefits schemes were introduced in Vietnam in 1961 for the public sector. The reforms took place in 2005 (medical benefits), 2006 (social insurance), 2008 (health insurance), and 2014 (social insurance). Since 2006, workers with a compulsory social insurance card can receive a sick leave allowance during their recovery from illness or when they take care of sick children under 7 years of age. To claim the allowance, certification from a medical institution is required. The number of leave days depends on the how long the worker has been paying for social insurance premiums. However, the time is not more than 30 days per year for less than a 15-year duration, and 60 days for more than a 30-year duration (Social Security, 2016). Compared with the 2006 Social Insurance Law, the 2014 version contains some new provisions for insurance card holders. The daily sickness benefit is computed by dividing the monthly sickness benefit into 24 days instead of 26 days as per the previous insurance law. The benefit level for convalescence and rehabilitation after illness equals 30% of the basic salary instead of 25% as previously. If workers need long-term treatment, especially to continue treatment after the prescribed 180-day period, they are entitled to 50% of their salary. This entitlement is subject to social insurance contributions in the month before they leave if they have paid fewer than 15 years (instead of 45% as previously). However, the maximum period that employees are entitled after the 180-day period equals the period they have made social insurance contributions.

Literature review

Social insurance systems that insure sickness risks vary widely across countries. According to Earle and Heymann (2006), one hundred and forty-five countries worldwide provided sick pay for workers in 2006. All OECD countries except the United States, Canada, and Japan, have some form of universal sick leave program (Heymann, Rho, Schmitt, & Earle, 2010). Statutory sickness insurance is integrated into the social insurance system in many European countries. Employers are usually responsible for short-term sickness compensation, while insurance agencies are responsible for long-term sickness compensation. EU countries consider sick leave policies to be a top priority as such policies use a sizable financial portion of their social security systems (Coggon et al., 2013). According to a report released in 2011 by the OECD, its members used 1.9% of their GDP on illness benefits (Organization for Economic Cooperation & Development, 2011).

The international literature

There is abundant literature on European sickness insurance systems. Recent research related to the effects of sick pay policies on short-term sick leave in Europe includes Blazinska, Rusnak, and Kacmarikova (2017), Dale-Olsen (2014), De Paola, Pupo, and Scoppa (2014), Goerke and Pannenberg (2015), Hall and Hartman (2010), Puhani and Sonderhof (2010), and Ziebarth and Karlsson (2010, 2014). All of these studies imply that an expansion (contraction) of sick leave benefits has caused a

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positive (negative) influence on the likelihood or the duration of sick leave. Blazinska et al. (2017), for instance, investigated a policy change that raised sick pay in the Slovak Republic and pointed out that the average sick leave duration increased to 6.5 days per year as a consequence of the legislation. Similar research showed that this figure in Germany was 1 day (Ziebarth & Karlsson, 2014). Dale-Olsen (2014) stressed that the repeal of the Norwegian public sick pay policy in 2004 led to a 4- to 5-day reduction among male employees, but female employees did not significantly respond to the law change. Hall and Hartman (2010) suggested that a reduction of sickness benefits under Swedish social insurance legislation resulted in a fall of 36% in the incidence of sickness absence. Evaluating the impact of a reform that lowered sickness benefits and increased monitoring of workers' health status in Italy, De Paola et al. (2014) found a 5.6 to 3% reduction in the likelihood of absence in the public sector during the period 2007–2008.

The United States and Canada pay more attention to work-related sickness, injury, and disability. Therefore, compensation insurance (for work-related sickness and injury) and disability insurance are more prevalent in these countries than are nonwork-related sickness and injury insurance. There is no federally mandated sickness insurance in the United States. However, a handful of major cities (e.g., San Francisco, Washington DC, and Seattle) and states (e.g., Connecticut, California, Massachusetts, and Oregon) have introduced some type of sick pay mandate. Past studies exploring this topic, such as those by Ahn and Yelowitz (2015), Garthwaite, Gross, and Notowidigdo (2014), and Stearns and White (2018), have used difference-in-differences as the primary tool. Ahn and Yelowitz (2015), for instance, provided evidence of a modest negative impact (around −0.3%) of Connecticut's sick leave legislation on the rate of labor force participation.

Another relevant source for our study is literature on long-term absenteeism (e.g., Puhani & Sonderhof, 2010). Using German data, Puhani and Sonderhof (2010) found that a cut in sick leave compensation from 100% to 80% in 1996 drove a decline of 2–3% in long-term sickness. This article also gains from studies on other social policies (e.g., Nguyen, 2019; Nguyen et al., 2018; Olsson & Thoursie, 2015). Olsson and Thoursie (2015) found that a rise in the sickness replacement rate in Sweden in 1987 caused a surge in the number of sick days of 9.1 and 6.1% among spouses of treated men and treated women, respectively.

A few studies have investigated sick leave policies in Asian countries (e.g., Earle & Heymann, 2006; Eling, Jia, & Yao, 2015). However, most of them did not thoroughly analyze the effects of the policies on labor behaviors. Literature quantifying the labor impact of health insurance policies has been useful for our study. Chou and Staiger (2001), for instance, found a connection between the introduction of an insurance scheme for nonworkers and a 4% reduction of labor supply among Taiwanese wives. Liao and Taylor (2010) evaluated the effectiveness of Taiwan Health Insurance in 1995 and documented a 9.6–13.6% decline in the supply claimed by married female farmers.

The Vietnamese literature

The literature on social insurance and work behaviors in Vietnam is scanty but nevertheless provides some relevant insights. Past social insurance systems have been well described and scrutinized. Also, the connection between these systems and household welfare and related policies has been seriously discussed (e.g., Evan & Harkness, 2008; Nguyen, Nguyen, Nguyen, & Tran, in press). Le, Groot, Tomini, and Tomini (2019) is a rare study examining the link between a health program for the poor and labor supply. The authors found a reduction of between 5.2 and 5.8 hours worked among the treatment group following the program. Both Le et al. (2019) and Tran (2015) have provided helpful arguments on matching methods and covariates for labor analyses. However, our article is the first work to estimate the effects of social insurance reform (and sickness insurance reform) on labor force participation in Vietnam.

The subgroup analysis evidence

The literature on determinants of sickness absence is also of interest. Previous studies have investigated sick leave and income distribution and found a negative link between earnings and sick leave rates (Leaker, 2008) and the duration of the sick leave in some employee subgroups (Johansson & Palme, 2002). Income levels may cause both income effect and substitution effect: the income effect implies that higher earnings allow a person to reduce the volume of work to attain a given consumption level. The substitution effect means that the cost of absence is also high (Allebeck & Mastekaasa, 2004).

This study also benefits from the literature on the association between job tenure and sickness absence. Barmby, Ercolani, and Treble (2002), for instance, showed a higher rate of sickness absence among employees who had longer tenure. Gimeno, Benavides, Amick, Benach, and Martinez (2004) probed data from 15 EU countries and found that sickness absence was usually higher for employees in permanent employment than for temporary employees. The exception was those on full-time fixed-term employment.

Also data investigations from Italy have revealed a positive association between the likelihood of sickness absence and job tenure (D'Amuri, 2017; De Paola et al., 2014).

Concerning gender disparity, Barmby et al. (2002) discovered that the average gender difference in sick leave among eight European countries and Canada was 1.2%. The fact that women were more sensitive to sickness benefits expansion than were men was confirmed by other studies from Sweden and Italy (e.g., De Paola et al., 2014; Henrekson & Persson, 2004). De Paola et al. (2014) argued that the Italian social insurance reform in 2008 has affected men and women differently. For the public sector, the probability of sickness absence declined by 2.58 and 4.74% for male and female workers, respectively.

The methods used in the literature

The difference-in-differences (DID) approach is a major tool used in the above studies. Other approaches used to enhance estimation, or check robustness, or reduce bias include propensity score matching (PSM), zero-inflated negative binomial (ZINB) fixed effect, and Cox hazard. PSM has been substantially utilized in studies in labor economics and health policy, for example, to quantify the association between alcohol consumption and work absence (Balsa & French, 2010), between hepatitis C, labor supply, and sick leave (DiBonaventura et al., 2011), and between insurance status and the working behaviors of childless adults (Dague, DeLeire, & Leininger, 2017). Regarding sickness insurance, Ziebarth and Karlsson (2010, 2014) provided evidence on the impact of German sick pay policies in the 1990s on sickness leave, labor costs, and workers' health.

Summary of the literature review

In brief, researchers from Europe and the United States have made a great effort to gauge the influence of sickness insurance reform on sickness absence. Their findings are intuitive and quite consistent: There is a link between sickness benefits generosity (reduction) and sick leave, although an insignificant estimation has sometimes been detected. Meanwhile, the topic has been neglected in Asia, albeit there is evidence on the employment effects of health-related policies. In addition, the literature on the association between gender status, income distribution, work experience, and sickness absence suggests the need for examining the role of these individual and characteristics on the link between the benefits change and sickness absence. Concerning the identification strategy, differences-in-differences (DID) was the dominant strategy used. The literature review also pointed out that a combination of DID and PSM is a promising approach to investigating reform.

Empirical strategy and data

Empirical strategy

In this section, we introduce the DID technique. DID is a widely used approach in policy analysis. Scientific research usually requires causal relationship evaluation, but that is a challenge. A randomized controlled trial (RCT) is often recommended to address confounding problems in assessing causal effects. However, as a large-scale RCT is rare, non-randomized experiments, such as the introduction of new laws and regulations, become popular. A conventional DID study is a quasi-experimental analysis that quantifies the impact of the new legislation on various outcomes based on a comparison of treatment and control groups before and after the legislation's introduction (Wing, Simon, & Bello-Gomez, 2018). The DID technique can be used for both repeated cross-sectional and panel data. A typical formula for DID estimation is given by:

$$\mathrm{DID} = \left(\text{avg. } K_T\text{aft} - \text{avg. } K_T\text{bef}\right) - \left(\text{avg. } K_C\text{aft} - \text{avg. } K_C\text{bef}\right) \quad (1)$$

where avg. means average; K_T and K_C denote outcomes of the treatment and control groups; aft and bef represent the outcomes after and before the legislation's introduction, respectively. The equation for the DID estimation is as follows:

$$K_{it} = \lambda \mathrm{Time}_t + \mu \mathrm{Treated}_{it} + \tau \mathrm{Treated}_{it} \times \mathrm{Time}_t + \varepsilon_{it} \quad (2)$$

where K_{it} stands for outcomes of individual i at time t; Treated is a dummy variable for treatment group (equals 1 if an individual is treated, zero otherwise); Time is a dummy variable for time effect (equals 1 if the time period is after the introduction, zero otherwise). It is easy to show that the coefficient of interest τ (i.e., the interaction term of the two dummies) in the equation (2) reflects the DID estimation in the equation (1).

The key premise of the DID approach is the "common trend assumption." It is said that in the absence of legislation, the two groups exhibit common trends in the outcomes. This assumption is not directly testable. However, common pre-legislation trends can usually be considered as supporting the assumption's credibility. There are several ways to improve the possibility of the common trends, of which a combination of matching with DID on pre-trends is dominant and the approach chosen in this research. (More discussion on the combination is presented below). Other possible ways of supporting the credibility are

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graphical evidence and legislation characteristics analysis, presented in the subsection *"Sickness absenteeism over time"* of the next section (Wing et al., 2018).

The DID approach, however, has some drawbacks. For instance, when pre-legislation data are not available, researchers should use other methods to evaluate the impact and not DID. Additionally, the control and treatment groups should be comparable and have the same level of data availability. DID may not be appropriate if there are insufficient time periods to perform the credibility of the common trend assumption as well. Nevertheless, these limitations were not an issue for our study. Still, Ai and Norton (2003) suggested that when outcomes are binary, OLS regressions are better than logit and probit ones for a DID model, as the model does not work well with nonlinear regressions. That could explain the recent popularity of OLS in economic literature (including our work) when dealing with binary outcomes in DID models.

To apply a DID technique, in this study, we assumed that employees' sickness absence is a function of the time of social insurance law reform, compulsory insurance status, personal controls, and job-related controls. The treatment group in the study consisted of employees who held a compulsory insurance card. The control group consisted of other employees who did not have an insurance card. The parametric DID model to quantify the causal effect of the legislative change is similar to equation (2), except for the inclusion of exogenous variables:

$$K_{it} = \lambda T_t + \mu D_{it} + \tau D_{it} * T_t + \psi X_{it} + \rho_t + \phi_p + \varepsilon_{it} \quad (3)$$

where K_{it} stands for either incidence of illness absence or duration of illness absence (the number of working hours absent) for individual i in period t; T_t is the time effect; D_{it} is the dummy for the treatment group; $D_{it}*T_t$ represents the law change effect with τ being the coefficient of interest; ρ_t is time dummies that may affect absence; ϕ_p is regional dummies, accounting for permanent differences across six Vietnamese regions that control for the variance in the regional labor market; ε_{it} stands for unobserved heterogeneity with normal distribution and zero mean; and X_{it} includes personal characteristics (e.g., gender, age, education), and job-related characteristics (e.g., job tenure, firm types).

Matching DID approach

Equation (3) describes a parametric model for the outcome. However, we could also use nonparametric estimates as the functional form assumptions for parametric specification might be incorrect. Let us suppose that the individual causal effect is $K_{it}(1) - K_{it}(0)$ for pre- and posttreatment, where K_{it} is the outcome under control for worker i in period t: $t = 0$ and 1 for pre- and posttreatment, respectively. Since no one is treated at time $t = 0$, we have $D_{i1} = D_i$ with $K_{i0}(0) = K_{i0}$ and $K_{i1} = D_i*K_{i1}(1) + (1- D_i)*K_{i1}(0)$, for $D_i = 1$, we have $K_{i1} = K_{i1}(1)$. We consider the Average Treatment Effect on the Treated (ATT), a key estimator in impact evaluation studies. In this study, ATT measured the impact of the insurance law reform on workers with a compulsory card.

$$\begin{aligned} \text{ATT} &= E\left[K_{it}(1) - K_{it}(0) \,|X_i, D_i = 1\right] \\ &= E\left[K_{it}(1) \,|X_i, D_i = 1\right] - E\left[K_{it}(0) \,|X_i, D_i = 1\right] \end{aligned} \quad (4)$$

ATT is identified only if the outcomes from the two groups do not differ in the absence of intervention. In observational research such as this one, one has to adjust any disparity in mean outcomes for disparities in characteristics before initiating the treatment to reduce the bias in estimating the effects of the intervention (Blackwell, 2013). There are several dominant adjustment approaches. This study used the "propensity score matching" method.

Matching techniques have been developed in many previous studies. The identification and consistency of matching estimators are ensured only if the following assumptions are satisfied: (i) *Unconfoundedness*: the outcomes are independent of intervention, conditional on the covariates; and (ii) *Overlap or common support condition*: the probability of intervention lies in the interval zero to one: $0 < Pr(D = 1|X) < 1$. *Unconfoundedness* means that, in a large group of control workers, we have to match workers similar to the workers in the treatment group in all relevant characteristics (X). Since the match is limited if vector X is highly dimensional, we can use "propensity score," namely, the probability of being treated. Matching procedures deriving from it are called propensity score matching (PSM). PSM helps in reducing the matching dimensionality to a single dimension. Matching may be better than OLS because the common support condition emphasizes the comparison of comparable workers. Also, it helps to avoid misspecification, as it is a nonparametric technique (Grilli & Rampichini, 2011).

Data

The Vietnam Labor Force Survey (VLFS), a monthly data set compiled by the General Statistics Office since 2007, was chosen for this study. The VLFS is the only data set in Vietnam that contains rich information on labor behaviors. The surveys contain household data, including basic demography, migration status, vocational training skills, employment, work absence, type of firms, main business industries, types of labor contracts, social insurance status, income, current

and previous work, reasons for a work break, working hours and remuneration, unemployment and economic inactivity status, under-employment status, and work mobility. Each survey covers around 700,000 to 830,000 individual observations. (Wave 2017 includes 825,102 observations). The data are representative of urban, rural, and six regions. We chose waves 2014–2017 because before 2014 there was no information on compulsory insurance status, and the 2017 wave was the latest available at the time of the study. As the new insurance law was enacted in 2014 but took effect first on January 1, 2016, the choice led to a balanced sample, with 2 years of pre-reform and two-years of post-reform. VLFSs have short panel data for each year. However, as the panel structure is not suitable for the DID technique due to the effective date, we considered the whole sample as cross-sectional data.

Table 1. Variable descriptions.

Variable	Description
Outcome variables	
Incidence of sickness absence (D)	Dummy variable, equal 1 if a worker worked less than usual in the last 7 days; and zero otherwise
Hours of sickness absence	Number of hours that a worker worked less than usual in the last 7 days due to sickness; and zero otherwise
Personal characteristics	
Age	Age of worker
New migration (D)	Dummy variable, equals 1 if migrating into the town under 12 months
Female (D)	Dummy variable, equals 1 if being woman
Child aged 0–5	Number of children from 0 to 5 years old
Log education	Log of education levels[a]
Job characteristics	
Job tenure	Duration worked for the current organization[b]
Fulltime (D)	Dummy variable, equals 1 if working at least 35 hours/week
Cooperative (D)	Dummy variable, equals 1 if working for a cooperative
Log regional unemployment rate	Log of annual regional unemployment rate
Other useful variables	
Compulsory social insurance (D)	Dummy variable, equals 1 if being compulsory SI cardholder
Registered organization (D)	Dummy variable, equals 1 if working for a registered organization
Actual hours worked (last week)	Actual hours worked in the last 7 days
Usual hours worked (a week)	Usual hours worked in a week
Monthly income	Monthly earning of worker
Farm household (D)	Dummy variable, equals 1 if working farm household business
Self-employment (D)	Dummy variable, equals 1 if being self-employed
Army force (D)	Dummy variable, equals 1 if working in army force

Note: (D): dummy variable; [a]Eight educational levels: 1. Never attended school, 2. Not completed primary school, 3. Primary school, 4. Lower secondary school, 5. Upper secondary school, 6. Mid-term professional school, 7. Professional college, 8. University and above; [b]Five duration levels: 1: under 6 months; 2: 6 months–1 year; 3: 1–5 years; 4: 5–10 years; 5: More than 10 years.

VLFS data sets allow us to identify whether an employee has worked less than usual because of sickness and the number of fewer hours. Table 1 describes the variables used for investigation. The main outcome variables were incidence of sickness absence (i.e., a dummy variable equals 1 if an employee worked less than usual or did not work at all in the last 7 days due to sickness; zero is used to indicate other reasons or work at least as usual); and duration of sickness absence (i.e., the number of hours that a worker worked less than usual in the last 7 days due to sickness. This number is zero if he or she worked at least as usual or less than usual, or did not work, due to reasons other than illness). This definition of outcome variables is similar to that of D'Amuri (2017).[1] Missing observations were not included in the final sample. (The percentages of missing values were 3.9, 3.9, and 4.5% for incidence, duration of sickness absence, and compulsory social insurance, respectively).Consistent with previous literature, explanatory variables included personal characteristics (age, new migration status, female dummy, the number of children aged under 5, and educational attainment) and job characteristics (job tenure, full-time work dummy, the dummy for work for a cooperative, and a log of the regional unemployment rate).

Other variables were useful for setting restrictions, splitting the sample, or checking robustness. A compulsory insurance dummy, for instance, helped to determine the treatment group compulsory insurance card holders. The control group included those without an insurance card. We restricted our final sample to persons in the labor force, residing in Vietnam during the survey period, and usually working at least 1 hour a week. Also, our sample did not include people who were working in the agriculture/forestry/fishing sector. As the law reform took effect in January 2016, we excluded the observations for this month to avoid contamination. Furthermore, the sample did not include very young workers (under 15 years old) or those above the retirement ages (i.e., 55 and 60 years old for women and men, respectively, according to the current Vietnam Labor Code). Appendix Table A1 presents summary statistics of the variables.

In the next section, we first illustrate the trends of the likelihood of sick leave and the average duration of

[1] Other reasons include schooling/training, personal matters, holidays/leaves/vacation, starting new work, shift work, off-season, starting/ending/changing job, leaving for family reasons. Similar to D'Amuri (2017), our estimations do not include observations of which workers did not work or worked less than usual for reasons they could not control (i.e., being laid off, production temporarily halted, bad weather, lack of clients/orders, job loss, loss of farming land, strikes, opening/remodeling a business unit, and other uncontrolled reasons). Sickness in this study included pregnancy-related illness.

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sick leave overtime using raw data. Then, we use the matching technique to reduce the bias of the raw sample. The PSTEST function of STATA checks balance quality. In all subsequent estimations, we applied the DID technique on the matched sample (i.e., the sample excluded observations that were not in the common support area). We estimated the reform impact on sick leave duration using the whole sample and the sample disaggregated by gender. We then gauged the effects of the legislative change on different income and job tenure groups. Next, we repeated the DID estimates but now using the incidence of sick leave rather than the duration of sick leave as the dependent variable. Finally, we checked for robustness by attrition samples.

Estimated results

Sickness absenteeism over time

The basic setup for DID analyses requires the common time (or parallel) trend assumption. That is, in the absence of the 2014 insurance law, the difference between the two groups should be unchanged over time. We could not test this assumption directly, but we argue in the study that it is probably true. First, as the 2014 law was enacted at the national level, enhanced previous legislation, and targeted a large group of workers (i.e., compulsory insurance card holders), it was exogenous to a change in absenteeism (Besley & Case, 2000). Second, the figure of

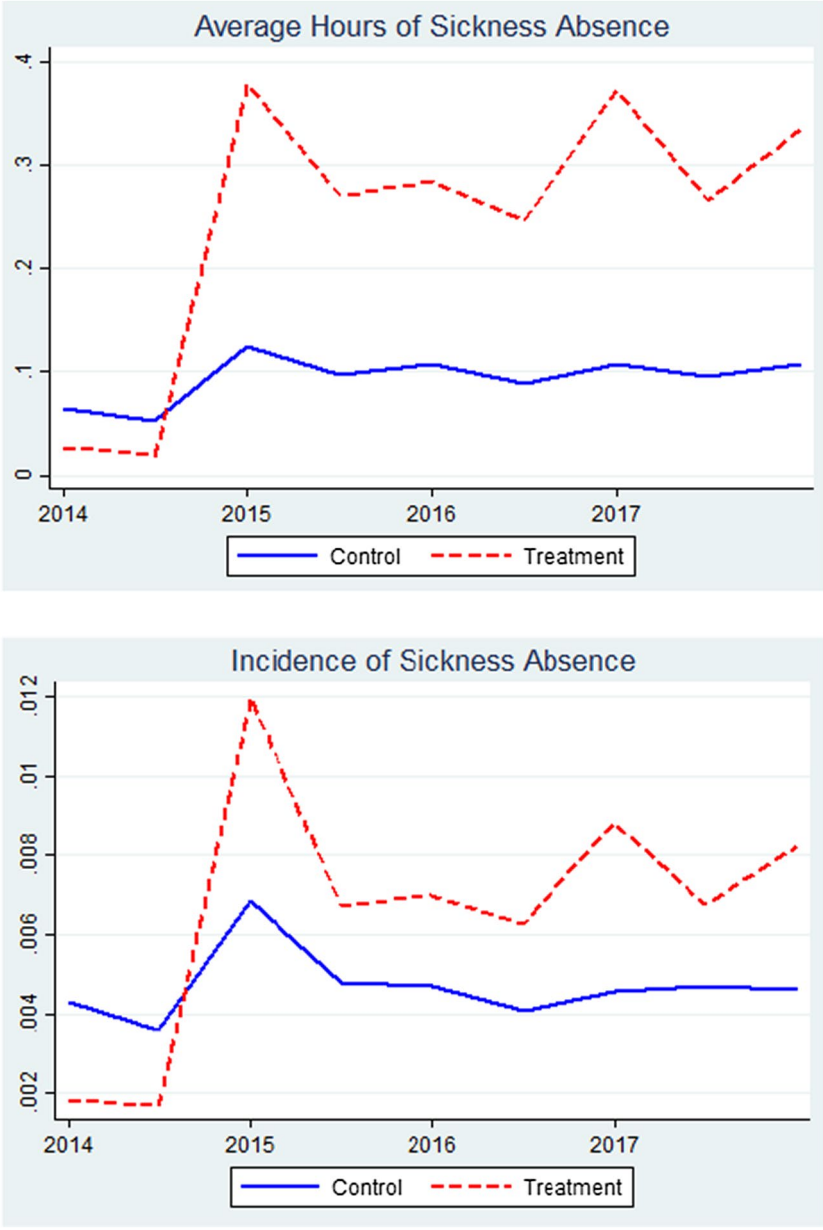

Figure 1. Sickness absenteeism over time by treatment group.

sickness absenteeism over time may provide another proof for the assumption. Third, by checking robustness with the attrition of the labor market, we might exclude contamination effects.

Figure 1 illustrates the incidence and average hours of sick leave for both control and treatment groups. The average weekly hours of sick leave reached the bottom in the middle of 2014 (0.05 hours and 0.02 hours for control and treatment groups, respectively). They reached a peak 6 months later (0.12 hours and 0.38 hours, respectively). Concerning the incidence of sickness absence, the lowest and the highest levels occurred at the same time as those of the duration of sick leave. The incidence was lower in 2014 for both control and treatment groups, probably due to changes in the survey questionnaire. The question relating to "worked less than usual" in the waves 2015–2017 had more options than in 2014. Therefore, the respondents had more opportunity to answer rather than skip the question, which led to a higher likelihood of sick leave. We see remarkably parallel curves in the 6 months before the law reform in both the upper and the lower half of the figure. After the reform took effect in January 2016, both curves had run parallel until the middle of 2016. (Note that the figure is based on raw data). These facts probably support the common time trend assumption.

Matching quality

As mentioned in the previous section, we combined DID with a matching technique to evaluate the influence of the insurance law change. We obtained propensity scores with balancing property by using logistic regression models controlling for personal and job characteristic variables, which are described in Table 1. One way to evaluate matching quality is by checking the balance for each covariate in the matching model. The most frequent indicator used to test the balancing is the percentage bias, namely, the percentage disparity between the treatment and control groups. We may calculate the bias for each covariate prior to and after the matching to estimate how the matching reduced the imbalance in the raw sample. Caliendo and Kopeinig (2008) suggested that the balancing quality is acceptable for values of the bias after the matching of less than 5%. The common practice is to obtain a bias of 5% or lower for as many variables as possible and a satisfactory bias for the remainder.

Table 2 displays PSTEST results for raw and matched samples, using 1-to-1 nearest neighbor matching with caliper 0.395 and non-replacement. Both the control and treatment groups in the raw sample differed for many covariates. The covariates with the most significant differences in the raw sample included the log of education, the number of children aged under 5, job tenure, female dummy, and the log of the regional unemployment rate. After matching, the difference was good for all covariates (i.e., percentage bias after matching <5%). The average percentage bias of the matched sample became 0.6% compared with 28.4% for the raw sample. $p > \text{Chi}^2$ of 0.151 implied no significant mean difference between the control and treatment groups in the matched sample. Other indicators, Pseudo R^2, B, and R, were also good. Thus, the matching performance was satisfactory.

The effects on absenteeism

Next, we analyzed the impact of the insurance law change on absenteeism using a matched DID approach. Namely, applying DID in the matched sample (i.e., the sample excluded observations that were not in common support). All estimations were based on equation (3). Table 3 shows the results of estimating the number of absence hours disaggregated by gender. In all columns, we controlled for personal and job characteristics. The columns in Panel A did not include time and regional fixed effects or the log of monthly income. In contrast, the columns in Panel B included them all. Table 3 shows that there were significant increases by 5.9 and 6.8% absence hour for mixed-gender and female samples, respectively, (equivalent to 3.5 and 4 minutes per week, respectively) for Panel A. Male workers did not significantly respond to the legislative change. Table 3 also shows that the results were not significantly changed when the fixed effects and log of income (see the last three columns) were included.

The estimations in Table 4 were designed similar to those of panel B of Table 3 (i.e., all columns control for personal and job characteristics, time and regions fixed effects, and the log of income). Table 4 displays the estimated results for two subsamples: (i) workers with a monthly income lower than the announced mean income (AMI - VND 5.5 million), and (ii) those with an income equal to AMI or higher[2]. We found a significant consequence related to the mixed-gender and female groups with a monthly income lower than AMI only. The logic here is that those with lower incomes might pay more attention to an increase in illness benefits. Alternatively, people with higher income may not be interested in such a small benefit, or they may try to attend work despite health problems in order to keep a well-paying job. Again, male workers were not sensitive to the

[2] AMI is the average monthly income announced by the government at the beginning of 2016. AMI is used in many public policies. AMI is VND 5.53 million for 2015, equivalent to VND 5.5 million at the base year 2014 of this study (Mai, 2016).

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Table 2. PSTEST unmatched and matched samples.

	Unmatched sample			Matched sample		
Variables	Treated mean	Control mean	%bias	Treated mean	Control mean	%bias
Personal characteristics						
Age	37.831	38.943	−10.6	37.023	37.092	−0.7
New migration (D)	0.0168	0.0139	2.4	0.0189	0.0191	−0.2
Female (D)	0.4991	0.3972	20.6	0.4675	0.4657	0.4
Child aged 0–5	0.3125	0.2221	17.3	0.2595	0.2645	−1.0
Log. education	1.7961	1.3098	141.0	1.5114	1.5104	0.3
Job characteristics						
Job tenure	4.0159	3.5894	40.4	3.7530	3.7513	0.2
Fulltime (D)	0.8725	0.8914	−5.9	0.9090	0.9103	−0.4
Cooperative (D)	0.0022	0.0046	−4.0	0.0034	0.0030	0.6
Log. of reg. unemployment rate	−4.0058	−3.9321	−13.9	−3.9573	−3.9500	−1.4

Sample	Ps	R2	LR chi2	$p > chi2$	Mean Bias	Median Bias	B	R	%Var
Unmatched	0.313		107235.29	0.000	28.4	13.9	148.7[a]	0.94	100
Matched	0.000		13.26	0.151	0.6	0.4	2.3	1.17	80

Note: (D): dummy variable; [a]If B > 25%, R outside [0.5; 2].

Table 3. Estimation on duration of sickness absence, by gender.

	Panel A			Panel B		
Variables	All	Women	Men	All	Women	Men
DID	0.059***	0.068***	0.019	0.058***	0.065**	0.019
	(0.014)	(0.025)	(0.013)	(0.014)	(0.025)	(0.013)
Personal characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Work characteristics	Yes	Yes	Yes	Yes	Yes	Yes
Log. of monthly income	No	No	No	Yes	Yes	Yes
Time dummies	No	No	No	Yes	Yes	Yes
Regional dummies	No	No	No	Yes	Yes	Yes
Number of observation	507,814	222,830	284,984	507,814	222,830	284,984
R_squared	0.022	0.031	0.17	0.023	0.032	0.018

Note: All columns control for personal and job characteristics. Columns in Panel B also control for time fixed effects, regional fixed effects, and log of income. The estimates use 1-to-1 nearest neighbor (with caliper 0.395, non-replacement) for matching. 1st number: Elasticity, 2nd number, in parentheses: Standard error. Standard errors are adjusted for clustering of personal identifiers; *Significant at $p = 10\%$; **Significant at $p = 5\%$; ***Significant at $p = 1\%$.

Table 4. Estimation on duration of sickness absence, by gender and income.

	<VND 5.5 mill.			≥VND 5.5 mill.		
Variables	All	Women	Men	All	Women	Men
DID	0.129***	0.164***	0.025	0.002	−0.008	0.013
	(0.019)	(0.032)	(0.017)	(0.018)	(0.044)	(0.018)
Number of observation	342,522	162,405	180,117	165,292	60,425	104,867

Note: VND 5.5 million is the announced mean income (AMI). All columns control for personal and job characteristics, log of income, time fixed effects, and regional fixed effects. The estimates use 1-to-1 nearest neighbor (with caliper 0.395, non-replacement) for matching. 1st number: Elasticity, 2nd number, in parentheses: Standard error. Standard errors are adjusted for clustering of personal identifiers; *Significant at $p = 10\%$; **Significant at $p = 5\%$; ***Significant at $p = 1\%$.

reform, no matter their income level. Hence, we have not reported the estimated results for the male group or groups with an income higher than AMI in the next analyses.

All estimated results are consistent with our expectation of a link between the reform and sick leave duration. Estimates were robust when time and regional fixed effects were included and quite close to the raw DID estimations. A handful of past studies have also examined sickness absence together with income distribution (e.g., Allebeck & Mastekaasa, 2004; Johansson & Palme, 2002; Leaker, 2008). Their arguments can partly explain the results of this study. The fact that men are less sensitive to sickness policy

changes than women is intuitive as men are usually the main breadwinner in Vietnamese families. Housework and childcare are mostly the wife's responsibility while the husband is engaged in the labor market. If a young child gets sick, the wife becomes the caregiver while her husband works for the family's income (Nguyen, 2013).

Concerning job tenure, the current Vietnam Labor Code stipulates that an employee working for an organization for at least 12 months is entitled to fully paid annual leave. Besides the annual leave, an employee is entitled to receive 1 day extra for every 5 years of continuous employment with the same organization. On the contrary, although the national labor policies make a clear distinction between annual leave and sick leave, in practice many employers do not comply with the policies and consider their employees' sick leave as annual leave. In these cases, the Vietnam Social Security Agency still pays for the employees' sick leave while their number of annual leave days is reduced following their absence duration. Hence, it is useful to evaluate the influence of the reform on job tenure.

The first row in Table 5 shows results from the whole sample. The second row indicates those with an income lower than AMI, and the last row exhibits women with an income lower than AMI. Table 5 shows significant tenure effects on workers in the last three tenure groups. The signs of estimations were consistent among the columns. The findings imply that the relationship between sick leave and the legislative reform was mostly driven by those employed for 1 year or more and having a monthly income lower than VND 5.5 million.

The association between job tenure and sickness absence has been frequently cited in previous literature (e.g., Barmby et al., 2002; Benavides, Benach, Diez-Roux, & Roman, 2000; D'Amuri, 2017; Gimeno et al., 2004). Nevertheless, this study addresses a different question and sheds light on the effects of a law change by varying tenure groups.

Finally, the incidence of work absenteeism was taken into consideration. Table 6 is a combination of Tables 3 and 4 but odds of sickness absence are used as the dependent variable. The findings presented in Table 6 indicate that the effects were minimal, but significant for both genders. The findings are in line with those in Tables 3 and 4, except for the male group. However, the estimations were not imprecise for those with an income higher than AMI. Thus, we analyzed the tenure effect for both genders in the next table. Table 7 disaggregates the reform effects by job tenure, income, and gender. We found that the labor behaviors of those with 1 to 5 years of service at their current organization ware linked to the legislative change, regardless of gender. For other tenure groups, estimations were imprecise for men. Still, workers with tenure longer than 10 years were sensitive to the legislation, except for the male group.

Taken together, the estimated results for the incidence of absence were consistent with those for the duration of absence in terms of the sign and the significance, especially for the mixed-gender and female samples. Sickness absenteeism among those with a compulsory insurance card resonates with the legislative reform. Demographic and job characteristics, such as gender, job tenure, and income levels, were critical in evaluating the association.

Robustness checks

Table 8 describes robustness checks using the duration and incidence of absence hours as the outcomes. We quantified the reform impact on five subsamples: (i) the sample of workers of registered organizations; (ii) the sample excluding workers of farm-household businesses; (iii) the sample excluding individuals in the military forces;

Table 5. Estimation on duration of sickness absence, by job tenure.

	Dependent variable: number of absence hours			
	Job tenure			
	<1 year	1–5 years	5–10 years	>10 years
All	0.027	0.102***	0.075**	0.050**
	(0.038)	(0.026)	(0.030)	(0.021)
	[36,636]	[175,169]	[130,386]	[165,623]
Monthly income < VND 5.5 mill.	−0.001	0.135***	0.179***	0.120***
	(0.039)	(0.032)	(0.043)	(0.035)
	[32,283]	[138,156]	[85,517]	[86,566]
Female, monthly income < VND 5.5 mill.	−0.043	0.199***	0.211***	0.154***
	(0.061)	(0.053)	(0.071)	(0.051)
	[15,266]	[66,414]	[41,556]	[39,169]

Note: VND 5.5 million is the announced mean income (AMI). All columns control for personal and job characteristics, log of income, time fixed effects, and regional fixed effects. The estimates use 1-to-1 nearest neighbor (with caliper 0.395, non-replacement) for matching. 1st number: Elasticity, 2nd number, in parentheses: Standard error, 3rd number, in brackets: number of observations. Standard errors are adjusted for clustering of personal identifiers; *Significant at p = 10%; **Significant at p = 5%; ***Significant at p = 1%.

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Table 6. Estimation on incidence of sickness absence, by gender and income.

	Panel A			Panel B		
Variables	All	Women	Men	All	Women	Men
DID	0.0019***	0.0021***	0.0011**	0.0019***	0.0020***	0.0011**
	(0.0004)	(0.0007)	(0.0005)	(0.0004)	(0.0007)	(0.0005)
	[507,814]	[222,830]	[284,984]	[507,814]	[222,830]	[284,984]
		<VND 5.5 mill.			≥VND 5.5 mill.	
	Both gender	Women	Men	Both gender	Women	Men
	0.0037***	0.0045***	0.0011*	−0.0006	−0.0006	−0.0010
	(0.0005)	(0.0009)	(0.0006)	(0.0007)	(0.0015)	(0.0007)
	[342,522]	[162,405]	[180,117]	[165,292]	[60,425]	[104,867]

Notes: This table is a combination of Tables 3 And 4, but dependent variable is incidence of sickness absence. VND 5.5 million is the announced mean income (AMI). All columns, except for those of Panel A, control for personal and job characteristics, log of income, time fixed effects, and regional fixed effects. The estimates use 1-to-1 nearest neighbor (with caliper 0.395, non-replacement) for matching. 1st number: Elasticity, 2nd number, in parentheses: Standard error, 3rd number, in brackets: number of observations. Standard errors are adjusted for clustering of personal identifiers; *Significant at $p = 10\%$; **Significant at $p = 5\%$; ***Significant at $p = 1\%$.

Table 7. Estimation on incidence of sickness absence, by job tenure.

	Dependent variable: number of absence hours			
	Job tenure			
	<1 year	1–5 years	5–10 years	>10 years
All	0.0016	0.0036***	0.0010	0.0020**
	(0.0015)	(0.0007)	(0.0009)	(0.0007)
	[36,636]	[175,169]	[130,386]	[165,623]
Monthly income < VND 5.5 mill.	0.0010	0.0044***	0.0033**	0.0039***
	(0.0016)	(0.0008)	(0.0013)	(0.0011)
	[32,283]	[138,156]	[85,517]	[86,566]
Female, monthly income < VND 5.5 mill.	−0.0004	0.0059***	0.0031	0.0057***
	(0.0024)	(0.0014)	(0.0020)	(0.0016)
	[15,266]	[66,414]	[41,556]	[39,169]
Male, monthly income < VND 5.5 mill.	0.0025	0.0018**	0.0001	0.0007
	(0.0023)	(0.0008)	(0.0013)	(0.0015)
	[17,017]	[71,742]	[43,961]	[47,397]

Note: VND 5.5 million is the announced mean income (AMI). All columns control for personal and job characteristics, log of income, time fixed effects, and regional fixed effects. The estimates use 1-to-1 nearest neighbor (with caliper 0.395, non-replacement) for matching. 1st number: Elasticity, 2nd number, in parentheses: Standard error, 3rd number, in brackets: number of observations. Standard errors are adjusted for clustering of personal identifiers; *Significant at $p = 10\%$; **Significant at $p = 5\%$; ***Significant at $p = 1\%$.

Table 8. Robustness check results.

	Registered organization	Nonfarm household business	Non-army force	Non-self-employment	Prime age
Panel A		Dependent variable: duration of sickness absence			
DID	0.0388**	0.0494***	0.0573***	0.0267*	0.0613***
	(0.0150)	(0.0142)	(0.0142)	(0.0152)	(0.0156)
Number of obs.	317,939	456,634	504,967	369,680	429,767
Panel B		Dependent variable: incidence of sickness absence			
	0.0011**	0.0014***	0.0019***	0.0009*	0.0018***
	(0.0005)	(0.0004)	(0.0004)	(0.0005)	(0.0005)
Number of obs.	317,939	456,634	504,967	369,680	429,767

Note: 1st number: Elasticity, 2nd number, in parentheses: Standard error. Standard errors are adjusted for clustering of personal identifiers; *Significant at $p = 10\%$; **Significant at $p = 5\%$; ***Significant at $p = 1\%$.

(iv) the sample excluding the self-employees; and (v) the sample of prime-age workers. Prime-age persons are defined as those in the age range 25 to 50 years old, who mostly completed their formal education, but are still too young to retire. (The retirement age for Vietnamese women is 55). By doing so, we tried to control for sample attrition effects. All estimations were controlled for, similar to those of Table 4. The coefficients were somewhat smaller than those of the full sample, except for column 5 of Panel A and column 3 of Panel B. However, all the

robustness checks confirmed the association between the legislative reform and illness absence. They also strengthened the parallel assumption.

The fact that in this study sick workers were remarkably responsive to economic incentives is consistent with the current labor economic theory. Understanding the effects of economic incentives on labor supply is essential as it helps policymakers to determine how to maintain national labor productivity, especially in a low-productivity country such as Vietnam.

Discussion and conclusion

Although a large volume of literature has indicated that sickness absence is affected by economic incentives, very few studies have assessed the impact of illness benefits in Vietnam. Our study adds to the debate on sickness insurance and enhances Vietnam's policy system's knowledge which has been ignored in previous literature. Our identification strategy is a combination of DID and matching approaches. Two main outcomes were expected: incidence and duration of sickness absence. Overall, the study pointed out that the generosity of sickness benefits under the 2014 insurance law positively affected the duration of sickness absence by around 0.058 hours (i.e., 3.5 minutes) per week on average. Its link to the incidence of sick leave was significant but small. Female workers seemed very sensitive while male workers showed indifference to the legislation, especially in terms of duration of sick leave. The augment in sick leave was primarily driven by workers employed for 1 year or more and having monthly income lower than VND 5.5 million. Various robust checks confirmed these findings with attrition samples.

Our research contributes to international literature that utilizes natural experiments to measure the impact of incentives on sick leave. The findings of this study are consistent with previous studies to some extent. Johansson and Palme (2002), Puhani and Sonderhof (2010), and Ziebarth and Karlsson (2014) have all shown that there is a positive link between sick pay and sick leave duration in their investigations of EU countries' data. Other research from Germany and Italy (e.g., D'Amuri, 2017; De Paola et al., 2014; Ziebarth & Karlsson, 2010) found that the legislative changes affect sick leave incidence. Nevertheless, the European sickness insurance schemes are not entirely similar to that of Vietnam as the Europe scheme covers nonwork-related sick leave only. The results of this study show how Vietnam's social insurance policies interact with the labor market. Its novelties include the Vietnamese context and the disaggregation of the effects by gender, income, and job tenure group.

Compared with the studies in Europe, which investigated reforms affecting some sectors of employment such as the private and public sectors (e.g., De Paola et al., 2014; Puhani & Sonderhof, 2010; Ziebarth & Karlsson, 2014), or all sectors but the central government (Pettersson-Lidbom & Thoursie, 2013), our work has focused on a different aspect, that is, on all sectors subjected to compulsory insurance.

A traditional interest reflected in the earlier literature is the gender difference in labor supply. In most countries, female employees take more sick leave than male employees. This fact happens even among unmarried employees. Explanations for this phenomenon include the difference in vulnerability, gender roles in both family and society, gender segregation within the labor market (Leao, Barbosa-Branco, Turchi, Steenstra, & Cole, 2017), nonlabor market responsibilities (Barmby et al., 2002), and even biological characteristics (Ichino & Moretti, 2009). Different policies are suggested for different genders (De Paola et al., 2014). Our research findings are in line with that literature.

Previous studies have also discussed the link between income level and sickness absence and its direction. Wages may have both the income effect (i.e., a higher wage means greater consumption level) and the substitution effect (i.e., a higher wage means greater income loss from absence). In general, as an insurance agency compensates only a fraction of the lost income, it is costly for high-income employees to take sick leave. Hence, a downward bias may occur when evaluating the cost of absence due to sickness benefits' generosity. On the contrary, employees with the desire to take leave may cause upward bias (Allebeck & Mastekaasa, 2004; Johansson & Palme, 2005). Our study contributes to the discussion by indicating that workers with a low income may be more sensitive to a small increase in sick pay (e.g., an income lower than VND 5.5 million).

Concerning job experience, a definite connection between the incidence of sickness absence and job tenure has previously been noted (e.g., Askildsen, Bratberg, & Nilsen, 2005; Barmby et al., 2002; Tompa, Scott-Marshall, & Fang, 2008). An interpretation of this fact is the "job security effect": Employees working for a longer time with an organization might assume that their jobs are secure, and therefore, the costs of absence are lower. Moreover, tenure is usually related to age, and the connection between age and sick leave is another interpretation. Our research adds to the literature by providing evidence of labor effects on job tenure of 1 year or longer.

On the contrary, people may argue that our estimated results are a consequence of a trend in sickness absence instead of the legislative change. This argument,

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however, does not hold up to scrutiny. First, it should be noted that there is a distinction between the reform effects on the likelihood and duration of sick leave. These effects are not in the same direction. Hence, it is unlikely that two distinctive trends would have a close link. Second, the application of the DID technique provides another explanation.

The Vietnamese government is committed to making further national social insurance reforms by 2030 to enhance social protection. A Master Plan on Social Insurance has been developed, intending to raise coverage and enhance insurance benefits for all citizens. Taking sick leave may give workers the time needed to rest and recuperate, and thus, improve their health status. Also, paid sick leave helps them to receive proper treatment. Generous sickness benefits can give people more opportunities to care for their family and contribute significantly to family health, especially that of children and older-aged parents. Increasing sickness benefits may also reduce the likelihood of presenteeism, namely, showing up for work while sick. This helps to reduce a negative influence on productivity, health, and the spread of workplace diseases, which causes more absenteeism (Heymann et al., 2010).

The implications of this study are straightforward, as knowledge of the impact of the reform contributes to improving the effectiveness of future policies. Our work points out the importance of sickness benefits as well as demographic and job characteristics (e.g., gender, income, and job tenure) in evaluating workers' behavior. Further health care and insurance reforms may need to consider these characteristics. To ensure better long-term health through sickness benefits while maintaining national labor productivity, our research provides important pointers for policymakers in Vietnam and other developing countries.

The findings of this study, however, should be interpreted with care. Given the small change in illness benefits, we have to be cautious in describing its specific effects. Moreover, as VLFS does not include detailed information on workers' health, we cannot quantify the impact of the legislative change on health status. Despite such limitations, we believe that our empirical strategy using a quasi-experiment has revealed important findings for future studies in labor economics and health policy.

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Appendix

Table A1. Summary statistics.

Variables	Mean	Std. Dev.	Min	Max	Obs
Outcome variables					
Incidence of sickness absence (D)	0.0051	0.0712	0	1	568,076
Hours of sickness absence	0.1292	2.1779	0	90	568,076
Personal characteristics					
Age	38.969	10.444	15	60	568,076
New migration (D)	0.0145	0.1198	0	1	568,076
Female (D)	0.4865	0.4972	0	1	568,076
Child aged 0–5	0.2614	0.5239	0	5	568,076
Education	4.9150	1.9011	1	8	568,076
Job characteristics					
Job tenure	3.8409	1.0581	1	5	568,076
Fulltime (D)	0.8823	0.3221	0	1	568,076
Cooperative (D)	0.0026	0.0515	0	1	568,076
Regional unemployment rate	0.0207	0.0089	0.0032	0.0496	568,076
Other useful variables					
Compulsory social insurance (D)	0.4109	0.4920	0	1	568,076
Registered organization (D)	0.6641	0.4723	0	1	567,521
Actual hours worked (last week)	45.543	11.815	0	99	567,973
Usual hours worked (a week)	47.325	9.6131	2	99	568,074
Monthly income	5370.5	5254.1	33.876	871586	568,076
Farm household (D)	0.0901	0.2863	0	1	568,076
Self-employment (D)	0.2431	0.4290	0	1	568,076
Army force (D)	0.0070	0.0837	0	1	568,023

Note: (D): dummy variable.