Play Hard, Shirk Hard? The Effect of Bar Hours Regulation on Absence

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Abstract

Government influence on labour supply behaviour through taxation and transfer policies is well understood. However, they also act to regulate the timing and extent of leisure activity. This has the potential to influence labour-leisure decisions. Legislative changes in bar opening hours provide a potential quasi-natural experiment of the effect of government regulation on working effort. This paper examines two recent policy changes, one in England/Wales and one in Spain that increased and decreased opening hours, respectively. A robust positive causal link between opening hours and absence is demonstrated. Further evidence is provided that longer opening hours cause poorer health outcomes, particularly amongst regular bar attendees.

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I. INTRODUCTION

How taxation and transfer payments influence individual labour supply behaviour is well understood. There is, for instance, a large body of research on the incentive effects of taxation and how hours of work are affected by taxes and transfers (see for instance Burtless and Hausman, 1978; Blundell, Meghir, Symons and Walker, 1988; Blomquist et al., 1990; Bourguignon et al., 1990; Heckman, 1993; Blundell, Duncan and Meghir, 1998). In turn, how governments influence leisure decisions is typically thought of through the lens of income and substitution effects. However, governments commonly intervene and regulate leisure activity directly. The reasons for and the forms of regulation are numerous. These include prohibition and restriction of the use of recreational substances, but also restrictions on the timing of the consumption of a range of leisure activities. These timing based interventions are typically justified on the basis of reducing negative externalities from leisure behaviour. For instance, restrictions in opening hours for live music venues (noise pollution), restrictions in the timing of night time sports in urban areas (light pollution) and restrictions on the opening hours of licensed venues. As a result, a body of research has examined the effect of bar opening hours on a range of health and socio-economic outcomes including alcohol consumption (Bernheim, Meer and Novarro 2012), traffic accidents (Vingilis *et al*., 2005; Smith 1990) and crime (Carpenter and Dobkin, 2011b; Chikritzhs and Stockwell 2002; Biderman *et al*., 2010; Hough and Hunter, 2008 and Humphreys and Eisner 2010). In addition to these outcomes, these interventions have the potential to markedly influence workers' leisure-labour decisions, but this has received little attention to date. In general, individual labour supply behaviour could be influenced by leisure regulation if, for instance, it affects the timing proximity of leisure consumption and working hours. In the case where it involves intoxicating substances, like alcohol, the timing of consumption could also have spill-over effects into working hours.

This paper investigates this issue by examining how the regulation of licensed hours at establishments that serve alcohol influences working hours, focusing primarily on worker absenteeism. While, there is no existing evidence along these lines Carpenter and Dobkin (2011a) have previously suggested that alcohol legislation in the form of minimum drinking ages can influence workforce productivity. We use recent changes in legal pub and club (herein bars for simplicity) opening hours in the UK and Spain to identify the effect on absence, which provides a readily measurable proxy for work effort (Audas et al, 2004).

These two legislative changes provide a nice point of comparison, as one involves a substantial liberalisation of opening hours (the UK) while the other involves a similarly substantial decrease in opening hours (Spain). These changes have the potential to affect working behaviour due to the proximity of leisure activity to normal working hours, but also through the timing of the consumption of alcohol. It is difficult to definitively disentangle these two channels of influence. However, we provide evidence on transmission channels by further examining the causal effect of these legislative changes on individual health outcomes and expenditure on alcohol in licensed premises which may be indicative of the role of variations in alcohol consumption.

To summarise our results, we demonstrate a causal link between bar opening hours and worker absenteeism, longer opening hours increase absence. We do this by taking advantage of a `quasi-natural' experiment that entails a liberalisation of drinking licensing hours in the UK and a contraction in Spain. These results are symmetric for Spain and the UK; decreasing opening hours (Spain) reduces absenteeism, increasing opening hours (UK) increases absenteeism. These results are robust across a range of specification and differing identification strategies within both countries. For instance, whilst we can identify the causal effect using difference in difference approaches, we also identify the policy effect within a panel fixed effects strategy for the UK, and demonstrate the robustness of our results to other common sources of bias in the estimates derived from applying a difference-in-difference methodology. In particular, both the fixed effects approach and the multiple country nature of our study reduces the concern that our policy effect is being driven by common unobserved random shocks. We demonstrate that the policy effect is concentrated among young workers and in the UK amongst women in particular. This policy effect may reflect the impact of the removal of constraints on the proximity of leisure timing to work timing and/or the effect of alcohol consumption on labour supply. In further estimates we provide evidence of a causal effect of drinking laws on individual health outcomes, and weak evidence of an expenditure increase on alcohol at bars. This is suggestive that the main channel of the absence effect we have identified is through alcohol consumption.

II. DATA AND INSTITUTIONAL BACKGROUND

Changes in Drinking Laws, Spain and the UK

The identification strategy in this paper is based on two legislative changes; a reduction in the permitted hours that bars could remain open in Spain and an extension of legal closing hours in two parts of the UK, England and Wales. In the Spanish case, this reduction in opening hours consisted of a requirement that licensed venues, such as bars, were legally required to close at 3:00 am (with some minor variation noted below). Prior to the legislative change the legal closing time was 6 am, and the majority of drinking venues did not close until this time. This legislation was enacted at different times regionally across Spain, and varied in terms of the actual new time of closing ranging from 2:00 am to 3:30 am.[1] Specifics of the actual legislative changes are reported in Table 1. Column 2 of Table 1 shows the quarter and year the reform came into force in Spain in each of the regions (reported in column 1).

For England and Wales, prior to the legislative change pubs were not allowed to stay open (and serve alcohol) after 11:00 pm. Following the Licensing Act of 2003, licensed venues could apply to remain open for longer up to a maximum of 5:00 am. This came into effect in all of England and Wales as of the 24th of November 2005, as at 1st April 2006 (the first available official statistics) some 50114 venues had been granted these licenses. By 2010 this had increased to 78879 venues. Hence the main expansion occurred in the initial time period that the legislation was enacted. It is worth noting that the stated reasons for these two legislative changes were markedly different. In Spain, it primarily reflected concerns over noise pollution and general disruption to residents near licensed venues. While in the UK, it reflected a view that the prior regime of 11pm closing was needlessly restrictive and that shorter opening hours may encourage binge drinking insofar as individuals would increase the speed of alcohol consumption.

INSERT TABLE 1

[1] The differential timing of the reform in Spain reflects the devolution of certain legislative powers to regional levels. In the case of public entertainment and recreation policy, devolution was completed by 1996. This meant that whilst the key legislative change in opening hours was mandated at a federal level, some regional autonomy in the timing of the adoption and actual closing times was permitted. We investigate the potential for bias of our results from non-random timing of adoption later.

When comparing the effects of these types of legislative changes such as England, Wales and Spain one must be aware of the substantial cross-country differences in culture and habits related to alcohol consumption and the attendance of licensed venues. While both have the same legal age of drinking, 18 years, the difference in the culture related to drinking in the UK and southern European countries such as Spain are large and well-known. The UK has long recognised problems with excessive and binge alcohol consumption. For instance MacDonald and Shields (2004) report problem drinking rates for males in the UK of around 20%, and that 10% of the male population aged 22-64 drank at least 45 units of alcohol per week.[2] Alcohol consumption in Spain is common, Gual (2006) reports that approximately 60% of male and 35% of females drink alcohol weekly. However, excessive and binge drinking has traditionally been uncommon. In comparison to the figures for the UK above, less than 20% of males and 10% of females in Spain report drinking more than 5 units of alcohol at least once per week. An additional key difference between the two countries is the demographics of bar attendance. In the UK, pub attendance is common across age groups. For instance data from the British Household Panel Survey reveal that in 2000 62% of males and 55% of females aged 16-24 years report `going out for a drink' at least once a week, this drops to 44% and 27%, for males and females respectively aged 25 to 34 years, but stays remarkably high after that; 38% of males aged 35-64 report going out for a drink at least once a week, while the figure for females is 19%. In contrast, it is generally understood that bar attendance is heavily concentrated among young people in Spain (Calafat et al., 2002). As a result, whilst there are some statistics available on young people's bar attendance in Spain,[3] there is no comparable statistics for the over 30's. These differences in the demographics of bar attendance help to inform our country specific identification strategies later.

Data

This paper uses two data sets that are very similar in basic structure, the UK Labour Force Survey (UK LFS) and the Spanish Labour Force Survey (SLFS). Both are quarterly representative surveys that provide a range of information on individual and work

[2] Where a unit of alcohol is defined as 10 millilitres in the UK, which is 0.564 of a US standard unit of alcohol (17.7 millilitres). While a Spanish unit of alcohol is 12.7 millilitres.

[3] For instance the Youth in Spain Report (2008) shows that 48% of 15-19 years olds report that they go out to bars either all or almost all weekends, the percentage is 47% for 20-24 years old and by the age of 25 to 29 this has decreased to 31%.

characteristics. A key feature of the data for our purposes is that they both have an internationally consistent definition of absence (Barmby, Ercolani and Treble, 2002), which we describe in more detail below.

The SLFS is a quarterly survey from which we have data available from the 1st quarter of 1996 to the 4th quarter of 2007. It is a repeated cross-section and contains 3,090,703 observations (our estimated sample consists of 1,719,510). For the UK, a 5 quarter rotating cohort version of the LFS is available which we use. This follows individuals for 5 consecutive quarters from entry. It is a rotating panel insofar as every quarter one cohort enters and another exits (after their 5 quarters). For the UK we narrow the data period to be more closely centred around the policy change. As a result we use 2003 to 2008 and this provides 472,017 observations for 128,444 different individuals.

We use information on usual and actual hours of work per week to generate two indicators of absence. The first is the hours a worker is absent per week. We calculate this variable as the difference between usual hours and actual hours $A_{it} = H_{it}^{u} - H_{it}^{e}$.[4] For ease of interpretation we multiply this number by 60 so that the estimated coefficients are in terms of minutes of absence. The second variable is the absence rate. It is defined as the ratio of the hours reported absent to contracted hours in the reference week $AR_{it} = A_{it} / H_{it}^{u}$. These measures of absenteeism may include variations in time at work that are outside of the control of the worker and as a result should not be readily affected by changes in drinking laws. Both the SLFS and the UK LFS contain information on why hours varied in the reference week. This allows us to construct absence measures that are more narrowly defined, excluding (for instance) variation due to flexible working hours, variations due to changes in jobs, training episodes and industrial disputes. Importantly, our key estimates are robust to using these narrower definitions of absence. This is discussed in more detail in the results section. Finally, this measure of absence may also capture any variation in contractual hours caused by the policy. In unreported estimates we found no effect of the policy change on contractual hours in either England/Wales or Spain. We also found no effect of the policy changes on the probability of being employed.

[4] We consider usual hours as synonymous of contractual hours. This is similar in spirit to the approach used in previous research by Barmby et al (2002), Lozano (2010) and Green and Navarro (2011).

Both data sets have quite a rich set of potential control variables, including many of the candidates that have been shown to be important determinant's of worker absenteeism in previous research. Thus, we incorporate socio-demographic variables, including the age and the age squared, gender, marital status, education level. We also include labour market variables which denote whether the individual works in the public sector, the type of contract, industry dummies, occupation dummies and size of the firm/establishment. We also control for year and quarter to take account of seasonal and time variations.

An important issue is that certain individuals' working hours may be directly affected by the change in drinking laws, most notably those who work in bars. We exclude all individuals working in these establishments, and to be especially sure, those working in allied industries such as hotels and restaurants. Finally, workers on part time work may have more natural variability in their working hours; in the results we investigate the robustness of our results to excluding part time workers. Appendix Table A1 provides summary statistics for the resultant samples for both Spain and the UK.

III. METHODOLOGY AND IDENTIFICATION

The differences in the nature of the legislation, data and institutional factors lead to variations in the identification strategy we adopt for Spain and England/Wales.

For Spain we adopt a difference in difference approach to estimate the effect of reducing opening hours on absence.[5] To assign individuals to treatment we rely upon the age-concentration of individuals drinking and who attend licensed premises, especially beyond 3 am in Spain. Specifically, our treatment group is young people. Our comparison group is older individuals (> 45 years old). Workers' minutes of absence per week can be specified as follows:

$$A_{ijt} = \phi + \delta Policy_{ijt} + \gamma\beta Treatment_i + \beta Policy_{ijt} \times Treatment_i + \alpha X_{ijt} + \tau Y_{it} + \sigma R_{ij} + \varepsilon_{ijt} \qquad (1)$$

Where A_{ijt} corresponds to the minutes of absence of worker *i*, in region *j* in period *t*. *R* and *Y* are sets of regional and year dummies, respectively. $Policy_{ijt}$ takes the value 1 if the worker is

[5] In principle the policy effect could be identified using just regional-time variation in policy implementation (i.e. without the comparison group). In unreported estimates, these provide estimates of 5 minutes as a result of the reduction in licensed drinking hours.

observed in region *j* and at time *t* that region has reduced drinking hours. $Treatment_{ij}$ is a dummy variable that equals one if the worker is 30 years or younger, and 0 if the worker is older than 45 years (we exclude all workers aged 31-45). The interaction term $Policy_{ijt} \times Treatment_{ij}$ equals one for treated individuals in the post-treatment period (after the legislation came into force in the region of the individual's residence). Estimates of β is equivalent to the Differences-in-Differences (DiD) estimator and this provides the absence caused by the reform for the treated group (i.e. the reduction in absenteeism for young workers caused by shutting bars and pubs earlier). A nice feature of our institutional setting is that we observe 11 regions changing drinking hours at different times. This makes it less likely that our policy effect is being identified by some unobserved shock occurring at the time of policy implementation.

Our identification strategy for the UK differs. In England and Wales there was no differential timing of the reform and there is a substantially less pronounced variation across age in drinking habits and attendance of licensed venues which makes assigning treatment based on age not viable. Instead, we have a clear comparison group, workers in Scotland and Northern Ireland where there was no change in drinking laws at this time. This leads to the following difference in difference strategy:

$$A_{ijt} = \phi + \delta Policy_{it} + \gamma\beta Treatment_{ij} + \beta Policy_{it} \times Treatment_{ij} + \alpha X_{ijt} + \tau Y + \varepsilon_{jit} \qquad (2)$$

where A_{it} corresponds to the minutes of absence of worker *i* in period *t*. $Policy_{it}$ equals one if the worker is observed after 24th of November 2005, 0 otherwise. $Treatment_{i}$ is a dummy variable that equals one if the worker resides in England or Wales and 0 if he/she is in Scotland or Northern Ireland. The interaction term $Policy_{it} \times Treatment_{i}$ equals one for treated individuals (those living in England or Wales) in the post-treatment period. Again, the OLS estimate of β is equivalent to the Differences-in-Differences (DiD) estimator and this provides an estimate of the increase in absence caused by the licensing laws for workers in England or Wales compared to those living in Scotland or Northern Ireland.

The single timing of the policy implementation in the UK makes the underlying assumptions needed to interpret β as the true policy effect on absence stronger. Along with a range of robustness tests reported later, we also exploit the longitudinal nature of the data for the UK to use an additional complementary identification strategy. Given that we observe an individual’s absence behaviour before and after the policy change we can estimate a model of absenteeism including worker level fixed effects:

$$A_{it} = \phi + \sigma_i + \eta Pubs_t + \alpha X_{it} + \varepsilon_{it} \qquad (3)$$

The estimate of η follows from the within worker variation in absence behaviour across the period of reform for the treated individuals only. Hence this model is identified only for those workers who resided in England and Wales and that we observe for at least one quarter before and one quarter after the reform came into effect. As a result we estimate (3) only on workers in England and Wales. This provides an indication of how given workers changed their behaviour after the increase in drinking hours. Because we can observe the week of interview in the UK-LFS we can identify this policy effect separately from quarter controls aimed to pick up seasonality in absence.

A standard concern in the literature on policy evaluation using difference in difference approaches is that spurious inference may result if the error structure is not modelled correctly. Specifically, a concern in our case would be the assumption that the error term is normally distributed within the regions in which our workers are embedded. This may lead to standard errors which are artificially low. To address this, in all of our models we cluster standard errors at the regional level (18 regions for Spain and 20 regions for the UK). In additional robustness checks reported later we also adopt the approach suggested by Bertrand et al (2004) to assess any downward bias in standard errors due to serial correlation in the dependent variable. Finally, in all models we estimate variants of (1), (2) and (3) where the dependent variable is instead the absence rate (AR) as computed above.[6] This dependant variable is more flexible insofar as it explicitly allows for variations in contractual hours.

Figures 1 and 2 provide some illustrative information on the changes in the dependent variables with respect to the policy change. Specifically they show absence behaviour in the immediate periods around the policy changes, recalling that these absence figures capture any variation in hours worked from contractual hours. These figures provide a preliminary indication of three key things. First there are variations in absenteeism behaviour across the policy regimes for the treatment groups. Second, there is almost no change, and no

[6] We also estimated, but do not report, the policy effect on the incidence of absence; where this took the value of 1 if usual hours exceed actual hours in the reference week and 0 otherwise. The pattern of sign and significance of the key policy estimates for this alternative measure of absenteeism were identical to those reported for minutes difference and the absence rate.

statistically significant change, in absence behaviour for the control group in Spain. Finally, there is some minor reduction in minutes of absence for the control group in the UK case which could potentially bias upwards our policy estimates. This change in Scotland and Northern Ireland is not, however, statistically significant at standard levels. In addition, it will not influence the within worker fixed effects estimates for England and Wales.

IV. RESULTS

Table 2 provides the estimates for the effect of the drinking law regulation in Spain on worker absenteeism. Two groups of estimates are reported, Tobit estimates for absence rate and OLS estimates for hours difference.[7] Initially, for means of comparison we also report estimates for a narrower treatment group, workers aged 25 or less. The control group in all cases are workers aged more than 45. It is worth noting however that our estimates are not substantively altered by using less restrictive comparison groups such as greater than 30 or 35 year old workers. Moreover, the influence of this choice affects our policy estimates in an expected manner. Less restrictive comparison groups lead to some decrease in the magnitude of our estimates of policy effects. A set of standard control variables are included covering gender, marital status, education, sector of employment, contract type along with occupation, industry, regional, year and quarter dummies which are not reported for the sake of brevity but are available on request from the authors. Age is controlled for both with a quadratic functional form, but our key estimates are unaffected by more flexible parameterisation such as including age dummies. The standard errors clustered at a regional level (18 regions) are presented in parentheses below the parameter estimates.

Looking at the coefficients on the key variable of interest (Policy $\times$ Treatment) demonstrates a substantial effect of the drinking law regulation on worker absence. For instance, the effects range from a decrease in the absenteeism rate of between 3.5% and 4.6% and a corresponding reduction in working minutes lost through absence of between 14 and 17 minutes. This is a marked effect when compared to our sample means for the treatment group ($\leq$ 30 years) of 8.8% absence rate and 200 minutes of absence.[8] Moreover, this effect

[7] In unreported results, we also estimated the absence rate models by OLS, the sign and significance of our estimates were unaffected by this.

[8] Overall sample means are an absence rate of 10.5% and 239 minutes of absence.

increases in magnitude as we look at our younger treatment groups. These estimates suggest that reducing the legal opening hours of licensed bars in Spain substantially reduced worker absenteeism among younger workers.

INSERT TABLE 2 HERE

Table 3 displays the corresponding Tobit and OLS estimates for the increase in licensed bar opening hours in England and Wales. The chief difference here is that we do not focus on worker age to assign treatment status but instead exploit the lack of legislative change in Scotland and Northern Ireland. Again we report the effect of the legislative change on worker's absence rate and minutes lost due to absence. The control vector is similar to that for the case of Spain, with only a slight difference in the education controls reflecting cross-country differences in qualification structure. Again for brevity we do not report the estimates for the occupational, industry, year and quarter controls.[9] As in the Spanish case the impact of the legislative change on absenteeism is substantial for workers in England and Wales. In this case, increasing opening hours increased worker absence by approximately 3% and lead to an increase in time lost through absence of 14 minutes per week. The standard errors clustered at a regional level (20 regions) are presented in parentheses below the parameter estimates. This, when combined with the results for Spain suggest a positive causal relationship between licensed opening hours and worker absenteeism.

INSERT TABLE 3 HERE

In the case of England and Wales we can go a step beyond difference-in-difference estimation and exploit the panel dimension of the UK LFS to examine how given workers absence behaviour changed post-reform. We use a sample of workers from England and Wales only for the period 2004 to 2008. Table 4 reports estimates from panel fixed effects models of absence rates and minutes of absence. These models are identified for workers who we observe in our five quarter panel before and after the legislative change in England and Wales. Again these results show that the extension of drinking hours substantially affect worker absence behaviour. For instance, the policy effect on minutes of absence is slightly larger than that reported earlier in the difference-in-difference estimation 21 versus 14. The

[9] It is worth noting that in both the Spanish and UK cases the estimates on the control variables largely follow those previously reported in the literature on absence. For instance, temporary workers take less absence (Bradley, Green and Leeves, 2011, Ichino and Riphahn, 2005), public sector workers take more absence and female and married workers take more absence (Barmby, Orme and Treble, 1991).

effect of the absenteeism rate is less marked, it reduces from approximately 3% to 1.5%, but still remain statistically significant at standard levels. This suggests that these earlier results were not driven entirely by, for instance, some compositional change in the unobservable characteristics of workers pre-and-post reform, or due to some change in behaviour of our control group of Scottish and Northern Irish workers that was contemporaneous to the legislative change. In unreported estimates we also investigated whether there was a within-worker change in absenteeism in Scotland and Northern Ireland at the time of the policy introduction as a form of placebo policy test. The resultant fixed effects estimate of the placebo policy effect whilst positive was far from statistical significant at standard levels (9.59 [S.E. 15.64] and 0.007 [S.E. 0.007] for hours difference and absence rate, respectively).

INSERT TABLE 4 HERE

Before moving on to robustness tests of our key estimates, in Table 5 we provide estimates for two subsamples that are of particular interest. First, our current estimates include all workers irrespective of working hours. A concern is that part-time workers have hours of work that naturally vary and this may somehow bias our policy estimates. The first columns in Table 5 provide estimates for full-time workers only. In all cases the difference between these estimates and those reported earlier are at most modest. Another issue is that the policy may impact differently across gender, hence we report males and females separately in subsequent columns. For Spain, there is some suggestion that the policy impact was larger for male working hours. However, for the UK a more dramatic pattern appears. The policy impact seems concentrated almost entirely in female workers.[10] In a recent study of the effects of reducing alcohol tax in Finland, Johansson et al (2011) find that female absence increased much more dramatically (13%) than male absenteeism (5%) in Swedish regions near the Finnish border.

INSERT TABLE 5 HERE

FURTHER ROBUSTNESS CHECKS

[10] We have no definitive explanation for why there is this marked gender difference but it is worth noting that female drinking has increased markedly in the UK within the last decades. For instance it has been reported that female binge drinking rates have doubled in the UK since the early 1990’s (Smith and Foxcroft, 2009).

Our results demonstrate that drinking law regulations have the potential to influence an individual's intra-marginal labour supply decisions. A strength of our approach is that the effect is found for two different countries where the policy was operating in different directions, and moreover, where different identification strategies are used. Nonetheless in this section we conduct a range of other robustness checks on our main estimates.

First, we examine whether our results reflect disruption associated with the implementation of the policy. We do this by excluding the year of reform from our sample. In the case of Spain the excluded year varies by region due to the differential timing of reform. These estimates are reported in Table 6 for both countries (for Spain and the UK). These reveal that omitting the year of the reform does not change the main results. For instance, the estimated effect of the legislative change on the absence rate in Spain and the UK is not statistically different from those reported earlier. The same is true for the minutes difference models. We can use this type of approach to assess whether the policy continued to influence behaviour in the years after implementation. These estimates reveal two things, the policy is not being identified by some form of disruption or other implementation effect, and by extension the policy continues to exert an influence on worker absence behaviour at until, at least, one year after the reform year. This second point is important as it suggests that the policy has a lasting impact and individuals do not revert to pre-policy behaviour after some period of adaptation. Finally, we re-estimated our main models with the reform lagged one year, this provides another form of placebo test. Estimates from these models revealed no effect of the lagged reform on absence behaviour.

As mentioned earlier, our measures of absenteeism may be too broad insofar as they capture all variations in working time, including those that occur for reasons out of the control of workers. In unreported estimates we used information in the SLFS and UK LFS on reasons for variation of working hours to exclude categories that were least likely to be in the control of workers and hence, be affected by the policy. Specifically we excluded those workers who's hours `usually vary' along with absence due to changing or loss of job, undertaking training, and union representation, strike or labour conflict and technical partial stop or employment regulation within a firm because of financial problems. This does change the tenor of our estimated policy effects. For Spain, the policy estimates are essentially unchanged. For the UK there is some increase in the estimated effect of the policy change. For instance, UK minutes of absence due to the reform increases to 18.59 (from 14.08) and the absence rate is 10% (from 3%) in the difference-in-difference models. The corresponding figures for the fixed effects models are 63.05 minutes of absence and an absence rate increase of 3.6%.

Two further issues relate to policy implementation, in the case of Spain there was some discretion in the timing of the adoption of the policy, as reported in Table 1. It could be that regions where there were more marked problems related to extended drinking hours adopted the policy early and this may bias our results. To investigate this we re-estimated our DiD models for those three regions that adopted early, La Rioja, Balearic Islands and Pais Vasco. Whilst this lead to some loss in precision, the policy estimates for these early adopters were essentially the same as our main results, For instance the minutes difference and absence rate effects were -16.89 and - 2.5% and, respectively.

A related issue with the estimates for England and Wales, is that unlike Spain, the change in licensing were in effect not mandatory. That is, individual venues had to apply for an additional licence to remain open later. We use this to further investigate if actual variations in drinking hours are causing the change in absence behaviour. The UK Department for Culture, Media and Sport reports the number of licenses granted by region. In areas where there is a greater density of venues that increased hours, we might expect a larger absence response. Most regions have quite a similar density of extended hours licences per head of population (16 years or older) of between 0.94 licenses per thousand people and 1.47 per thousand people. However, three regions have particularly high densities, the South West of England, London and the North East of England (1.47, 1.44 and 1.25, respectively). We re-estimated our DiD models for these regions only (again using Scotland and Northern Ireland as control groups) and these reveal slightly higher estimates of the policy effect than those for England and Wales in total, for instance the estimate of 14 minutes rises to 21 minutes. These estimates remain statistically significant at standard levels.

A standard concern with DiD estimates in repeated cross-sections with a long time dimension is that if there is serial correlation in the dependent variable this leads to standard errors that are biased downwards and hence incorrect inference (Bertrand, Duflo and Mullainathan, 2004). To investigate this we collapse our data (by group characteristics in Spain and by individual in the UK) into two periods, pre and post reform. We then re-estimate equations (1) and (2) on this collapsed data. For the case of Spain, the regressions are weighted by the number of observations in each cell to replicate the underlying micro data. Again, the standard errors are clustered at a regional level. The results are reported in Table 7 and demonstrate that the policy caused young worker's absence rate to decrease by 1% in Spain and an increase by 1.9% in worker's absence rate in the UK. Young workers reduce the minutes they are absent from work in Spain by 21.77 while workers in England and Wales increase the minutes they are absent from work due to increase in pub closing hours by 17 minutes. Importantly, these estimates remain statistically significant at standard levels and do not suggest that our previous policy inference was incorrect due to serial correlation in absenteeism.

Why Do Drinking Laws affect Workplace Absence? The Role of Drinking, Health and Consumption.

To this point we have demonstrated a robust causal effect of changes in bar opening hours on worker absenteeism. However, we cannot directly distinguish whether the effect comes from a pure leisure-labour trade off due to, for instance, the timing and the choice of sleeping hours (Biddle and Hamermesh, 1990) or in an indirect way through a spill-over of alcohol consumption and intoxication into working hours. Here we use further household data for Spain and the UK to examine whether the policy changes affected individual health outcomes. This, we argue, may be indicative of a channel of effect via changes in the level of alcohol consumption. A literature exists that examines how relaxing the sale of alcohol at off-premise locations influenced alcohol consumption in Sweden (Olsson and Wikstrom, 1982; Norstrom and Skog, 2003, 2005) and the effect of so called Blue Laws or Sunday sales bans in the US (Stehr, 2007) and Canada (Carpenter and Eisenberg, 2007, 2009). At the same time, it has been demonstrated that there is a link between increased alcohol consumption and absenteeism (Norstrom and Moan, 2009; Johannson et al, 2008; Balsa and French, 2010; Johannson et al, 2011). It is important to note, however, that any effect of licensing hours on health outcomes need not only reflect a direct role of alcohol consumption. For instance previous research has demonstrated that liberalisation of off-premise alcohol sales increased traffic accidents (Lovenheim and Steefel, 2011; McMillan and Lapham, 2006; Stehr, 2010; Heaton, 2012) and crime (Gronqvist and Niknami, 2011). These could in turn influence individuals' health outcomes.

We use two data sets, again with similar structures. For Spain we use the European Community Household Panel Survey (ECHP) 1994-2001, while for the United Kingdom we use the 1997-2007 data from the British Household Panel Survey (BHPS). Both ask individuals a variant of the following question, do you have any physical or mental health problem, illness or disability.[11] These include health related problems unlikely to be directly affected by alcohol consumption. This introduces measurement error with the resultant bias in our policy estimates towards zero. We use these responses to construct a binary dependent variable of health problems which we include in analogous regression specifications to (1) and (2).

INSERT TABLE 8

Again our empirical strategy varies between Spain and the UK. For Spain, because of the time period of the ECHP the policy change is only observed occurring in the Balearic Islands. We use an analogue of the DiD model for absenteeism, where 30 years olds or younger are our treatment group and those older than 45 years of age are the comparison group. A further difficulty is that in the regional disaggregation available (NUTS 2) in the ECHP the Balearic Islands are grouped with the regions of Catalonia and Valencia where there was not a policy implemented at that point. As a result our estimates provide a lower bound for the Balearic Islands, and one that may not be generalisable to the rest of Spain. These estimates reveal a negative but statistically insignificant reduction in health problems due to the policy of -0.01 [s.e. 0.008]. If we choose a younger treatment group, 25 years or younger a larger more negative policy effect results, -0.021[s.e. 0.012] which is statistically significant at the 10% level. This provides some weak evidence that the reduction in drinking hours improved youth health outcomes.

INSERT TABLE 8

For the UK the data is more advantageous in a number of ways. The BHPS allows us to replicate directly the difference in difference specification from before but also has additional information that provides more confidence that the estimated effect is actually being driven causally by the policy change. Given the gender disparities in policy effect revealed earlier for England and Wales we report all health estimates separately for males and females. The initial difference in difference estimates reveal that licensing laws increased the incidence of health problems of males and females in England and Wales by 1.8 percentage

[11] The questions are, in the ECHP, do you have any chronic physical or mental health problem, illness or disability? while in the BHPS it is do you have any health problems or disabilities?

points and 2.1 percentage points, respectively.[12] To help pinpoint the source of this apparent policy effect we use a question in the BHPS which asks how often, on average, the interviewee goes out to licensed venues to drink. We use this information to estimate separate models according to whether the individual reports going out at least once a week, at least once a month or at a frequency less than once a month. If it is changes in licensing hours driving these health effects then it should be more pronounced in more frequent attendees of bars, and zero for those who do not frequently go to bars. The estimates reported in the last 3 columns of Table 9 fit with this intuition. Regular drinkers, who are most likely to be affected by the policy, had a substantial increase in the incidence of health problems due to the change in licensing hours. This effect is absent for infrequent attendees of bars.

As noted before, these effects on health outcomes could arise from a range of potential channels, including increased alcohol consumption. To provide some final evidence along these lines we sought to examine whether these health and absence effects were matched by a change in expenditure on alcohol at licensed venues. To do this we used the 2001 to 2008 waves of the UK Expenditure and Food Survey (EFS), which provides a representative sample of household's expenditure in the UK as an annual repeated cross-section.[13] The EFS asks respondents to keep a two week diary detailing expenditure items and the value of purchases. In particular, it provides information on expenditure on alcohol at licensed venues. A difficulty with this data is the excess of zeros which could reflect either that these individuals never consume alcohol at bars, or merely that their consumption was zero in the reference weeks. If we estimate a simple analogue of our DiD model for the UK with log alcohol expenditure (£) at licensed venues as the dependent variable and again England/Wales as the treatment group we find no effect of the policy on consumption. Limiting our sample to non-zeros we find that individuals in the treatment group in the policy period increased expenditure by approximately 5 percent. This again hides gender differences, whereby female drinkers increased expenditure by nearer 10 percent, and male

[12] Of course these results could reflect unobservable factors influencing both drinking and health. Whilst there are well known problems with implementing conditional logit models with respect to sample selection, we re-estimated our model of the policy effects on health to account for unobserved time invariant characteristics. This revealed a marginal effects coefficient of 0.078 [S.E. 0.007] of the policy effect on having a health problem.

[13] A Spanish FES equivalent exists. However a lack of consistent data on the particular expenditure group of interest across our policy period means that we cannot estimate the policy impact on alcohol expenditure at drinking establishments in Spain.

expenditure did not increase. This provides some weak evidence of a policy effect on consumption, at least amongst the sub-group of the population who choose to drink at bars. Again, this is suggestive that the effect of changing licensing hours on absence is related to alcohol consumption.

V. CONCLUSION

This paper sought to examine how changes in the regulation of leisure activities can influence individual labour supply decisions. Specifically, we used two recent and symmetric changes in the legal opening hours of licensed premises in Spain and England and Wales. These are particularly advantageous insofar as they provide policy changes in opposite directions, a reduction in drinking hours in Spain and an extension in England and Wales. Focusing on one dimension of intra-marginal labour supply, absenteeism, we demonstrate a causal effect of these legislative changes. Reducing opening hours in Spain reduced absenteeism, whilst increasing opening hours in England and Wales increased worker absenteeism. This result proves robust to a variety of specifications, alternative treatment and control groups and identification approaches.

This change in behaviour may result from changes in the proximity of working and leisure hours and/or changes in alcohol consumption and the likelihood of the effects of intoxication being felt during working hours. We provide further evidence that the change in legislation had a causal effect on individual health. UK evidence demonstrates that this is most acute for those who report regularly attending licensed premises. This, coupled with evidence of an increase in alcohol expenditure at bars, is suggestive that the channel of effect is through alcohol consumption. In turn, this indicates that the policy in England and Wales did not have the desired effect of reducing health problems related to drinking.

How governments influence work-leisure decisions is typically thought of through the lens of income and substitution effects. However, governments also often intervene and regulate leisure activity directly. Our results suggest that government intervention in the regulation of leisure activities has the potential to have unintended consequences on labour supply decisions. An important implication of our paper then is that governments influence leisure-work trade-offs not only through taxation and transfer payments but also through direct regulation of leisure activities.

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Figure 1. Minutes of Absence and absence rate for workers in Spain

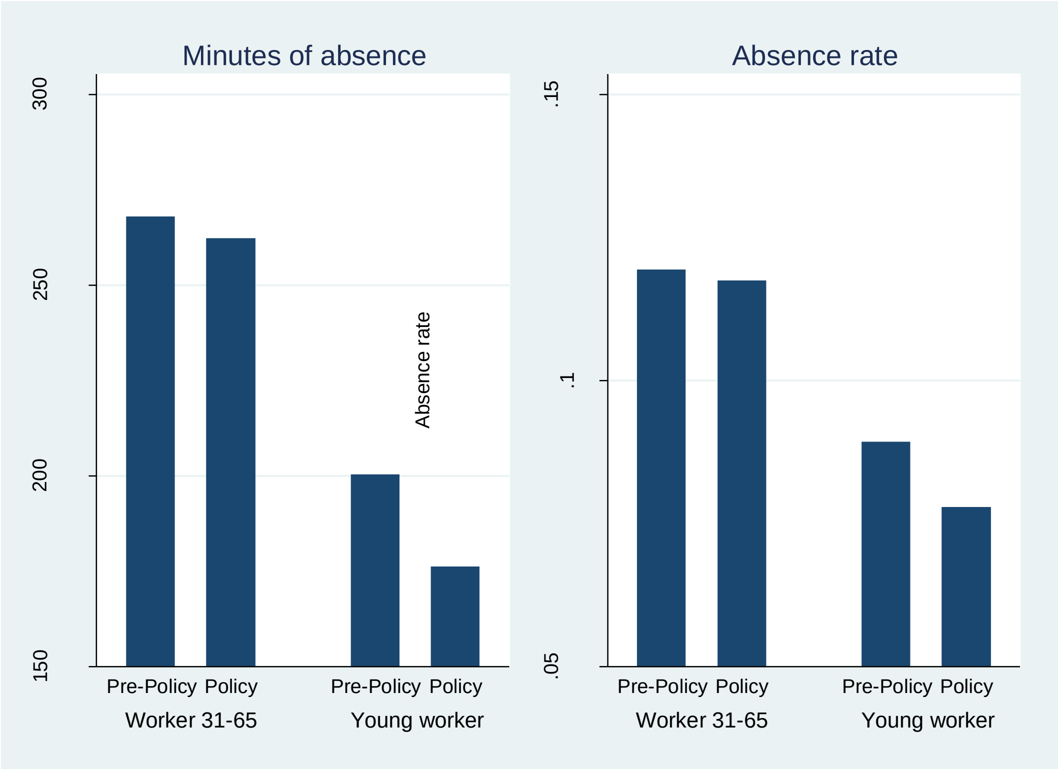

Figure 2. Minutes of absence and absence rate for workers in the UK

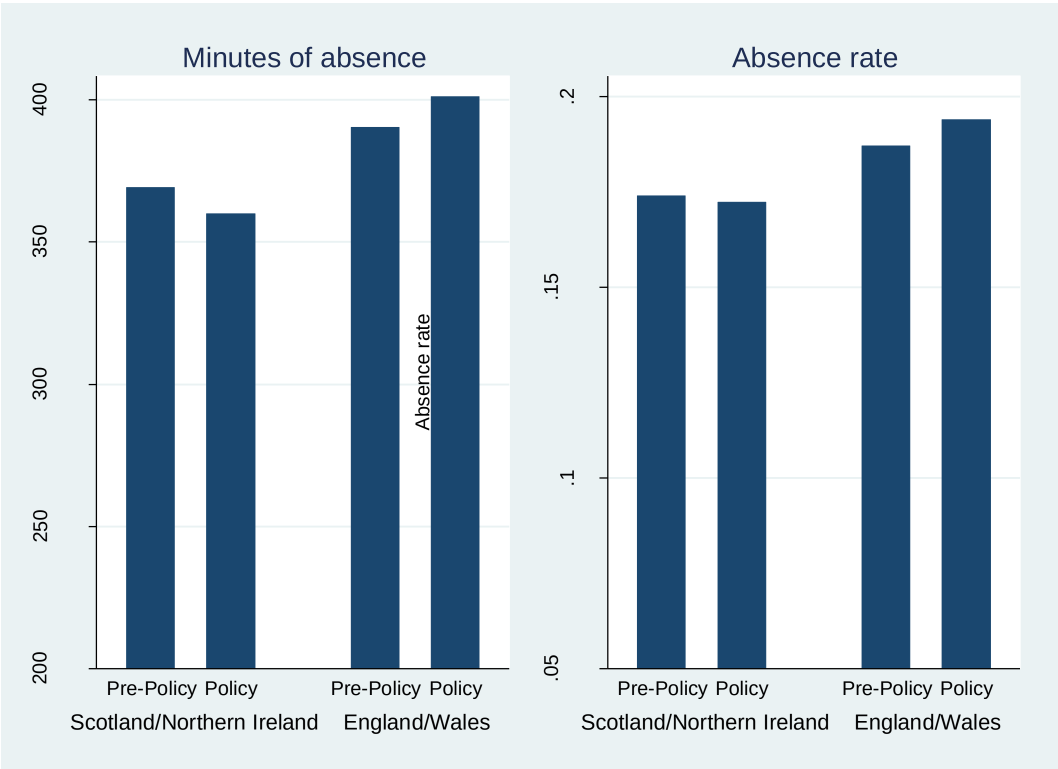

Table 1. Regional Timing of Drinking Hours Law Changes in Spain and England/Wales.

Regions (CCAA)	Law came into force	Law	Closing time
Spain			
Andalucia	1st quarter 2003	Ley 13/1999, de 15 de diciembre, de Espectáculos Públicos y Actividades Recreativas de Andalucía (BOE núm. 15, de 18 de enero), modificada por la Ley 10/2002, de 21 de diciembre (BOE núm. 14, de 16 de enero de 2003).	3:00am*
Aragon	1st quarter 2006	Ley 11/2005, de 28 de diciembre, reguladora de los espectáculos públicos, actividades recreativas y establecimientos públicos de la Comunidad Autónoma de Aragón (BOE núm. 23, de 27 de enero).	3:30am*
Canary Islands	2nd quarter 2002	Ley 1/1998, de 8 de enero, de Régimen Jurídico de los Espectáculos Públicos y Actividades Clasificadas (BOE núm. 27, de 31 de enero). Corrección de errores en BOE núm. 68, de 20-03-98 y modificada por la Ley 2/2002, de 27 de marzo (BOE núm. 97, de 23 de abril).	3:30am
Castilla Leon	4th quarter 2006	Ley 7/2006, de 2 de octubre, de espectáculos públicos y actividades recreativas de la Comunidad de Castilla y León (BOE núm. 272, de 14 de noviembre).	3:00am
Comunidad de Madrid	3rd quarter 2002	Ley 17/1997, de 4 de julio, de Espectáculos Públicos y Actividades Recreativas (BOE núm. 98, de 24 de abril de 1998), modificada por la Ley 24/1999, de 27 de diciembre (BOE núm. 48, de 25 de febrero de 2000), por la Ley 5/2000, de 8 de mayo (BOE núm. 126, de 26 de mayo) y por la Ley 5/2002, de 27 de junio (BOE núm. 176, de 24 de julio).	3:00am**
Navarra	2nd quarter 2004	Ley Foral 2/1989, de 13 de marzo, Reguladora de los Espectáculos Públicos y Actividades Recreativas (BOE núm. 84, de 8 de abril), modificada por la Ley Foral 26/2001, de 10 de diciembre (BOE núm. 39, de 14 de febrero de 2002). 27 de octubre de 2003, 656/2003 Decreto Foral (BON145 de 14/11/2003), entrada en vigor 1 de abril de 2004.	3:30am**
Comunidad Valenciana	1st quarter 2004	Ley de las Cortes Valencianas 4/2003, de 26 de febrero, de los Espectáculos Públicos, Actividades Recreativas y Establecimientos Públicos (BOE núm. 81, de 4 de abril). Ley 4/2003, de 26 de febrero, Orden de 19 de diciembre de 2003, entrada en vigor en 2004.	3:30am
Balearic Islands	2nd quarter 1999	Ley 7/1999, de 8 de abril, de Atribución de Competencias a los Consejos Insulares de Menorca y de Eivissa i Formentera en materia de Espectáculos Públicos y Actividades Recreativas (BOE núm. 124, de 25 de mayo).	3:00am
La Rioja	4th quarter 2000	Ley 4/2000, de 25 de octubre, de Espectáculos Públicos y Actividades Recreativas. (BOE núm. 287, de 30 de noviembre).	3:30**
Pais Vasco	3rd quarter 1998	Ley 4/1995, de 10 de noviembre, de la Comunidad Autónoma del País Vasco, sobre normas reguladoras de Espectáculos Públicos y Actividades Recreativas (BOE núm. 230, de 1 de diciembre). 210/1998 de 28 de Julio 1998.	2:00am*
Asturias	1st quarter 2005	Ley 8/2002, de 21 de octubre, de Espectáculos Públicos y Actividades Recreativas. (BOE núm. 278, de 20 de noviembre). Decreto 90/2004, de 11 de noviembre, por el que se regula el regimen de horarios de los establecimientos, locales e instalaciones para espectáculos públicos y actividades recreativas en el Principado de Asturias.	3:30am*
UK			
England and Wales	24th November 2005	Licensing Act 2003	

Source: http://www.mir.es/SGACAVT/juegosyespec/espectaculos/legislacionxCA.html and BOE for the case of Spain and the Licensing Act 2003 for the UK.
* Fridays and Saturdays are allowed to stay open for an hour more.
** Fridays and Saturdays are allowed to stay open for half an hour more.

Table 2. Effect of licensing laws on worker absence behaviour in Spain, 1996-2007.

	Absence rate		Minutes difference	
	Treatment $\leq$25yrs	Treatment$\leq$30yrs	Treatment $\leq$25yrs	Treatment $\leq$30yrs
Policy $\times$Treatment	-0.046	-0.035	-17.016	-14.486
	(0.015)***	(0.015)**	(5.392)***	(5.687)**
Treatment	0.160	0.162	62.306	61.956
	(0.052)***	(0.020)***	(14.577)***	(8.263)***
Policy	0.067	0.063	18.408	18.101
	(0.040)*	(0.042)	(9.275)*	(9.969)*
Age	0.011	0.008	5.909	4.445
	(0.005)**	(0.002)***	(1.392)***	(0.774)***
Age^2	-0.002	0.001	-2.713	-1.361
	(0.005)	(0.002)	(1.430)*	(0.893)
Female	0.101	0.115	18.771	26.264
	(0.012)***	(0.011)***	(4.248)***	(3.938)***
Secondary education	-0.011	-0.010	-7.108	-6.173
	(0.009)	(0.008)	(2.902)**	(2.445)**
Higher education	-0.003	0.001	-2.925	-1.562
	(0.008)	(0.006)	(2.335)	(1.847)
Public sector	0.195	0.174	81.415	73.474
	(0.013)***	(0.011)***	(2.933)***	(2.544)***
Temporary contract	-0.126	-0.126	-50.637	-51.532
	(0.008)***	(0.008)***	(2.172)***	(2.509)***
Observations	918149	1181051	918349	1181269

Note: Treatment is defined as worker aged $\leq$ 25 or 30 years old where the comparison group are workers older than 45 years old. Policy takes value 1 if region j at time t has shortened hours. Controls for marital status, industry, workers' occupation, establishment size, region, year, and quarter are included but not reported. Standard errors clustered at a regional level (18 regions) are reported in parentheses.
*, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 3. Effect of Licensing Laws on Absence Behaviour in the UK, 2003-2008

	Absence rate	**Minutes difference**
Policy $\times$ Treatment	0.030	14.079
	(0.018)*	(0.018)***
Treatment	0.090	36.603
	(0.085)	(0.637)**
Policy	-0.043	15.765
	(0.020)**	(1.687)*
Age	0.014	5.682
	(0.002)***	(0.261)**
Age^2	-0.000	-0.060
	(0.000)***	(0.000)***
Female	0.080	52.720
	(0.005)***	(4.312)*
Degree or higher	0.091	37.605
	(0.006)***	(0.187)***
Vocational training/Diploma	0.092	40.011
	(0.010)***	(0.988)**
A-Levels	0.068	26.611
	(0.008)***	(1.757)**
Temporary contract	-0.005	-41.470
	(0.011)	(2.759)**
Part time job	0.050	-180.902
	(0.011)***	(9.573)**
Public sector	0.095	64.877
	(0.010)***	(1.965)**
Observations	268654	269350

Note: Treatment corresponds to workers in England/Wales and the comparison group are workers in Scotland and Northern Ireland. Policy takes value 1 if worker *i* is observed after the 24th November 2005. Time period: 2003-2008. Controls for marital status, presence of dependent children, industry, workers' occupation, year, and quarter are included but not reported. Standard errors clustered at a regional level (20 regions) are reported in parentheses.
*, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 4. Effect of licensing laws on absence behaviour in England and Wales, Fixed Effects Estimates, 2004-2008

	Absence rate	**Minutes difference**
Policy	0.015	21.083
	(0.004)***	(9.964)*
Degree or higher	0.006	-0.475
	(0.011)	(24.043)
Vocational training/Diploma	0.015	30.801
	(0.006)**	(14.538)*
A-Levels	0.013	28.390
	(0.004)***	(9.691)***
Temporary contract	-0.002	-16.506
	(0.005)	(7.597)**
Part time job	-0.029	-208.496
	(0.004)***	(9.463)***
Public sector	0.011	16.153
	(0.006)*	(13.506)
Observations	328335	330016
Number of individuals	87931	88072

Note: Policy takes value 1 if worker *i* is observed after the 24th November 2005. Controls for marital status, presence of dependent children, industry, workers' occupation, region and quarter are included but not reported. Robust standard errors are in parentheses. Time period: 2004-2008. *, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively. Estimates for our full period, 1997-2008, are the same.

Table 5. Effect of Changes in Drinking Laws for UK and Spain, Full-Time Workers Only and Males vs Females

	Full time workers		All workers			
	Total		Male		Female	
Spain (DD)	AR	MD	AR	MD	AR	MD
Policy × Treatment	-0.033	-15.106	-0.040	-17.752	-0.029	-10.716
	(0.017)**	(5.900)**	(0.014)***	(4.947)***	(0.027)	(9.706)
Treatment	0.134	60.510	0.051	22.008	0.367	122.661
	(0.018)***	(8.066)***	(0.018)***	(6.628)***	(0.032)***	(11.042)***
Policy	0.062	19.302	0.058	17.903	0.072	18.462
	(0.040)	(10.310)*	(0.038)	(9.586)*	(0.052)	(12.293)
Observations	1064203	1064365	730771	730915	450280	450354
UK (DD)						
Policy × Treatment	0.041	17.471	0.022	-0.883	0.040	29.585
	(0.026)	(0.120)***	(0.027)	(0.181)	(0.016)**	(0.157)***
Treatment	0.079	46.558	0.080	50.232	0.101	23.323
	(0.080)	(0.647)***	(0.074)	(1.114)**	(0.099)	(0.792)**
Policy	-0.053	12.141	-0.039	14.191	-0.049	8.067
	(0.028)*	(2.617)	(0.028)	(0.569)**	(0.018)***	(3.119)
Observations	195638	196095	131556	131931	137098	137419
UK (FE)	0.012	27.912	0.005	7.493	0.025	34.663
	(0.004)***	(10.015)**	(0.004)	(10.636)	(0.005)***	(11.307)***
Observations	210955	211513	161984	162815	166351	167201

Note: All controls as per tables 2 & 3. Spain: Treatment is defined as worker aged ≤ 30 years old where the comparison group are workers older than 45 years old. Policy takes value 1 if region j at time t has shortened hours. Controls for marital status, industry, workers' occupation, establishment size, region, year, and quarter are included but not reported. Standard errors clustered at a regional level (18 regions) are reported in parentheses. Time period: 1996-2007

UK: Treatment corresponds to workers in England/Wales and the comparison group are workers in Scotland and Northern Ireland. Policy takes value 1 if worker *i* is observed after the 24th November 2005. Controls for marital status, presence of dependent children, industry, workers' occupation, year, and quarter are included but not reported. Standard errors clustered at a regional level (20 regions) are reported in parentheses. Time period: 2003-2008.

*, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 6. Effect of Licensing Laws on Absence Behaviour for Spain and the UK; Robustness test for Implementation Effects (DiD estimates excluding year of policy implementation).

	Spain	UK
Absence rate		
Policy $\times$ Treatment	-0.032	0.030
	(0.015)**	(0.018)*
Treatment	0.165	0.090
	(0.020)***	(0.085)
Policy	0.084	-0.043
	(0.047)*	(0.020)**
Observations	1096150	268654
Hours difference		
Policy $\times$ Treatment	-14.490	14.079
	(5.782)**	(0.018)***
Treatment	62.527	36.603
	(8.615)***	(0.637)**
Policy	24.464	15.765
	(11.003)**	(1.687)*
Observations	1096360	269350

Note: All controls as per tables 2 & 3. Spain: Treatment is defined as worker aged $\leq$ 30 years old where the comparison group are workers older than 45 years old. Policy takes value 1 if region j at time t has shortened hours. Controls for marital status, industry, workers' occupation, establishment size, region, year, and quarter are included but not reported. Standard errors clustered at a regional level (18 regions) are reported in parentheses. Time period: 1996-2007

UK: Treatment corresponds to workers in England/Wales and the comparison group are workers in Scotland and Northern Ireland. Policy takes value 1 if worker *i* is observed after the 24th November 2005. Controls for marital status, presence of dependent children, industry, workers' occupation, year, and quarter are included but not reported. Standard errors clustered at a regional level (20 regions) are reported in parentheses. Time period: 2003-2008.

*, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 7. Effect of Licensing Laws on Absence Behaviour, Collapsed Samples, Spain and the UK.

Absence rate		
Policy $\times$ Treatment	-0.008	0.019
	(0.004)**	(0.008)**
Treatment	0.010	0.036
	(0.006)*	(0.032)
Policy	-0.015	-0.013
	(0.010)	(0.007)*
Observations	285166	88707
Hours difference		
Policy $\times$ Treatment	-21.771	17.438
	(5.988)***	(4.544)***
Treatment	41.679	32.131
	(7.699)***	(35.198)
Policy	40.803	-20.180
	(8.768)***	(1.775)***
Observations	285242	88760

Note: All controls as per tables 2 & 3. Spain: Treatment is defined as worker aged $\leq$ 30 years old where the comparison group are workers older than 45 years old. Policy takes value 1 if region j at time t has shortened hours. Controls for marital status, industry, workers' occupation, establishment size, region, year, and quarter are included but not reported. Standard errors clustered at a regional level (18 regions) are reported in parentheses. Time period: 1996-2007

UK: Treatment corresponds to workers in England/Wales and the comparison group are workers in Scotland and Northern Ireland. Policy takes value 1 if worker *i* is observed after the 24th November 2005. Controls for marital status, presence of dependent children, industry, workers' occupation, year, and quarter are included but not reported. Standard errors clustered at a regional level (20 regions) are reported in parentheses. Time period: 2003-2008.

*, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 9. Effect of Licensing Laws on Health Problems in the UK, BHPS 1997-2007.

	Total	Male	Female	Goes Out to Drink		
				At least		< than once a month
				Weekly	Monthly	
Policy × Treatment	0.020	0.018	0.021	0.050	0.020	-0.005
	(0.008)***	(0.010)*	(0.006)***	(0.021)**	(0.016)	(0.022)
Treatment	0.039	0.027	0.049	0.025	0.035	0.055
	(0.025)	(0.019)	(0.030)	(0.030)	(0.019)*	(0.029)*
Policy	-0.014	-0.018	-0.015	-0.050	-0.061	0.011
	(0.006)***	(0.012)	(0.004)***	(0.013)***	(0.022)***	(0.016)
Age	0.015	0.019	0.013	0.011	0.013	0.026
	(0.003)***	(0.003)***	(0.003)***	(0.003)***	(0.002)***	(0.002)***
Age^2	-0.000	-0.000	-0.000	0.000	-0.000	-0.000
	(0.000)**	(0.000)***	(0.000)**	(0.000)	(0.000)	(0.000)***
Female	0.064			0.093	0.069	0.027
	(0.011)***			(0.015)***	(0.009)***	(0.010)***
Public sector	-0.012	0.001	-0.022	-0.011	-0.016	0.005
	(0.012)	(0.012)	(0.020)	(0.017)	(0.013)	(0.022)
A-Levels	-0.006	-0.004	-0.013	-0.012	-0.008	-0.007
	(0.009)	(0.015)	(0.012)	(0.010)	(0.008)	(0.016)
Vocational training/Diploma	-0.029	-0.037	-0.024	-0.008	-0.024	-0.068
	(0.015)*	(0.020)*	(0.022)	(0.025)	(0.017)	(0.027)**
Degree or higher	-0.046	-0.058	-0.042	-0.035	-0.041	-0.060
	(0.010)***	(0.013)***	(0.013)***	(0.014)**	(0.011)***	(0.019)***
Temporary contract	0.033	0.013	0.046	0.035	0.018	0.023
	(0.012)***	(0.014)	(0.014)***	(0.023)	(0.011)*	(0.025)
Drink often	0.016	0.018	0.017			
	(0.002)***	(0.003)***	(0.002)***			
Observations	58429	26858	31571	18925	123593	13493

Note: All controls as per tables 2 & 3. Treatment corresponds to workers in England/Wales and the comparison group are workers in Scotland and Northern Ireland. Policy takes value 1 if worker *i* is observed after the 24th November 2005. Standard errors clustered at a regional level (12 regions) are reported in parentheses.

*, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

APPENDICES:

Table A1. Descriptive statistics

	Spain			UK	
	Mean	Std		Mean	Std
Minutes of absence	239.226	601.700	Minutes of absence	390.615	728.538
Absence rate	0.105	0.264	Absence rate	0.187	0.336
Age	40.352	14.915	Age	41.608	11.830
Female	0.369	0.482	Female	0.511	0.500
Married	0.533	0.499	Married	0.615	0.487
Primary education	0.532	0.499	A-Levels	0.217	0.412
Second education	0.193	0.395	Vocational training/Diploma	0.140	0.347
Higher education	0.274	0.446	Degree or higher	0.218	0.413
Public sector	0.164	0.370	Public sector	0.320	0.467
Temporary contract	0.455	0.498	Temporary contract	0.050	0.218
			Part time job	0.273	0.446
			Dependent children	0.799	1.090
Observations	1719510			472017	