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

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Abstract

The regulation of alcohol availability has the potential to influence worker productivity. In particular, opening hours of bars could affect worker effort through the proximity of leisure consumption to working hours and the timing of alcohol consumption spilling over into working hours. This paper uses legislative changes in bar opening hours to provide a potential quasi-natural experiment of the effect of alcohol availability on working effort, focusing on worker absenteeism. 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 is demonstrated between opening hours and absenteeism. However, the decrease in absence caused by shorter opening hours in Spain is short-lived, disappearing after two years. The effect is more robust for the UK where we provide evidence which suggests that increased alcohol consumption is a key mechanism.

KEY WORDS: Labor Supply, Absenteeism, Alcohol Policy
JEL Classification: J22, K32

[*] The authors would like to thank Silke Anger, Steve Bradley, Paul Devereux, Mirko Draca, Javier Gardeazabal, Juan Ramon Garcia, John Heywood, Gareth Leeves, Owen O'Donnell, Efthymios Pavlidis, Aydogan Ulker, Ian Walker, Anwen Zhang, seminar participants at the University of Wisconsin-Milwaukee, participants at the 4th Annual Meeting on the Economics of Risky Behaviors, the 11th IZA/SOLE Transatlantic Meeting, the 2012 EEA conference in Malaga, the 2012 EALE conference in Bonn and the 2013 RES conference in Royal Holloway for helpful comments. María Navarro gratefully acknowledges financial support from the Spanish CYCIT Research Project ECO2008-06395-C05-03/ECON. Corresponding author: Colin Green, Economics Department, Lancaster University, Lancaster, LA1 4YX, UK. e-mail: c.p.green@lancaster.ac.uk. Maria Navarro, Economics Department, Lancaster University, Lancaster LA1 4YX UK, e-mail: m.navarropaniagua@lancaster.ac.uk.

The data was provided by the U.K. Data Archive. The data are available on request, subject to registering with the Data Archive. The usual disclaimer applies.

I. INTRODUCTION

The regulation of alcohol consumption remains a highly contentious area of public policy and has generated a large literature in both public health and economics (see Anderson et al, 2009 and Carpenter and Dobkin, 2011a for recent reviews). Policies aimed at regulating alcohol are numerous including taxation, minimum pricing, age based restrictions, place-based restrictions and restrictions on the timing of sales. Timing restrictions can take many forms, including restrictions on permissible days and hours of alcohol availability both on and off premises. For instance, day restrictions include laws that prevent Sunday sales of alcohol in the United States and Australia; and Saturday sales in Scandinavia (Gronqvist and Niknami, 2011; Heaton, 2012). This type of regulation appears to influence consumption behavior. For instance, Carpenter and Eisenberg (2009) show that allowing Sunday sales in Ontario, Canada, increased Sunday-specific drinking by 7-15%, mainly in the form of substitution from drinking on Saturdays. At the same time, restrictions of the timing of on premise drinking are the norm, but with a wide variation in actual opening hours across jurisdictions. The aim of these laws is not solely to restrict alcohol availability, but also to reduce negative externalities from leisure behavior such as noise pollution and disruption to residents near venues. This regulation of opening hours has been shown to influence a range of health and socio-economic outcomes including alcohol consumption (Bernheim, Meer and Novarro 2012), traffic accidents (Green *et al*., 2013; 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).

Whilst mainly aimed at reducing social externalities associated with excessive consumption, alcohol regulations also have the potential to influence labor market outcomes. In terms of alcohol taxation this idea has attracted some attention. For instance, Johansson *et al* (2012) examine the effect of a cut in alcohol taxation in Finland which led to large differences in alcohol prices between Finland and Sweden. They show that this had a weak effect on mortality and alcohol related illnesses but substantially increased workplace absenteeism, a 5% increase for males and a 13% increase for females in Swedish regions near to the Finnish border when compared to Swedish regions that are over 100 km away from the border. In contrast, Dave and Kaestner (2002) find no evidence that alcohol taxes are related to labor market outcomes such as employment, hours of work and wages in the US.

Restrictions on opening hours also have the potential to influence labor market outcomes, but this has received little attention to date. Traditionally, the effect of government regulation of leisure is thought of through the lens of income and substitution effects (see for instance Burtless and Hausman, 1978; Heckman, 1993; Blundell, Duncan and Meghir, 1998). However, individual labor supply behavior could be influenced directly by leisure regulation if, for instance, it affects the temporal proximity of leisure consumption and working hours (Biddle and Hamermesh, 1990). Moreover, the timing of consumption could also have spill-over effects into working hours when intoxicating substances like alcohol are involved. Along these lines, previous evidence suggests that alcohol consumption is linked to absenteeism (Balsa and French, 2010; Johansson *et al*, 2008; Norstrom, 2006; and Norstrom and Moan, 2009).

This paper investigates 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, it has been previously suggested that other forms of alcohol legislation such as minimum drinking ages and reductions in alcohol taxation can influence workforce productivity (Carpenter and Dobkin, 2011a; Johansson et al, 2012). We use recent changes in legal pub and club (herein bars for simplicity) opening hours in the UK and Spain as 'quasi-natural' experiments to identify the effect of on-premises alcohol availability on absence.[1] These two legislative changes provide a nice point of comparison, as one involves a substantial liberalization 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 behavior due to the proximity of leisure activity to normal working hours, but also through the timing of the consumption of alcohol.

To summarize our results, we demonstrate a causal link between bar opening hours and worker absenteeism, longer opening hours increase absence. In terms of the direction of the effects, the results are symmetric for the UK and Spain: increasing opening hours (UK) increases absenteeism, decreasing opening hours (Spain) reduces absenteeism. We identify the causal effect of opening hours on absenteeism using difference in difference approaches. In addition, we demonstrate the robustness of our results to standard concerns derived from applying a difference-in-difference methodology such as violations in the common trend

[1] Absenteeism can be viewed as a proxy for worker effort (Audas et al, 2004).

assumption and appropriate control groups. While, the UK policy effect appears robust, for Spain the effect appears to be relatively short-lived and concentrated amongst early adopting regions. Finally, we provide evidence on the mechanisms through which the policy influenced worker absenteeism for the UK. This evidence is suggestive of a central role for increased on premise alcohol consumption.

II. DATA AND INSTITUTIONAL BACKGROUND

Changes in Drinking Laws, UK and Spain

The identification strategy in this paper is based on two legislative changes; an extension of legal closing hours in two constituent parts of the UK, England and Wales and a reduction in the permitted hours that bars could remain open in Spain. 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 number had increased to 78879 venues. Hence, the main expansion occurred in the initial time period that the legislation was enacted.

In the Spanish case, the 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 in different periods regionally across Spain, and varied in terms of the actual new time of closing ranging from 2:00 am to 3:30 am. 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, a degree of 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. Importantly, other related policies, such as alcohol taxation and prices, age based regulation, location restrictions and drink driving policies and sanctions are enacted only at a national

level. This limits the potential of our policy estimates to be influenced by confounding changes in other related alcohol policy. 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).

INSERT TABLE 1

It is also worth noting that the stated reasons for these two legislative changes were markedly different. In the UK, it reflected a view that the prior regime of an 11pm closing time was needlessly restrictive and that shorter opening hours may encourage binge drinking insofar as individuals would increase the speed of alcohol consumption (IAS, 2007). While in Spain, it primarily reflected concerns over noise pollution and general disruption to residents near licensed venues. Finally, the margin at which these changes occurred means that the policy changes may affect different types of individuals in the UK and Spain. Extending hours from 11pm is likely to hit a broad cross-section of individuals. In contrast, reductions from 6am to 3am are more likely to be binding for younger individuals: they are simply more likely to attend bars at these times.

Data

This two main data sources used in this paper are very similar in their basic structure, the UK Longitudinal Labor Force Survey (UK LFS) and the Spanish Labor Force Survey (SLFS). Both are quarterly representative surveys that provide a range of information on individual and work 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.

For the UK, we use the 5 quarter rotating cohort version of the LFS. 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). We focus on the period 1997-2008. This provides 846,106 observations for 218,405 different individuals. 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 our estimating sample consists of 1,993,260 observations.

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 $A_{it} = H_{it}^{u} - H_{it}^{e}$.[2] 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. This is defined as the ratio of the hours reported absent to contractual hours in the reference week $AR_{it} = A_{it} / H_{it}^{u}$. A potential issue is that these measures of absenteeism may be affected by changes 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 UK LFS and the SLFS 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 and this is discussed in more detail in the results section.

INSERT FIGURE 1

Figures 1 and 2 provide some illustrative information on absence trends in our two settings. Figure 1 shows absence trends before and after the policy changes for England/Wales and Scotland/Northern Ireland, respectively. Both panels of Figure 1 provide some indication that the absence trends for England/Wales and Scotland/Northern Ireland follow similar patterns prior to the policy change. After the policy there is a clear divergence in behavior; workers in England/Wales take relatively more absence, while for Scotland/Northern Ireland the trend appears stable.

INSERT FIGURE 2

Figure 2 plots similar data for Spain. A complication here in terms of visualisation is that we have variation in the timing of policy implementation by region. Figure 2 shows raw data plotting our absence measures for treatment regions *prior* to the policy change and a 'control' series that consists of regions that never implemented the policy in our sample period. What is apparent is that these series are very similar in both trends and levels.

Both data sets have quite a rich set of potential control variables, including many of the candidates that have been shown to be important determinants of worker absenteeism in previous research. Thus, we incorporate socio-demographic variables, including the age and

[2] 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 (2012), among others.

the age squared, gender, marital status, education level. We also include 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. Finally, region, year and quarter dummies are introduced to take account of regional, time and seasonal variation.

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. In addition, we focus on only those workers who are beyond the minimum legal drinking age, excluding those who are less than 18 years old. Appendix Table A1 provides summary statistics for the resultant samples for both the UK and Spain.

III. METHODOLOGY AND IDENTIFICATION

We estimate the effect of changes in opening hours on absenteeism using variants of the following reduced form model:

$$A_{ijt} = \alpha + \beta Policy_{ijt} + \gamma X_{ijt} + \mu_j + \tau_t + t\mu_j + \varepsilon_{ijt} \qquad (1)$$

where A_{ijt} corresponds to the minutes of absence of worker *i* in region *j* in period *t*. The region and year fixed effects are captured by μ_j and τ_t, respectively. In the UK setting $Policy_{ijt}$ equals one for treated individuals (workers residing in England or Wales) in the post-treatment period (after the 24th of November 2005). The OLS estimate of β in equation (1) is equivalent to the Differences-in-Differences (DD) estimator and this provides an estimate of the increase in absence caused by the change in licensing laws for workers in England or Wales compared to those living in Scotland or Northern Ireland. In the UK-LFS we can observe the week of interview and thus we can identify this policy effect separately from quarter controls aimed to pick up seasonality in absence.

For Spain, we have regional and time variation in the adoption of the legislative change. We observe 11 regions (out of 18 in total) closing bars earlier in different periods. The models are identified using regional variation in the timing of policy adoption, controlling for differences across regions that were not treated over the same time period. $Policy_{ijt}$ takes the value 1 if the worker is observed in region *j* and at time *t* that region has

reduced drinking hours, its accompanying parameter is equivalent to the Differences-in-Differences (DD) estimator.

In both countries, region fixed effects μ_j are included for the model to absorb all persistent unobserved regional characteristics that may be correlated with the timing of the introduction of the policy and with absenteeism. This could for instance be difference in absence patterns between regions in the south and in the north. In a similar way, the year fixed effects τ_t removes national time variation in absenteeism common for all regions.

The key identifying assumption for any DD strategy is that the outcome in treatment and control group would follow the same trend in the absence of treatment. At first glance, the pre-treatment trends displayed for the UK in Figure 1 and for Spain in Figure 2 appear similar. However, as differences in absence trends pre-policy may be confounded with the policy effect we extend our specification further by re-estimating our main models incorporating region-specific linear time trends $t\mu_j$ to allow for different trends across regions. We do this by interacting the dummy for each region with a variable quarterly linear time trend that equals 1-12 the first year 13-24 the second year and so forth.[3]

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. One approach is to introduce robust standard errors clustered at the regional level. For Spain, we observe 18 regions, with 11 changing policy within our sample period. However, for the UK, while we observe 20 regions, the policy only varies at the country level. Standard results for the introduction of clustered standard errors with small numbers of groups are not encouraging (Bertrand et al., 2004). As an alternative for the UK we adopt the approach suggested by Cameron et al., (2008) and estimate our standard errors using a wild bootstrap procedure by cluster (regions) that aims to address this issue. A related concern is that while using longer panels of data help to address concerns regarding spurious estimates in DD models, the presence of serial correlation in the dependent variable will downwardly bias standard errors. In additional robustness checks we

[3] Our main results are robust to replacing this quarterly time trend with a yearly time trend.

also adopt the approach suggested by Bertrand et al (2004) and collapse our data into pre and post policy periods and re-estimate our models.

Finally, in all models we estimate variants of (1) where the dependent variable is the absence rate (AR) as defined above. This dependent variable is more flexible insofar as it explicitly allows for variations in contractual hours. This variable is censored at zero and many workers are observed at this bound. As a result, we estimate these models as a tobit. This has the advantage of being a more efficient estimator, but will lead to policy estimates which are not directly comparable to our minutes absence models in terms of magnitude of effect. To aid this comparison we also report OLS estimates of the effect of changing drinking hours on the absence rate.

IV. RESULTS

INSERT TABLE 2

Table 2 provides the reduced form estimates of the effect of the drinking law regulation in England/Wales on worker absenteeism. Three groups of estimates are reported, OLS estimates of minutes difference along with Tobit and OLS estimates of the absence rate. A vector of standard controls is included covering age, education level, sector of work, whether the individual is on a temporary contract, occupation, industry, region, year and quarter dummies. For brevity we do not report the estimates for all of these controls but these are available upon request. It is worth noting that in both the UK and Spanish 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, 2012, Ichino and Riphahn, 2005), public sector workers take more absence and female and married workers take more absence (Barmby, Orme and Treble, 1991).

There is an impact of the legislative change on absenteeism for workers in England and Wales. Our initial estimates in the first panel of Table 2 show that increasing opening hours increased worker absence by approximately 1.4 percentage points and lead to an increase in time lost through absence of 5.59 minutes per week per worker. For completeness, and to aid comparability we also report the OLS estimates of the policy effect on the absence rate. These estimates are small, 0.1 percentage points, but remain statistically significant. All standard errors are calculated using the Wild Bootstrap with assignment at the regional

cluster level (see Cameron *et al*, 2008). In the bootstrap procedure we use the Rademarcher distribution and set the number of replications to 400.

INSERT TABLE 3

Panel 1 of Table 3 provides the reduced form estimates of the effect of the drinking law regulation in Spain on worker absenteeism. Again we report separate OLS estimates for minutes difference and Tobit estimates for the absence rate. Standard errors are again calculated using a Wild Bootstrap procedure. The initial estimates (first 3 columns) contain standard controls; again suppressed for brevity, and year, quarter and region fixed effects. These estimates suggest a link between shorter drinking hours and reduced absenteeism, however the estimates are not statistically significant at standard levels.

We extend these initial specifications in both countries by including region specific time trends in addition to region and year fixed effects. The resultant estimates are reported in Panel 2 of Tables 2 and 3. The results for both the UK and Spain appear robust to controlling for violations in the common trends assumption. In both cases introducing these trends lead to an increase in the estimated impact of the policy, leading to estimates in Spain that remain negative but are now statistically different from zero. This is our preferred specification and all subsequent reported estimates will be based on this. For brevity, we omit the OLS estimates of the absence rate from this point on, but these are available from the authors on request.

INSERT TABLE 4

Due to the pattern of bar attendance, which is likely to be more concentrated among young people, we may expect treatment heterogeneity according to the age of workers. This may be amplified in the Spanish case due to the margin at which the legislative adjustment was made, young people are not only more likely to attend bars, but more likely to attend bars up to 6 am prior to the reform. This second effect may be less marked in the case of England and Wales due to the change being from 11pm. We re-estimated our main models for three groups of workers; those aged 18 to 30 years, 31 to 45 years and those aged 46 years or older, and this is reported in Table 4. There is, perhaps, surprisingly little age variation, although the pattern is more apparent for Spain where there is some suggestion that older workers were less affected by the policy change. Due to a lack of precision, however, none of these estimates is statistically different from each other. There is essentially no pattern in

terms of minutes absence in the UK, but some indication of an age gradient in the absence rate models.

Threats to Identification and Inference

Our results demonstrate that drinking law regulations have the potential to influence worker absenteeism. They suggest that longer hours in England and Wales increased absenteeism approximately 11 minutes per week; while shorter hours in Spain lead to a reduction of 17 minutes per week. A strength of our approach is that the effect is found for two different countries where the policy was operating in different directions. Nonetheless, in this section we examine a range of potential threats to our identification strategy, and in turn discuss standard concerns with inference in a difference in difference setting. We start by summarising a range of alternative estimates in Table 5, which we discuss in turn.

INSERT TABLE 5

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 bias our policy estimates. The first row of Table 5 provides estimates where our samples are restricted to full-time workers only. The resultant point estimates are essentially unchanged. 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. Using information in the SLFS and UK LFS on reasons for variation of working hours we exclude categories that were less likely to be in the control of workers and hence, could be affected by the policy. Specifically, we excluded those workers whose hours `usually vary' along with absence due to changing or loss of job, undertaking training, and union representation, strike or labor conflict and technical partial stop or employment regulation within a firm because of financial problems. Again the policy estimates are essentially unchanged. Insofar as the remaining reasons for absence can more readily be controlled by workers this makes us more confident that our estimates are picking up worker responses to changes in drinking hours.

We next investigate the sensitivity of our results to a range of issues related to regional variation in the timing of implementation and the extent of expansion. Unlike Spain, the change in licensing in England and Wales was, in effect, not mandatory. Individual venues had to apply for an additional licence to remain open later. This leads to potential

variation in the intensity of treatment. In areas where there is a greater density of venues that increased hours, we might expect a larger absence response. The UK Department for Culture, Media and Sport reports the number of licenses granted by region. Most regions have quite a similar density of extended hours licences per head of population (18 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 models for these regions only (again using Scotland and Northern Ireland as control groups) and these reveal higher estimates of the policy effect than those for England and Wales in total, for instance the estimate of 11 minutes rises to 27.27 minutes. This suggests that areas that were more intensively treated by the reform suffered greater increases in absenteeism.

A concern may be that border crossing between England and Scotland to consume alcohol could contaminate our results. This seems unlikely to be a concern due to geography. The Anglo-Scottish border is sparsely populated, there is only one city near the border on the English side, Carlisle, which has a population of 75,000 people, and relatively few small towns on the Scottish side within 50 kilometres from the border. This suggests that border crossing is unlikely to be a major factor in our estimates. Nonetheless we re-estimated our models excluding the regions in England bordering Scotland (Tyne and Wear and Rest of the Northern Region) and our main estimates were unaffected.

In the case of Spain, there was discretion in the timing of the adoption of the policy, as per 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 main models up to 2002 such that our policy estimate is only for those three regions that adopted early, La Rioja, Balearic Islands and the Basque Country. The estimates for these early adopters are much higher than our main results.

We next examine whether these policy effect on absence behaviour was short-lived or long-lasting. The initial approach taken is to simply exclude 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, while for the UK this amounts to removing the 24th of November 2005 to the 23rd of November 2006**.** For the UK there is essentially no difference between the resultant estimates and our previous results. This suggests an effect of extending hours that does not immediately diminish. This fits with the pattern of absence observed in Figure 1. For Spain,

however, the policy effect essentially halves in magnitude and is no longer statistically significant at standard levels.

While our results for Spain are robust to including region specific linear time trends, the fact that the policy effect is bigger for early adopters and that our results are not statistically significant when omitting the year of policy implementation is suggestive of a short term policy effect that may be concentrated in early adopters. It could be that regions more concerned with worker absenteeism implemented the policy earlier. This is investigated further by using a regression-adjusted event study in the spirit of Autor (2003). The key estimates are presented pictorially as Figure 3 with 95% confidence intervals around the point estimates. The regression that generates these estimates is our main specification where the policy indicator is replaced by a series of leads and lags. There are two striking features of the resultant pattern. First, there is some evidence of anticipation. There is a reduction in absenteeism in the year prior to adoption that is statistically different from zero at the 5% level. Nonetheless the majority of the reduction occurs in the year of implementation and the year afterwards. Both of these estimates are statistically different from zero at the 5% level. The estimates for these two years are also statistically different from that of the year prior to implementation. The other clear result is that the absence 'benefits' from shorter opening hours disappear completely within 2 years. Hence, workers in Spain respond to shorter opening hours in the short run with less absenteeism, there is no long term effect.

INSERT TABLE 6

In Table 6 we pursue a number of further robustness checks which relate specifically to the UK. First, one advantage of our UK data is the presence of worker fixed effects. This allows us to examine the potential for our estimates to be influenced by compositional change in the workforce pre and post policy. Worker level fixed effects are introduced into our main specification and reported in Table 6. The resultant estimates again show that the extension of drinking hours substantially affects worker absence behavior.

Our UK identification strategy leans heavily on Scotland and Northern Ireland as a suitable comparison group of workers and on the specific timing of the policy change. This motivates a number of robustness checks. First, we investigate 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. This is done by exploiting the panel dimension

of the UK LFS. The resultant fixed effects estimate of the placebo policy effect for Scotland and Northern Ireland, while positive, is far from statistically significant at standard levels.

A more general concern relates to whether workers in Scotland and Northern Ireland provide a good counter-factual for workers in England and Wales. Figure 1 suggests that the levels of absence are, on average, quite similar pre-treatment between these areas. Nonetheless we investigated this further using a synthetic control method in the spirit of Abadie and Gardeazabal (2003). While there is a significant population in our control groups (approximately 5,094,800 in Scotland and 1,724,400 in Northern Ireland) the LFS only provides three identifiable regions for these geographic areas (Strathclyde, the rest of Scotland and Northern Ireland). The optimal weights from this procedure are used such that the synthetic England/Wales (minutes absence = 389.46; absence rate = 18.49) more closely resembles the actual England/Wales (minutes absence rate = 392.10; absence rate =18.67) before the extension of opening hours. The resulting estimates are larger than before, but with a loss of precision. These estimates are 35.19 minutes absence and 0.008 increase in the absence rate, where the absence rate is estimated via OLS.[4]

A standard concern with DD 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 UK data by individual into two periods, pre and post reform. We then re-estimate equation (1) on this collapsed data and the results are reported in Table 6. These results remain positive and statistically significant at the 1% level. This suggests that our previous policy inference was not incorrect due to serial correlation in absenteeism.

INSERT TABLE 6

Mechanisms

To this point, we have demonstrated a robust relationship between bar closing hours and worker absenteeism. In practice, this effect could occur through a variety of channels. For instance, the proximity of hours of leisure consumption and work could influence worker absence decisions. Likewise, increases in alcohol consumption or consumption more proximate to working hours could spill-over into working hours. In this section, we use a

[4] Where the optimal weights are 0.873 (0) Strathclyde, 0 (0.931) rest of Scotland and 0.127 (0.069) Northern Ireland for the minutes absence models (absence rate models).

variety of approaches for the UK to provide suggestive evidence on the role of alcohol consumption as a factor linking opening hours and absenteeism. Corresponding evidence for Spain is not provided for two main reasons. The first reflects the lack of a long-term robust effect demonstrated in the previous section. The second is more pragmatic, there is simply a lack of data that allows us to pursue this line of investigation for Spain.[5]

Alcohol consumption could change with opening hours in the form of more consumption by those who attend licensed venues, or due to a change in venue attendance behavior. Using representative UK data we explore both. We use the Expenditure and Food Survey (EFS) data from 2001-2007 and the Living Costs and Food Survey 2008 which provides a representative sample of household's expenditure in the UK as an annual repeated cross-section. This data 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 premises (i.e. bars, hotels and restaurants) at a household level. Using this information we can plot this individual expenditure for England/Wales and Scotland/Northern Ireland. This is reported as Figure 4 and shows a fairly stable, but small, gap between these two groups up until the time of the policy; Scotland and Northern Ireland have higher average expenditure over this period. After the policy was implemented, this difference disappears and even reverses slightly.

Likewise, the Health Survey for England (HSE) can be used to examine an indicator of heavy drinking. This representative data asks the question how many units of alcohol you consumed on your heaviest day of drinking in the last 7 days. We show the annual average of this for England as Figure 5. We also plot this separately by broad age groups. Again, there is an indication of an increase in heavy drinking that coincides with the policy implementation. This is in line with the expenditure evidence. While there is an increase for all groups, it is most marked amongst young people.

INSERT TABLE 7

[5] 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.

Another channel through which longer opening hours could influence drinking behaviour is directly through frequency of bar attendance. The British Household Panel Survey (BHPS) contains information on how often on average the interviewee goes out to licensed venues to drink. As the BHPS contains longitudinal data for England/Wales and Scotland and Northern Ireland we can use this information to estimate a DD model where frequency of going out to licensed venues is the dependent variable. Specifically, we construct the variable drink often which takes value 1 if the individual declares that they go out to licensed venues to drink several times a year, at least once a month and at least once a week and 0 if the individual goes out to license venues either once a year or less or never/almost never. These estimates suggest that extending bar hours increases the probability of going out for a drink in England and Wales by 1.2% compared to Scotland and Northern Ireland.

Likewise the expenditure information is used to estimate a simple analogue of our DD model for the UK with log alcohol expenditure per week (£) at licensed venues as the dependent variable. Extending bar hours increased alcohol expenditure on-premises by 6.8% for individuals living in England and Wales compared to those living in Scotland and Northern Ireland (Table 7). This provides evidence that greater on-premise availability increases alcohol expenditure. This is in line with previous evidence of the effect of off-premise availability on consumption provided by Carpenter and Eisenberg (2009). Moreover, our estimates are broadly similar in magnitude to the 7%-15% found in their study of Sunday trading in Canada. They found little evidence that overall alcohol consumption increased, increases in Sunday drinking were matched by reductions in drinking on other days. In our case, we estimated the effect of the increase in opening hours on *off-premise* alcohol expenditure, also available in the EFS. In unreported estimates, we found no effect on this expenditure. Hence, our results do not seem to reflect substitution of drinking between on and off premises.

This information on expenditure also allows us to indirectly estimate the causal effect of alcohol consumption on premises on absenteeism that is due to longer opening hours. This is possible due to the UK LFS and the EFS sampling the same underlying population. Specifically we are interested in estimating the effect of alcohol consumption on premises on absenteeism, using longer opening hours as a source of exogenous variation in on premise drinking such that:

$$A_{ijt} = \phi + \rho \log(alc \exp)_{ijt} + \alpha X_{ijt} + \varepsilon_{ijt} \quad (2)$$

$$\log(alc \exp)_{ijt} = \alpha + \delta Policy_{it} + \eta Treatment_{ij} + \gamma Policy_{it} \times Treatment_{ij} + \lambda X_{ijt} + v_{ijt} \quad (3)$$

Where the key parameter of interest is $\hat{\rho}$, which provides the effect of alcohol expenditure on premise on absenteeism. This model is identified by an exclusion restriction, longer opening hours in England and Wales.

The main problem is the lack of expenditure information in the UK LFS and absence data in the EFS. This can be overcome by combining these two data sets, assuming that the sample moments are independent, and estimating a two samples two stage least squares model (Angrist and Krueger, 1992) such that:

$$A_{ijt} = \phi + \delta Policy_{it} + \gamma Treatment_{ij} + \beta Policy_{it} \times Treatment_{ij} + \alpha X_{ijt} + \tau Y_t + \varepsilon_{ijt} \quad (4)$$

$$\log(alc \exp)_{ijt} = \alpha + \delta Policy_{it} + \eta Treatment_{ij} + \gamma Policy_{it} \times Treatment_{ij} + \lambda X_{ijt} + v_{ijt} \quad (5)$$

To implement this approach we first estimate the effect of extended opening hours on absenteeism (4) using UK LFS data (Columns 4 and 7 from Table 2) where A_{ijt} is the absence of individual *i* in region *j* at time *t*. Subsequently, we estimate the effect of extending bar hours on the logarithm of alcohol expenditure (5) as previously reported in Table 7.

INSERT TABLE 8

The ratio of these two estimated coefficients $\frac{\hat{\beta}}{\hat{\gamma}}$ is identical to the 2SLS ρ coefficient in equation (2) for the exactly identified case where we have as many instruments (extended opening hours in England and Wales) as potential endogenous variables (log alcohol expenditure). This will give us a LATE estimate of alcohol consumption on absenteeism for those workers that are affected by the bar hours extension. Given that our estimate is a non linear combination of estimators we apply the delta method to compute the standard errors (Van Kippersluis et al., 2011 and Devereux and Hart, 2010). The results are reported in Table 8. These demonstrate very large and statistically significant effects of the alcohol consumption changes on absence for those affected by the policy. In fact, these effects, in the order of 3 hours more absence for a 1% increase in weekly alcohol expenditure on premises seem implausibly large. Nonetheless, they provide a further suggestion that a channel of transmission of the policy effect on absenteeism is through alcohol consumption.

V. CONCLUSION

This paper sought to examine how changes in permissible bar opening hours influences individual labor supply decisions. Specifically, we used two recent changes in the legal opening hours of licensed premises in England and Wales and Spain. These are particularly advantageous insofar as they provide policy changes in opposite directions, an extension in drinking hours in England and Wales and a reduction in Spain. Focusing on one dimension of intra-marginal labor supply, absenteeism, we demonstrate a causal effect of these legislative changes. Increasing opening hours in England and Wales increased worker absenteeism, whilst reducing opening hours in Spain reduced absenteeism. This result appears robust to a range of standard threats to identification for the UK where we provide further evidence that suggests that the channel of transmission is through an increase in alcohol consumption. The effect for Spain, however, disappears after 2 years of implementation. This suggests that the 'benefits' from shorter opening hours in terms of less absenteeism are short-lived.

This result contributes to the growing literature on the labor market consequences of alcohol availability. It provides the first evidence, specifically, that more availability in terms of the opening hours of licensed premises has an effect on worker absence behavior. In addition, we contribute to existing evidence that availability influences alcohol consumption. Finally, our results suggest that government intervention in the regulation of leisure activities has the potential to have unintended consequences on worker behaviour and workplace productivity.

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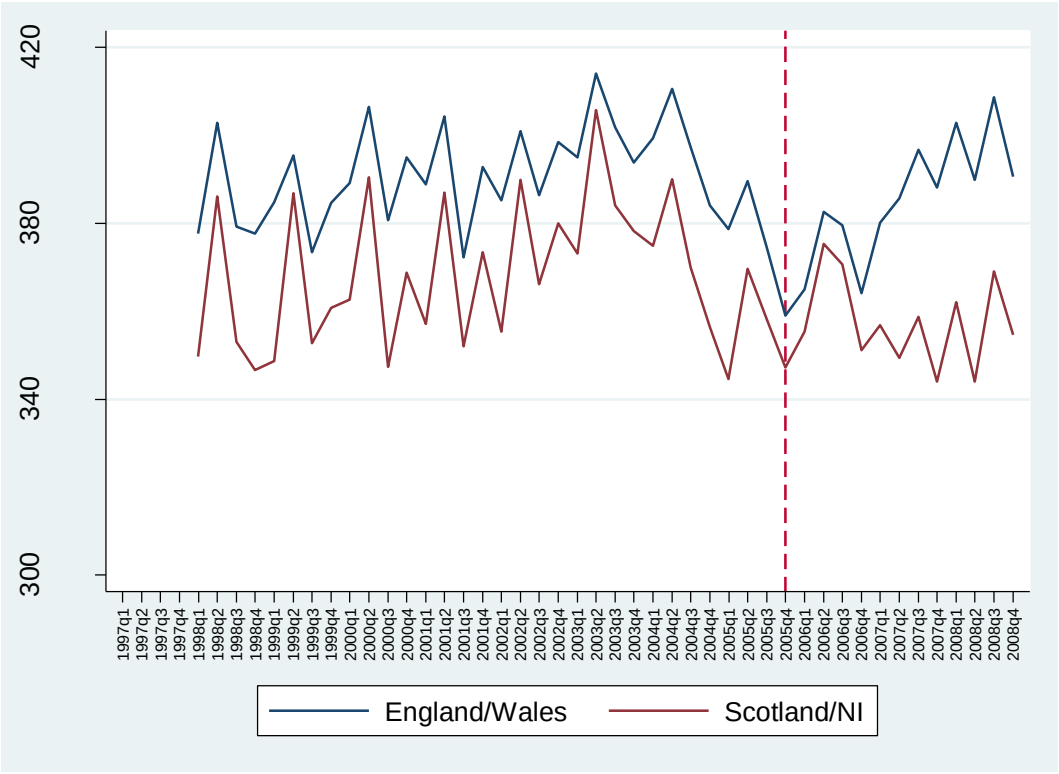
Figure 1

Panel A: Minutes of Absence for workers in the UK



Panel B: Absence rate for workers in the UK

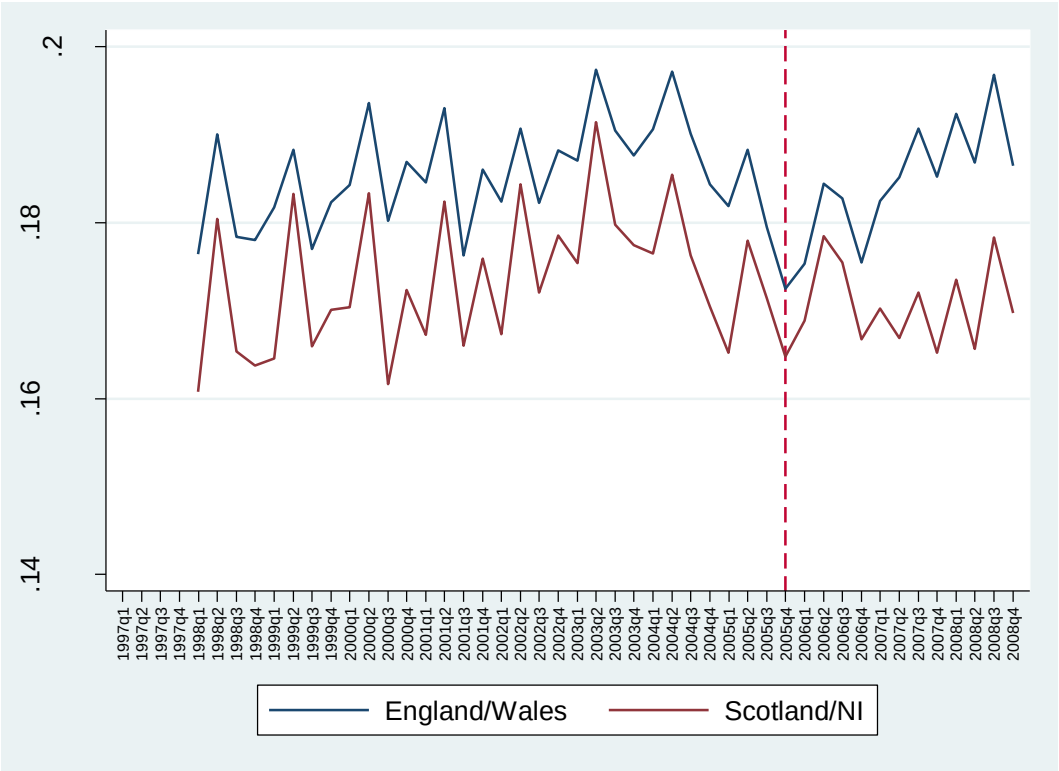
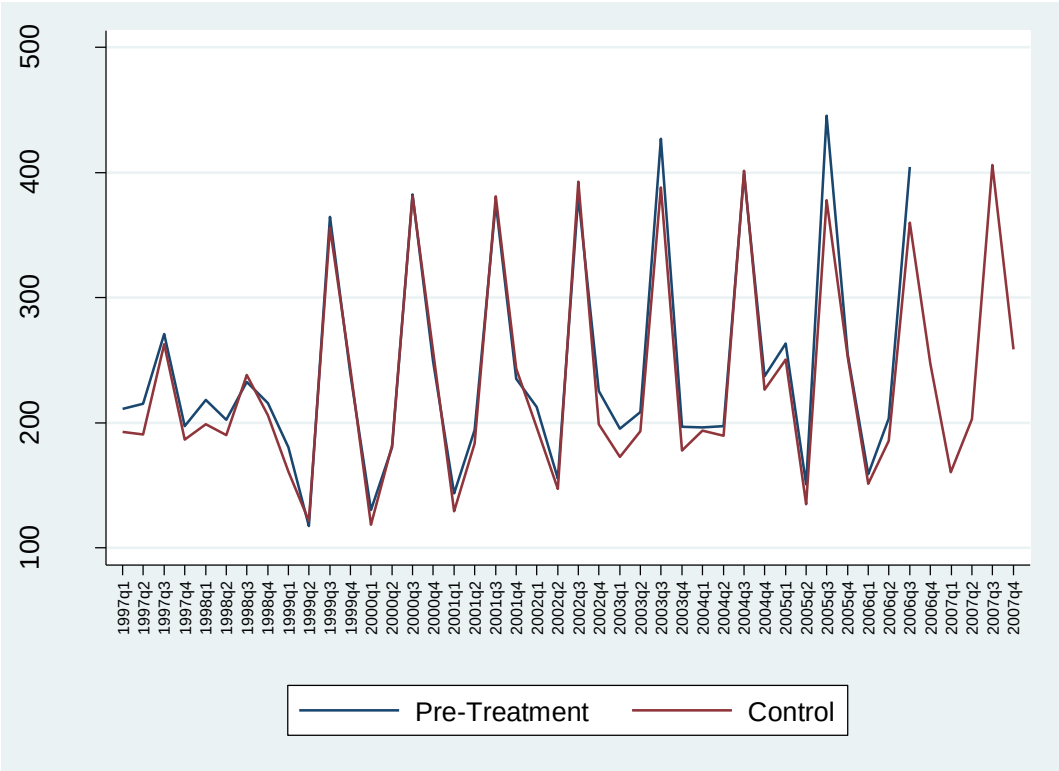


Figure 2

Panel A: Minutes of Absence for workers in Spain



Panel B: Absence rate for workers in Spain

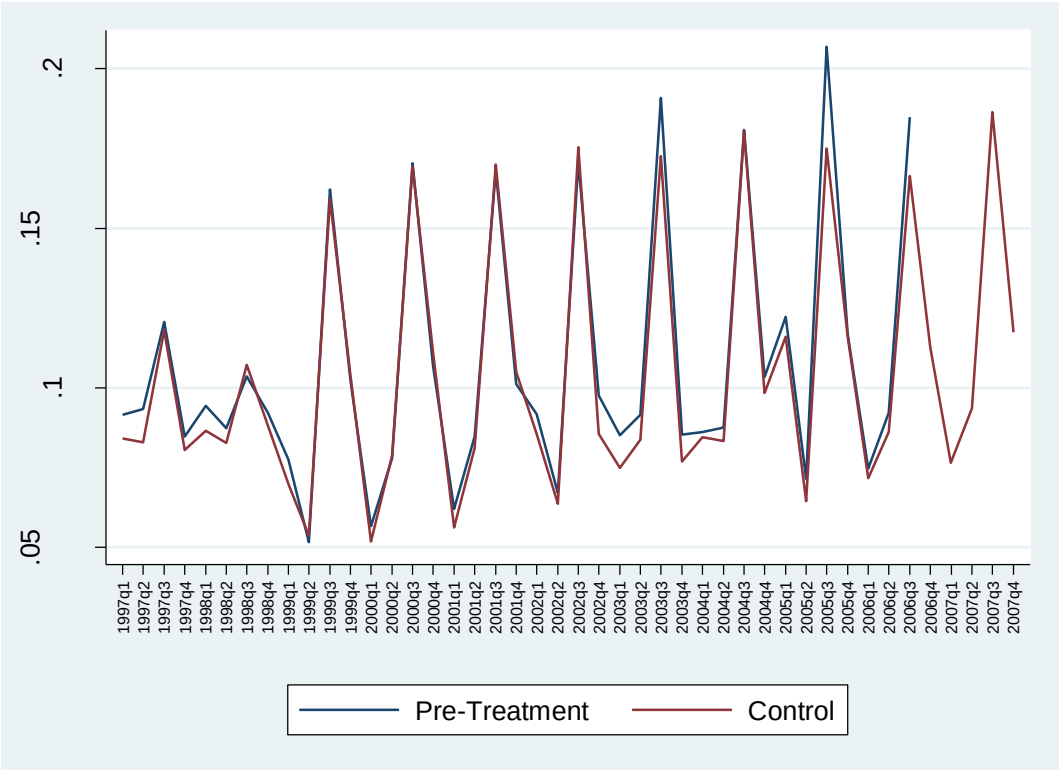
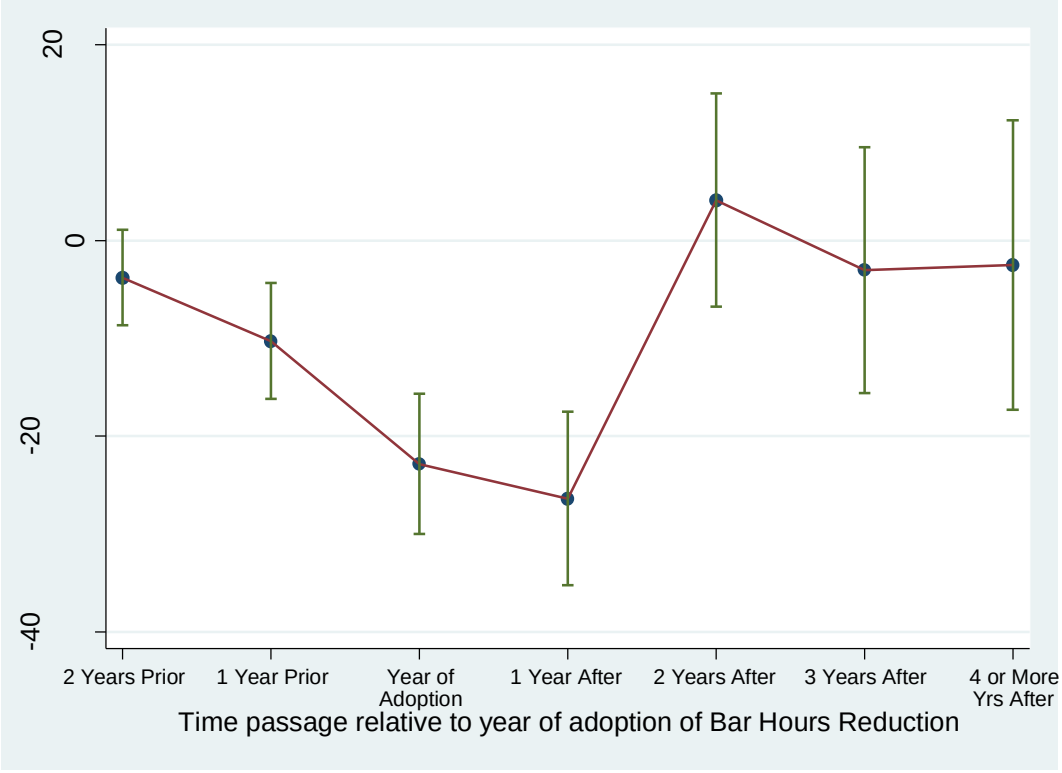


Figure 3. Estimated Impact of Reducing Opening Hours on Absenteeism for Years Before, During and After Adoption, Spain.



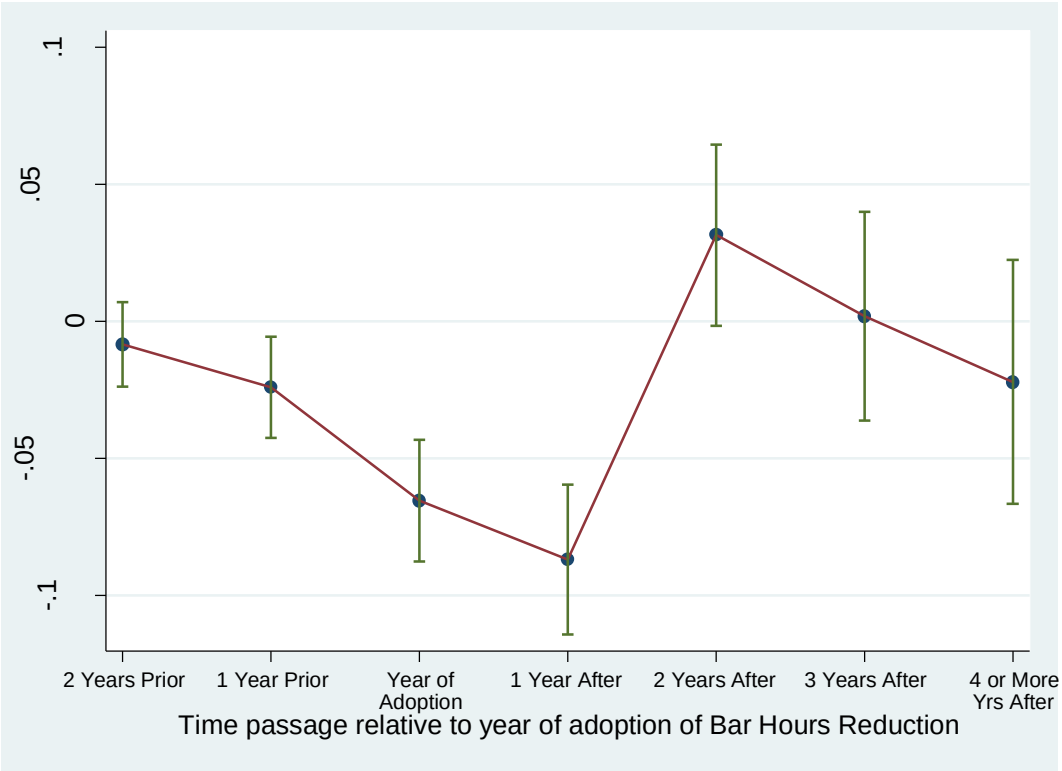
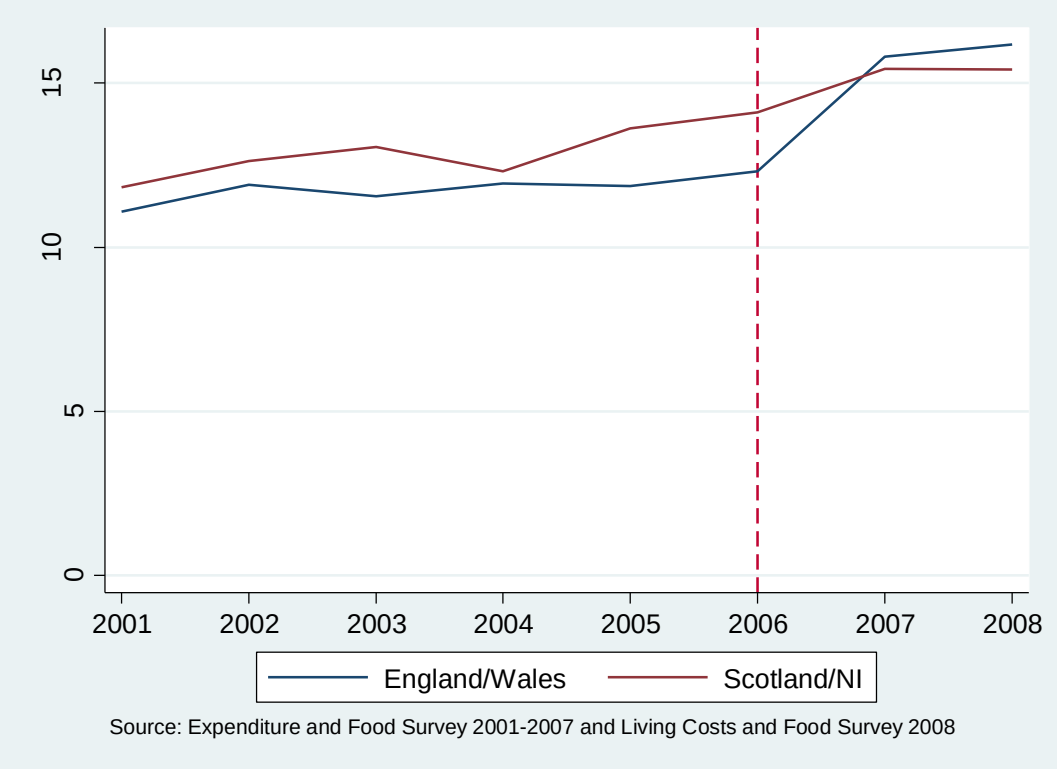


Figure 4. Total alcohol expenditure outside the household (pubs/restaurants/hotels) per week (£).



Source: Expenditure and Food Survey 2001-2007 and Living Costs and Food Survey 2008

Figure 5. Number of alcohol units drunk on heaviest day in the last 7 days (England)

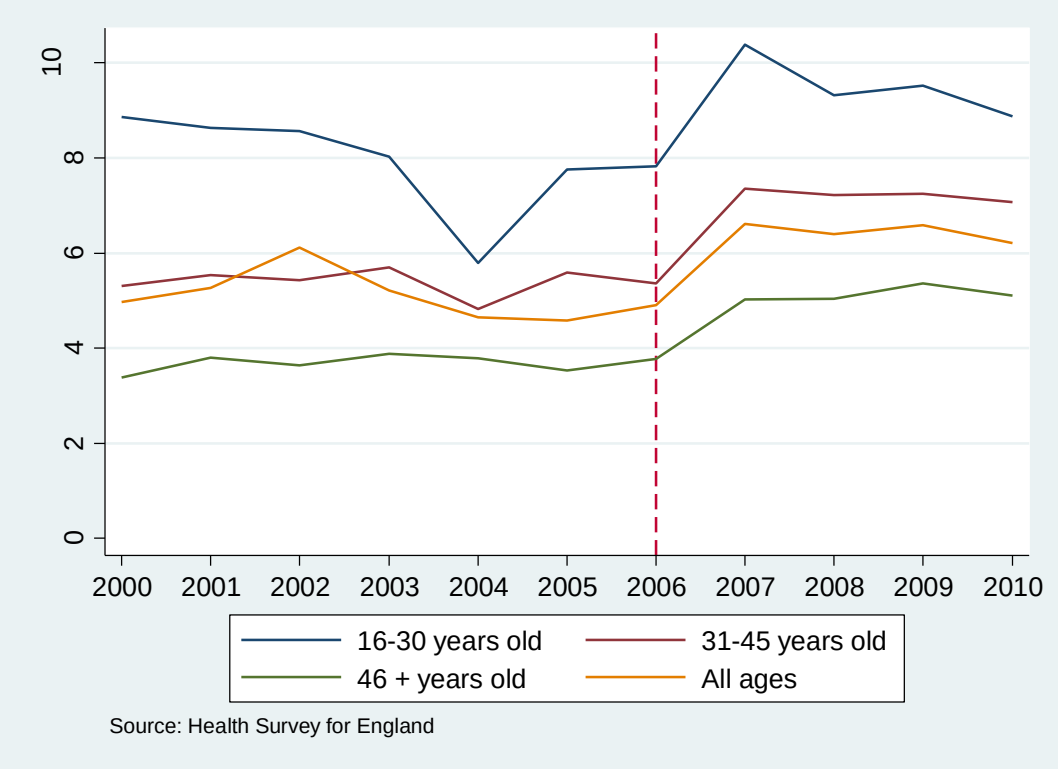


Table 1. Regional Timing of Drinking Hours Law Changes in Spain and England/Wales.

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 Absence Behavior in the UK, 1997-2008

	Panel 1			Panel 2		
	Minutes Absence	Absence Rate (Tobit)	Absence Rate (OLS)	Minutes Absence	Absence Rate (Tobit)	Absence Rate (OLS)
Policy	5.590	0.014	0.001	10.855	0.026	0.003
	(0.062)***	(0.000)***	(0.000)**	(0.199)**	(0.001)***	(0.000)**
Female	47.557	0.065	0.032	47.563	0.065	0.032
	(2.109)**	(0.006)***	(0.001)**	(2.118)**	(0.006)***	(0.001)**
Degree or higher	35.577	0.072	0.013	35.579	0.072	0.013
	(1.375)**	(0.007)***	(0.001)*	(1.378)**	(0.007)***	(0.001)*
Vocational training/Diploma	32.737	0.075	0.015	32.733	0.075	0.015
	(0.396)***	(0.006)***	(0.000)**	(0.400)***	(0.006)***	(0.000)**
A-Levels	25.320	0.052	0.011	25.328	0.052	0.011
	(0.652)**	(0.002)***	(0.000)**	(0.640)**	(0.002)***	(0.000)**
Temporary contract	-36.528	0.008	0.011	-36.530	0.008	0.011
	(2.814)**	(0.013)	(0.004)	(2.818)**	(0.013)	(0.004)
Public sector	65.960	0.097	0.035	65.960	0.097	0.035
	(0.781)***	(0.004)***	(0.000)***	(0.781)***	(0.004)***	(0.000)***
Region FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Trend * region	No	No	No	Yes	Yes	Yes
Observations	846106	846106	846106	846106	846106	846106

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. DD corresponds to the interaction between Treatment and Policy, that is, it takes value 1 for England/Wales after the policy was implemented. Time period: 1997-2008. Controls for marital status, presence of dependent children, industry, workers' occupation, year, and quarter are included but not reported. Wild cluster bootstrap-se at the regional level are reported in parentheses (400 reps).
*, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 3. Effect of licensing laws on worker absence behaviour in Spain, 1997-2007

VARIABLES	Panel 1			Panel 2		
	Minutes Absence	Absence Rate (Tobit)	Absence Rate (OLS)	Minutes Absence	Absence Rate (Tobit)	Absence Rate (OLS)
Policy	-3.424	-0.009	-0.001	-17.616**	-0.053**	-0.008**
	(4.631)	(0.017)	(0.002)	(7.502)	(0.021)	(0.003)
Female	35.089***	0.144***	0.025***	34.891***	0.144***	0.025***
	(3.751)	(0.011)	(0.002)	(3.507)	(0.010)	(0.002)
Secondary education	-1.977	-0.001	-0.001	-1.369	-0.001	-0.000
	(2.114)	(0.007)	(0.001)	(2.098)	(0.007)	(0.001)
Higher education	8.138***	0.027***	0.003***	8.572***	0.024***	0.003***
	(2.395)	(0.009)	(0.001)	(2.202)	(0.008)	(0.001)
Public sector	74.096***	0.178***	0.033***	75.900***	0.185***	0.034***
	(3.297)	(0.011)	(0.002)	(3.226)	(0.010)	(0.001)
Temporary contract	-52.586***	-0.127***	-0.022***	-51.758***	-0.129***	-0.021***
	(2.524)	(0.008)	(0.001)	(2.487)	(0.008)	(0.001)
Region FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Trend * region	No	No	No	Yes	Yes	Yes
Observations	1993260	1993260	1993260	1993260	1993260	1993260

Note: DD takes value 1 if region *j* at time *t* has shortened licensing hours. Controls for marital status, industry, workers' occupation, establishment size, region, year, and quarter are included but not reported. Wild cluster bootstrap-se at the regional level are reported in parentheses (400 reps). *, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 4. Age Treatment Heterogeneity

UK	18-30	31-45	46+
Minutes Absence	13.161	13.897	12.016
	(0.834)**	(0.465)**	(0.193)**
Observations	166353	359288	320465
AR Tobit	0.055	0.028	0.020
	(0.002)***	(0.002)***	(0.000)
Observations	166353	359288	320465

Spain	18-30	31-45	46+
Minutes Absence	-18.351**	-19.227**	-14.549*
	(7.073)	(8.353)	(8.022)
Observations	520413	823228	649619
AR Tobit	-0.059***	-0.054**	-0.047**
	(0.019)	(0.025)	(0.019)
Observations	520413	823228	649619

Controls for marital status, presence of dependent children, industry, workers' occupation, year, and quarter are included but not reported. Wild cluster bootstrap-se at the regional level are reported in parentheses (400 reps). *, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 5 Robustness Tests

	UK		Spain	
	Minutes Absence	AR Tobit	Minutes Absence	AR Tobit
Full time workers only	11.678 (0.356)**	0.033 (0.002)***	-17.973** (7.658)	-0.054*** (0.021)
Narrower Absence Definition	12.190*** (0.143)	0.049*** (0.003)	-17.544** (7.574)	-0.057*** (0.022)
Regions with high proportion licences (UK)	27.267*** (0.425)	0.048*** (0.004)		
Exclude Border Regions (UK)	11.769 *** (0.182)	0.027*** (0.001)		
Exclude year of policy implementation	10.855** (0.199)	0.026*** (0.001)	-9.986 (8.877)	-0.029 (0.029)
Early Adopters (Spain)			-53.261* (29.837)	-0.176* (0.101)

Controls for marital status, presence of dependent children, industry, workers' occupation, year, and quarter are included but not reported. Wild cluster bootstrap-se at the regional level are reported in parentheses (400 reps). *, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 6 Further Robustness Checks, UK

	Absence Rate	Minutes Absence
Regions with high proportion licences	0.048***	27.267***
	(0.004)	(0.425)
Exclude Border Regions	0.027***	11.769 ***
	(0.001)	(0.182)
Synthetic Cohort	0.008**	35.193**
	(0.004)	(16.095)\
Bertrand et al., (2004)	0.015***	14.407***
	(0.000)	(0.012)
Worker FE	0.013***	47.250***
	(0.000)	(0.313)
Placebo test Scotland/NI		
	0.011	5.190
FE	(0.023)	(48.071)

Controls for marital status, presence of dependent children, industry, workers' occupation, year, and quarter are included but not reported. Wild cluster bootstrap-se at the regional level are reported in parentheses (400 reps). *, **, and *** indicate statistical significance at the 10%, the 5%, and the 1% levels, respectively.

Table 7. Effect of Licensing Laws on Alcohol expenditure outside the household per week and the probability of going out to licensed venues to drink.

First-Stage	EFS: Log(alcexp)	BHPS: Drink often
DD	0.068**	0.012***
	(0.029)	(0.004)
Treatment	-0.096***	0.039***
	(0.019)	(0.006)
Policy	0.155***	-0.007**
	(0.027)	(0.003)
Observations	42523	139184

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

Table 8 Alcohol Expenditure and Absenteeism, Two Samples Two Stage Least Squares Estimates, UK.

SAMPLE 1 – UK LFS $A_{ijt} = \phi + \delta Policy_{it} + \gamma Treatment_{ij} + \beta Policy_{it} \times Treatment_{ij} + \alpha X_{ijt} + \tau Y_t + \varepsilon_{ijt}$		
	$\hat{\beta}$	$V(\hat{\beta})$
Minutes absence	15.5135	0.0415887
Absence rate	0.038589	0.00037587
SAMPLE 2 - EFS $\log(alc\exp)_{ijt} = \alpha + \delta Policy_{it} + \eta Treatment_{ij} + \gamma Policy_{it} \times Treatment_{ij} + \lambda X_{ijt} + v_{ijt}$		
	$\hat{\gamma}$	$V(\hat{\gamma})$
Log(alcexp)	0.06771	0.00084492
TS2SLS	$\hat{\rho}$	SE
Minutes absence	229.1171	98.40489425
Absence rate	0.569919	0.37662316

Standard Errors are calculated using the Delta Method $V(\hat{\rho}) \cong \frac{\hat{\beta}^2}{\hat{\gamma}^4} V(\hat{\gamma}) + \frac{1}{\hat{\gamma}^2} V(\hat{\beta})$

APPENDIX:

Table A1. Descriptive statistics

	UK			Spain	
	Mean	Std		Mean	Std
Minutes of absence	390.2009	721.8995	Minutes of absence	253.0919	628.9931
Absence rate	0.185599	0.331486	Absence rate	0.113595	0.277178
Age	41.09038	11.26567	Age	39.68434	11.67721
Female	0.505371	0.499972	Female	0.395422	0.488941
Married	0.644355	0.478708	Married	0.608853	0.488007
A-Levels	0.241515	0.428002	Primary education	0.493533	0.499958
Vocational training/Diploma	0.130792	0.337173	Second education	0.205629	0.404161
Degree or higher	0.181723	0.385616	Higher education	0.300838	0.458622
Public sector	0.306445	0.461017	Public sector	0.192055	0.393916
Temporary contract	0.054492	0.226986	Temporary contract	0.378426	0.484995
Part time job	0.25471	0.435699			
Dependent children	0.789453	1.114526			
	846106		Observations	1993260	