UNIVERSITAT DE BARCELONA

Essays in Health Economics

Grace Victoria Armijos Bravo

PhD in Economics

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1 Introduction

Health across the life cycle is an important component of well-being and quality of life. According to Deaton (2013), improvements in health and living standards have positively transformed our lives. Likewise, Deaton, in his book, "The Great Escape: health, wealth, and the origins of inequality" explains how good health contributes to economic growth and development. However, he also emphasizes that inequalities in health are a key element for gaping inequalities between people and nations.

The World Health Organization (WHO) defines health as "a state of complete physical, mental and social well-being and not merely the absence of disease or infirmity" (WHO, 1989). The WHO definition links health with well-being and as a resource for everyday life. Health is also linked to current and future socioeconomic outcomes that are part of an individual's well-being. For instance, health conditions at birth and during early childhood have been found to be a significant determinant of child and adulthood outcomes such as earnings, employment, educational achievement and future health (Lindeboom et al., 2010; Scholte et al., 2015; Deng and Lindeboom, 2022). Likewise, once in adulthood, current health also plays an important role in quality of life, labor market, and other socioeconomic outcomes.

This dissertation assesses how exogenous events can negatively affect health across different moments of life. Thus, the main purpose of this dissertation is to empirically assess whether external events have an impact on health. This thesis brings together four main empirical chapters that contribute to two research lines in the field of health economics. The first line, with two chapters, focuses on health at birth outcomes, and the second one explores changes in workers' health and how health shocks can affect women's labor market outcomes.

With respect to the first research line, there is some economic literature showing that good health at birth is linked to better outcomes later in life in terms of future health, wages, education attainment and human capital development. For example, low birth weight is associated with developmental difficulties (Almond et al., 2005), with worse educational achievement and earnings Black et al. (2007).

Many factors can influence health at birth. One of them, for example, is stress during pregnancy which, according to the medical literature, may have detrimental effects on the health of newborns (Beydoun and Saftlas, 2008; Torche, 2011). Economic literature assessing the effects of exogenous sources of stress (earthquakes, bombs, terrorist attacks) on health at birth outcomes is not very rich and has focused only on the general pregnant women population. The second chapter of this thesis focuses on assessing the effect of in-utero exposure to stress, generated by jihadist terrorist attacks, on the health of newborns of Muslim mothers. Another factor that can affect health at birth is mother's substance use/abuse. Chapter 3 explores how a public policy aimed at not criminalizing drug use may be linked to increased women's drug use and, through this, to bad newborn's health.

Bad health may affect individuals from before being born, through mother-fetus transmission. This refers to the so-called "fetal origins" hypothesis, in which Barker (1990), the author, argues that the intrauterine environment and nutrition "programs" the fetus to have specific metabolic characteristics, which can lead to future disease. The fetal origins literature suggests that in-utero health affects not only birth weight but also basic metabolism, which in turn affect future health outcomes. Thus, Barker (1990) argues that in-utero starving may contribute to higher risks of adulthood obesity, heart disease, and diabetes. Therefore, there is a transmission channel from in utero health to adult health, being the latter strongly associated with adult (socioeconomic) well-being.

The economic literature has also contributed with evidence that supports the fetal origins hypothesis. For instance, Currie and Vogl (2013) studied the link between fetal health and long-term outcomes for developing countries, and conclude that health in-utero is an important determinant in developing economies. Lindeboom et al. (2010) study how nutritional conditions in utero may affect health and mortality later in life. The authors find strong evidence for negative long-run effects caused by exposure to famines in the Netherlands. Likewise, Scholte et al. (2015) using data from the Dutch famine in year 1944/1945, find a significantly negative effect of in utero exposure to this famine on employment and hospitalization rates.

With regards to the link between child health and future socioeconomic outcomes, Currie (2009) finds evidence supporting that health could play a role in the intergenerational transmission of economic status. On this, the author explains that poor child health is likely to affect future health, which in turn may affect individual's labor market results. In addition, the author also highlights that child health may negatively affect the capacity of acquiring skills. Moreover, poor health is also strongly associated with lower rates of school attendance. Finally, poor nutrition during childhood may not only negatively affect education achievements in the short and long term, but also may contribute to worse later-life health Deng and Lindeboom (2022).

Moving to the second research line of this thesis, it explores the interactions between health and labor market outcomes. Chapter 4 investigates how the call for new civil servant positions may affect workers health, which we measure through sickness absences. In the economic literature, research on the effects of public examinations on workers socioeconomic outcomes such as health and well-being is very scarce. Most of the research is focused on the area of psychological and personnel management from a "best practice" perspective. During adulthood individuals may be exposed to work-related events that can affect their health. For example, Robone et al. (2011) find that working conditions are associated with mental health, with different results for men and women. Likewise, Niedhammer et al. (1998); Rugulies et al. (2006) study how high psychological demands relates to psychological well-being in the context of labor markets. They find that there is a positive association, which translates into an increase in the probability of experiencing lower emotional satisfaction. Likewise, Robone et al. (2011) find that job insecurity increases the probability of psychological distress. In this context, this chapter evaluates whether public examinations are generating negative externalities in terms of increased sickness absences.

Finally, Chapter 5 evaluates how external health shocks may influence individual's labor market outcomes, with a focus on women's employment. More precisely, this chapter focuses on the Covid-19 pandemic to evaluate whether non-pharmaceutical interventions aimed at reducing social contact, are linked to negative labor market outcomes for women workers.

This dissertation also pays special attention to the role that women play in a set of socioeconomic outcomes. First, in the intergenerational transmission of health from mothers to children, which is explored in Chapters 2 and 3. Second, in how women can be affected differently than men in the labor market in view of the entry exams for civil servants. In addition, this thesis also explores whether women are more affected than men in terms of job loss as a consequence of the Covid-19 pandemic, topics that are explored in Chapters 4 and 5.

Looking more in detail each of the empirical chapters, Chapter 2, with the title *"Terrorist Attacks, Islamophobia and newborns' health"*,[1] exploits the exogenous shock of the terrorist attack occurred in Catalonia-Spain in 2017 to study the impact of terrorist attacks on newborns' health. These attacks were perpetrated by a jihadist group, which contributed to the increase in Islamophobia against the Muslim population living in the target cities, which is problematic for pregnant women since stress is negatively associated with newborn's health Torche (2011); Ivandic et al. (2019).

This chapter contributes to the literature on, (i) maternal stress and birth outcomes, and (ii) discrimination and health. It also contributes to the discussion on

[1] This chapter is based on: Armijos Bravo and Vall Castelló (2021). Terrorist Attacks, Islamophobia and newborns' health. Journal of Health Economics, 79:102510.

the negative consequences that jihadist attacks may have on the Muslim community.

This chapter uses several databases. First, administrative data on live births in Catalonia from years 2015 to 2018 to estimate the effect of the attacks on newborns to Muslim mothers living in the target cities. Using a difference in differences in differences approach, the findings show detrimental results is terms of the share of low birth weight newborns and in the share of deliveries with complications. The effect is also different across the trimester of gestation. Additional results are provided to support the robustness of the baseline empirical method, for instance an event study and several placebo tests.

In the last part of the chapter, potential mechanisms are explored to support an increase in Islamophobia and stress-related outcomes in Muslim women. For this, this chapter employs survey and administrative data on rejection against the Muslim community, hate crimes and stress-related outcomes. The results point to an increase in Islamophobia and in emotional distress after the attacks. The conclusion is that Islamophobia-related stress is possibly one of the channels affecting health at birth. The results have important policy implications as newborns' health matters for future socio-economic outcomes such as adult health, earnings and quality of life, and increased discrimination may reinforce the differences already existing between Muslims and non-Muslim.

In the third chapter of the thesis, *"Drug Tenancy, Thresholds, Consumption and Newborn's Health"*, I study how a policy aimed at not criminalizing illicit drug consumers, to recognize their health care needs, might negatively influence newborn's health in Ecuador. In 2013, a policy that established maximum permissible amounts of possession of illicit substances for personal consumption was implemented in Ecuador. This regulatory change might be one of the drivers of increased availability of drugs, which in turn, may increase consumption in the population of reproductive age, and through this, affect newborn's health. I exploit variation in drug exposure at the province level in a difference-in-differences framework, which compares Ecuadorian provinces with different intensities of drug exposure, before and after 2013.

To measure drug intensity exposure before the policy change, I develop a composite index at the province level using administrative and survey data. I use this index to estimate changes in newborn's and women's health. With respect to newborn's health, this chapter uses administrative data from live birth registries and from neonatal hospital discharges. I find that the share of low birth weight and very low birth weight newborns raise significantly in provinces initially more exposed to drugs. In addition, there is an increase in neonatal hospitalizations to medical conditions related to maternal drug use/abuse.

Moving to women's health outcomes, using administrative data on hospital discharges, the findings support an increase in hospitalizations to drug-consumption related diagnostics. There is also evidence that supports an increase in the deliveries with complications. Overall, I document a significant increase in both women and men's substance use. I also report a set of robustness checks to support the preliminary findings.

While the results of this chapter are still preliminary, the findings seem to suggest that the increase in drug use, through increased availability, is possibly one of the channels affecting newborn's health living in provinces more exposed to drugs.

In the fourth chapter of the dissertation, entitled *"Job Competition in Civil Servant Public Examinations and Sick Leave Behavior"*,[2] we investigate the impact of the call for new civil servant positions on workers sickness absences. In several countries the entry system to access public service positions is the traditional public examination procedure. In this setting, candidates have to take passing exams that require a huge load of material to study, and therefore time. Candidates working while preparing for public exams may find it difficult to spend enough time on both tasks. Thus, they might experience increased stress/anxiety related to public service entrance exams.

This chapter uses novel administrative dataset on the universe of sickness absences and civil servant positions offered in Spain from 2009 to 2015. The findings point to a significant increase in health-related work absences several months before the examination date. This effect is stronger for individuals working in the education sector as well as for calls offering a large number of positions. This effect is mostly driven by stress related absences. In a second step, using data on medical visits (GP and specialist), we find evidence consistent with a deterioration in public sector workers' health. The results are important from a policy perspective as they highlight the existence of important negative consequences of the civil service recruitment process that have been previously overlooked.

The last empirical chapter of the thesis, *"Covid-19 Lockdown in Ecuador: Are there Gender Differences in Unemployment?"*,[3] examines whether lockdown policies during the first wave of the Covid-19 pandemic differently hit formal employment for women as compared to those of men in the Ecuadorian labor market. To fight Covid-19, governments imposed restrictions on personal mobility and social interactions which may have had negative consequences in the labor market. These consequences may be different across demographic groups, particularly for female workers.

[2]This chapter is based on: Armijos Bravo, G. and Vall Castelló, J. (2023). Job competition in civil servant public examinations and sick leave behavior. Instituto de Economía de Barcelona (IEB), Working Paper No. 2023/04.

[3]Paper coauthored with Segundo Camino Mogro.

This chapter employs a difference-in-differences-in-differences model to compare female employees working in restricted economic sectors with other workers, before and after the lockdown policy. We make use of unique administrative data on the universe of unemployment spells registered in the social security system of Ecuador during years 2019 and 2020. We complement this analysis using survey data to explore associations between the lockdown and the probability of being employed across different age and education level groups. The results show that the lockdown policy is associated with an increase in the number of unemployment spells of women working in the restricted economic activities, during the first wave of the Covid-19. We also document a decrease in the probability of being employed, which is particularly strong for the youngest women (15–24- years-old), oldest women (45-65 years-old), and less educated female workers. We conclude that the lockdown policy imposed in Ecuador is a plausible explanation for women's job loss in the formal sector.

In sum, this dissertation contributes to the literature studying the impact that exogenous events may have on health, with a particular focus on newborn's and worker's health. It also adds on the relation between the Covid-19 health shock and socioeconomic outcomes such as employment. In addition, to this introductory chapter, this dissertation includes four more empirical chapters, each of them contributing to specific strands of the health economics literature, a concluding chapter (Chapter 6) and a reference list.

Finally, in the conclusions, this thesis highlights the main findings of each chapter and the importance of evaluating public policies and events that may affect health across the life cycle.

2 Terrorist Attacks, Islamophobia and Newborns' Health

2.1 Introduction

During the last years, several terrorist attacks have been executed across western countries. For instance, those perpetrated by jihadist groups such as the 9/11 attacks, the bombing at Madrid's biggest train station in 2004, the London bombing of July 2005, the shooting in Paris in January 2015 and November 2015, and others perpetrated in Brussels, Nice and Berlin. In Spain, the last one was in Barcelona – Cambrils (August 2017) with 16 dead and a hundred of injured.

Terrorist attacks are events that impose considerable individual and societal costs (Krueger, 2007). With respect to jihadist attacks, they have generated an increase in the rejection against the Muslim community across countries, which in turn, may generate an increase in Islamophobia, especially in the affected cities after a jihadist attack (Elver, 2012; Gould and Klor, 2016; Swahn et al., 2003; Kaplan, 2006).

Islamophobia can be experienced by the Muslim community in many aspects of daily life in terms of feelings of rejection, harassment, fear, hate speech, among others (Union of Muslim Communities in Spain, 2018; Amer and Bagasra, 2013). According to the European Union Agency for Fundamental Rights (2018), as reported in its Second Survey of the European Union on Minorities and Discrimination for Muslims, 39% of the sample felt discriminated because of their origin. Focusing only on women, we see that 35% felt discriminated, which represents an increase of 11 percentage points compared to the same survey in 2008. In addition, 31% of them experienced harassment, 39% have been subjected to inappropriate looks or offensive gestures in the 12 months prior to the survey, and 22% have been subject to insults because of their immigrant and Muslim condition. In addition, there is also some evidence of an increase in hate crimes against Muslims after a jihadist attack (Ivandic et al., 2019; Gould and Klor, 2016).

This discrimination and stigma may be a source of stress for the immigrant Muslim community living in western countries. This stress is particularly problematic for pregnant Muslim women, even if they do not directly experience the discrimination, because the concern about what might happen would be a considerable source of stress as a post-terrorism-event psychological consequence[1] (DiMaggio and Galea, 2006; Hobfoll et al., 2006; Becker and Rubinstein, 2011; Bader and Berg, 2013). There is vast (medical) literature showing that intrauterine exposure to stress has a negative influence on health at birth. For instance, it increases probabilities of low birth weight[2] , and complications during pregnancy and delivery (Rondó et al., 2003; Beydoun and Saftlas, 2008; Torche, 2011; Khashan et al., 2008). There is also some economic literature on this regards; for example, Black et al. (2016) find that stress leads to worst health at birth outcomes in Norway, and Persson and Rossin-Slater (2018) find that stress during pregnancy depresses birth outcomes and raises the risk of perinatal complications.

The aim of this chapter, therefore, is to identify the impact of the Catalonia 2017 jihadist terrorist attack on health at birth for newborns whose mothers are from Muslim countries living in affected cities. To do so, we exploit the exogenous source of stress coming from the attacks occurred in Catalonia in August 2017, as the exact location, intensity, and timing were largely unpredicted by the general population. We use a difference-in-differences-in-differences strategy to estimate the effect of the attacks on health at birth comparing differences across mother's country of origin, Muslims versus non Muslims (first difference), and mother's municipality of residence, affected versus non-affected cities (second difference), before and after the attacks.

Health conditions at birth, such as birth weight, have been found to be a significant determinant of child and adulthood outcomes in terms of health, education and earnings. For example, Almond et al. (2005), using data for the U.S., found that lower birth weight newborns experience developmental difficulties. Black et al. (2007) found that lower birth weight babies have worse outcomes in terms of educational attainment and adult earnings in Norway. Currie and Vogl (2013) studied the links between fetal health and long-term outcomes for developing countries, and they conclude that health in early life is a more significant determinant in these types of economies. Likewise, Almond and Currie (2011) made an extensive review on the effects of birth weight on socio-economic adult outcomes.

Research assessing the causal link between terrorist attacks and health at birth is scarce. For instance, Camacho (2008) using data for Colombia, assessed the impact of exposure to landmines, as a source of exogenous stress, on birth weight. She found that exposure in the first three months of pregnancy decreases birth weight

[1] In addition, terrorism may cause devastating consequences in terms of lost human lives and lost infrastructure (Krueger, 2007; Bloomberg et al., 2004; Lenain et al., 2002).

[2] According to the World Health Organization (WHO), low birth weight is defined as weighting less than 2500 g at birth.

by 8.7 g. Despite this result, it should be mentioned that landmines in Colombia were not the primary source of stress for the general population as were car bombs and massacres, as pointed out by Mansour and Rees (2012). Mansour and Rees (2012) estimated the effect of intrauterine exposure to armed conflict on pregnancy outcomes using data on the al-Aqsa Intifada.[3] They found that stress due to the conflict decreases birth weight and increases the probability of having a low birth weight baby. However, as recognized by the authors, other channels in the Palestinian conflict context may be affecting birth weight, such as malnutrition and lack of access to medical facilities, due to mobility restrictions in the region. Quintana-Domeque and Ródenas-Serrano (2017) studied the effects of intrauterine exposure to terrorism in Spain. Focusing on the ETA[4] attacks, they found that exposure to terrorism in the first months of pregnancy has detrimental effects on average birth weight and the fraction of normal babies. Finally, there are other papers that exploit the attacks occurred in September 2001 in New York City and document a negative association between the attacks and health at birth (Berkowitz et al., 2003; Eskenazi et al., 2007; Smits et al., 2006; Brown, 2014).

This chapter contributes in several ways to two strands of the literature, (i) maternal stress and birth outcomes, and (ii) discrimination and health. Thus, to our knowledge, this is the first attempt to provide causal evidence on the effects of jihadist terrorism on health at birth outcomes. As described before, previous literature has focused on the impact of terrorism on newborns' health as a whole. We rather exploit a new channel through which Muslim pregnant women may be affected, that is, the increased Islamophobia that may be a source of stress. As emphasized by Sheridan and Gillett (2005) as part of their research on discrimination, a terrorist attack may influence perceptions that other social groups have about the social group to which the perpetrators belong.

We also contribute to the discussion on the negative consequences that jihadist attacks may have on the Muslim community. Several papers have found an increase in hate crimes right after a terrorist attack perpetrated by jihadist groups. For instance, Ivandic et al. (2019) document a causal link between local anti-Muslim hate crimes and jihadist terror attacks in the United Kingdom. Sheridan and Gillett (2005) and Gould and Klor (2016) find the same results for the New York 9/11 attacks. Thus, we provide new evidence on increased Islamophobia as a mechanism behind detrimental health outcomes at birth. In addition, we explore whether Muslim women living in target cities may be affected by increased levels of stress right

[3]Al-Aqsa Intifada is the name given to describe the armed conflict between Palestinians and Israel between years 2000 and 2004.

[4]ETA is the "Euskadi ta Askatasuna" or Basque Homeland and Freedom, terrorist group in Spain.

after a jihadist terrorist attack.

To do so, we take advantage of the attacks as an exogenous surge in discrimination that can be isolated from Muslim mothers' characteristics, and add evidence on the negative effects that discrimination may have on health. Research on this issue is scarce in the economics literature. For instance, Johnston and Lordan (2012) estimate the causal impact of increased Anti-Muslim discrimination on a range of objective and subjective health outcomes for the UK and find that discrimination is linked to worse health outcomes such as blood pressure, cholesterol, body mass index and self-assessed health. Likewise, Maas and Lu (2021) evaluate whether partisan politic discrimination affects health. Their results show that mortality rates for counties that vote for the loser candidate increase after elections. The main mechanism behind this result is increased feelings of social disintegration (stress, anxiety and isolation). In the same line, Samnaliev (2014) examines the link between discrimination, victimization, and social standing on health-related quality of life (HRQL) for sexual orientation minorities, and find that the affected group shows lower HRQL. Additional research has been conducted in the public health and epidemiology literature and, most of these studies, find that discrimination is associated with worse health (Lauderdale, 2006; Williams et al., 2003).

As an additional contribution, the analysis in this chapter is implemented at the local level using data from municipalities, rather than at a more aggregate level (provinces or regions) as is usually done in the literature on terrorist attacks and health at birth.[5] In this way, we can take into account that not everyone is equally affected within a province.

Finally, we exploit the exogenous and unpredictable elements of the timing, intensity, and location of the Catalonia 2017 attacks, and apply an identification strategy, triple differences, that allows us to control for two kinds of potentially confounding trends: changes affecting health at birth outcomes of Muslim newborns, unrelated to the attacks, and changes affecting health at birth outcomes of all Muslim and non-Muslims babies living in the treated cities (cities where the attacks where perpetrated).

As a first insight of our results, we find that newborns from immigrant Muslims mothers living in the affected cities have worse health at birth outcomes. In terms of birth weight, we find a higher probability of being born with low weight. Regarding deliveries, there is an increase in the share of complications. Our baseline estimates show that the Catalonia 2017 attack is translated into an increase in the probability of having a low birth weight newborn by 23.77%, and an increase in the share of complications by 13.02%. In addition, we find that being exposed (while in

[5]See for example Quintana-Domeque and Ródenas-Serrano (2017).

utero) to the attacks during the first trimester of gestation is linked to detrimental health at birth outcomes, although we also find negative effects on the second and third trimesters. Our results are robust to a battery of robustness checks such as a placebo treatment group, indirect treatment effects (using distance to affected cities) and city-specific effects. Moreover, we provide some evidence on the increase of Islamophobia in the affected zones, relative to the rest of the territory. We find an increase in the rejection against Muslims of 5.46 percentage points and an increase in reported hate crimes after the attacks. Additionally, we find evidence supporting increased levels of stress as well as stress-induced behavioral responses associated to the attacks.

The results have important policy implications regarding the situation of the immigrant Muslim population living in non-Muslim countries. As mentioned above, newborn's health matters for future socio-economic outcomes such as adult health, earnings and quality of life. Increased discrimination may reinforce the differences already existing between Muslims and non-Muslim. For example, Muslims living in Spain have, on average, lower levels of education than natives, are less likely to find a job due to their religion, and earn lower wages compared to Spaniards with the same degree and in a similar employment position (European Union Agency for Fundamental Rights, 2018). Thus, our results should raise awareness of the potential negative impacts for the Muslim population and should encourage authorities to implement and work towards more anti-discrimination policies that will foster a more inclusive society. Finally, our findings may be an useful tool across countries in which there is a high percentage of Muslim individuals, and have been also affected by terrorist attacks.

The rest of the chapter proceeds as follows. Section 2.2 provides the context of the Catalonia attacks and rejection against Muslims. Section 2.3 describes the data sets and shows some summary statistics. Section 2.4 presents the empirical strategy. Section 2.5 presents the results. In Section 2.6, we present a battery of robustness checks and, in Section 2.7, we assess possible mechanisms. Finally, Section 2.8 concludes with a discussion of the results.

2.2 Background

2.2.1 The Catalonia Attacks

Catalonia is one of the 19 Regions "Comunidades Autónomas" of Spain. It is made up of four provinces: Barcelona, Girona, Lleida and Tarragona. In August 2017 a group of individuals, using vehicles and knives, executed a terrorist attack in the cities of Barcelona and Cambrils, located in the provinces of Barcelona and Tarragona, respectively (Reinares and García-Calvo, 2017). A jihadist organization assumed the responsibility of these attacks, where, as described by Reinares and García-Calvo (2018), the perpetrators were a group of nine second-generation Muslim youth who were raised and residents of the city of Ripoll, located in the province of Girona.

These attacks took place, first in the city of Barcelona, during the afternoon of August the 17th 2017, when a vehicle was introduced in the pedestrian zone of the popular "Las Ramblas" hitting all the people walking in the area. As a result of this act, thirteen people were killed and more than a hundred injured. Some hours after this attack, the same mechanism was used in a pedestrian walkway in Cambrils, Tarragona. Together, the Catalonia attacks left sixteen dead, without including the jihadist terrorists killed, and around 140 injured (Reinares and García-Calvo, 2018).

Regarding the terrorist attacks' history in Spain, the last attacks before the Catalonia ones, were in March 2004 in the main train station of the capital, Madrid. Terrorist attacks are unexpected negative events affecting both individuals and society and, therefore, an exogenous source of post traumatic stress and fear among people. Several authors have found that, in addition to the direct consequences of terrorism, such as lost infrastructure and deaths, there are also psychological effects (Becker and Rubinstein, 2011; Schuster et al., 2001; Galea et al., 2003, 2002; Gidron, 2002).

As other European cities had been the target of terrorist attacks before the Catalan tragedy, one may be worried that it could have been easy to predict that additional attacks would follow. However, the exact location, timing, and severity of the attacks were very difficult to predict by the general population living in Spain. This unpredictable character of terrorist attacks allows us to draw reliable conclusions supporting our empirical strategy (see for example: Jaeger and Paserman (2008), Camacho (2008), and Mansour and Rees (2012) for statements supporting exogeneity in terrorist attacks.).

2.2.2 Terrorist Attacks, Islamophobia, Stress and Health at Birth

As mentioned above, terrorist attacks are violent events that have negative consequences both for society and individuals. When these events are perpetrated by one specific group, the rest of the society may change their perception towards individuals belonging to the same group as the perpetrators (Sheridan and Gillett, 2005). In this sense, the increase in jihadist terrorist attacks executed around the world in the last years have raised rejection levels against the Muslim community, also known as Islamophobia. According to Union of Muslim Communities in Spain (2018), Islamophobia is defined as an unfounded hostility to Islam and, consequently, a feeling of fear and aversion towards all Muslims (or most of them). Islamophobia is a phenomenon that may be experienced in several dimensions, from psychological to physical threats. It comprises expressions of prejudice, stigma, discrimination, critiques of the Islamic religion, hate speech and hate crimes against Muslims (Tartaglia et al., 2019).

Several authors have studied how, after terrorist attacks perpetrated in the western world, Islamophobia has experienced a dramatic rise. For instance, regarding the 9/11 New York and Washington attacks, Gould and Klor (2016) found that Muslim immigrants living in the affected cities experienced more exposure to hate crimes. In turn, those groups exposed to higher hate crime rates also exhibited lower English proficiency, greater chances of marrying another Muslim and lower female labor force participation rates. Therefore, the authors conclude that the 9/11 attacks generated a reaction against Muslims which, in turn, slowed their rate of assimilation. Swahn et al. (2003) using data from hate related attacks against the Muslim population after the 9/11, found that most of these crimes occurred in the 10 days immediately after the terrorist attacks. In addition, victims were mostly men that experienced Islamophobia in public places such as supermarkets, stores and schools. Sheridan (2006) investigated the levels of self-reported religious discrimination in British Muslims after the 9/11 attacks. She found that indirect discrimination increased by 82% and direct manifestations of Islamophobia increased by 76%. With respect to attacks occurred in Europe, a more recent paper by Ivandic et al. (2019), using data from the Greater Manchester Police, found an increase in Islamophobic hate crimes right after a jihadist terror attack in Europe.

Islamophobia, with its discrimination and stigma, may be a source of stress that can negatively affect health (Samari, 2016). In fact, several studies have found a strong relation between Islamophobia and increased levels of stress (Samari et al., 2018; Haque et al., 2019). This exposure to stress is particularly problematic for pregnant Muslim women who, in addition, are more easily identified as Muslims.

Intrauterine exposure to stress has been linked to worse health outcomes at birth such as birth weight and complications. This, in turn, may be a source of disparities in socio-economic outcomes later in life (in adulthood).

2.3 Data and Variable Definitions

2.3.1 Main Sources

Regarding **birth records in Catalonia**, we use administrative data provided by the National Institute of Statistics[6] (INE) from 2015 to 2018. In this data set the unit of observation is a live birth. We construct the sample according to the mother's province of residence; that is, we consider all entries that report that a mother's province of residence is one of the four provinces in Catalonia.

For every unit of observation, we have information on weight (in grams), month and year of birth, gender, whether the delivery was by c-section, complications during delivery and gestational length (in weeks). With this information, we construct our outcomes of interest: birth weight (in grams), an indicator of low birth weight (LBW) that takes the value of one for weight below 2500 grams, an indicator of pre-term delivery when the gestational length is less than 37 weeks, an indicator variable for complications during delivery, an indicator of the sex of the newborn (1 for females and zero otherwise), and an indicator of deliveries through a c-section. We also have information regarding the mother of the child: her country of birth, municipality and province of residence, age, and level of education.[7] Unfortunately, the administrative records do not have any information regarding mothers' risky behaviours, such as alcohol consumption, smoking or prenatal health care use.

To identify the trimester of pregnancy in which mothers were exposed to the attacks, and therefore, the in-utero exposure to stress, we use gestational age (in weeks). This measure is estimated (by healthcare professionals) as the period between the first date of the last menstrual cycle (which is also considered the conception date) and the date of birth. In this way, the accuracy of the gestational age is crucial for our estimates of trimester of exposure. However, we are not concerned about the reliability of this variable, as Spain uses international guidelines to estimate the conception date.[8] Thus, we use the reported gestational age and the date of birth to calculate the estimated conception date and, then, the trimester of gestation

[6]INE, www.ine.es

[7]Mothers' level of education is only available for some years In fact, for more than 50% of the observations the value is missing.

[8]In addition, in Spain all pregnant women have free access to prenatal care and to professional care during childbirth. For the years of analysis, 99.38% of all births took place in hospitals.

relative to the date of the attacks. Following Álvarez-Aranda et al. (2020), the first trimester is calculated from the estimated conception date until week 13; the second trimester ranges from weeks 14 to 26 of gestation; and the third trimester goes from week 27 until birth.

To identify whether a mother was exposed to the attacks during the first, second or third trimester of pregnancy, we construct a categorical variable with four groups. The first category includes mothers who were exposed to the attacks during the first trimester of pregnancy, the second one consists of mothers exposed in the second trimester, the third group includes mothers exposed during the third trimester of gestation, and the fourth group includes mothers that were not exposed. As one may be concerned about fertility responses, we limit the sample to births that were conceived before the attacks in all our models.

To assess possible mechanisms, we make use of four data sets. First, we employ data from the survey on **"Attitudes toward immigration"**, waves VIII, IX, and X from the "Centro de Investigaciones Sociológicas" of Spain. We use this data in order to examine changes in the rejection against Muslims living in affected cities after the 2017 Catalonia attacks. In these surveys, respondents (only Spaniards) are asked about their feelings towards all immigrants as well as towards specific groups of the immigrant population, such as Muslims. Other general questions are also asked. Wave VIII corresponds to the year 2015, IX to 2016 and wave X to September 2017,[9] which is a great fit for the time horizon of our analysis. This survey is nationally representative and samples all 52 Spanish provinces, including municipalities with more than 100,000 inhabitants.

Second, we use administrative data on **Reported hate crimes**[10] provided by the Catalan Police, "Mossos d'Esquadra", for years 2015 to 2017.[11] We have information on the total number of hate crimes reported in the Region of Catalonia at the municipality level and for every month for each of the three years included (2015, 2016 and 2017). To be used as a placebo test, we also use administrative records on falsification crimes; a type of crime that includes documents forgery, identity usurpation and other related crimes, across Catalan police areas. As a final data source on hate crimes, we use administrative records provided by the "Ministerio del Interior", in which we are able to identify the nationality of the victim. This dataset is provided at the province level for years 2015 to 2018.

To analyze the relation between the attacks and reported stress among women

[9]For this survey, the last year available is 2017.

[10]According to the "Mossos d'Esquadra" hate crimes are define as *"criminal offenses where victims are selected based on his or her race, national origin, ethnicity, religion, language, sexual orientation or other similar factors"*.

[11]To date, 2017 is the last year available for external use.

living in affected cities, we use data from the **Catalan Health Survey (ESCA, in Spanish)** for years 2015 to 2018 (as mentioned above, there is no health information available in the birth records data). ESCA provides information regarding health and health care use outcomes. It is conducted in several waves within a year, allowing us to precisely identify the pre and post attacks period. It samples all Catalan Health Areas (AGA) which includes one or more municipalities.

2.3.2 Treatment and Control Groups

As mentioned in the introduction, our main approach compares differences across Muslim vs non-Muslim mothers (first difference) and mother's municipality of residence, affected vs non-affected cities (second difference), before and after the attacks.

Regarding the first difference, we use the variable that identifies the mother's country of birth (from INE birth records) in order to construct the treatment group (Muslim mothers). We classify a woman as born in a Muslim country when in that country: 1) there is a majority or strict majority of Muslims (followers of the religion of Islam) or 2) the distinction between religion and the state is not clear. Therefore, our treatment group includes only immigrant Muslim mothers, as defined by the country in which they were born (following the definition above). Thus, when we mention "Muslim" mothers, we are referring to this definition. We always select mothers according to their country of birth.[12] For the control group, we consider all other immigrant women (born in a country that is not Spain and does not qualify for the Muslim country definition above) who had a child in the period of analysis, excluding mothers born in Spain.

Using these definitions, we have 28,682 live births (conceived before the attacks) from immigrant Muslim mothers, which represents 38.21% of the total live births from non-Spanish born mothers living in Catalonia for the years 2015-2018, as shown in Table 2.1. In the Appendix, in Table 2A.1, we show the list of countries from which Muslim mothers come from (in our sample) along with the share of population that is Muslim in each of these countries. We do something similar with the non-Muslim group and present, in Figure 2B.1, the distribution of countries in which these mothers were born.

For the second difference, our treated cities group is constructed according to mothers' municipality of residence. Thus, we define as "treated cities" those that

[12]As the reader might wonder, with this definition we are using only first-generation Muslims, but by comparing them with other immigrants we avoid, to a large extend, the inclusion of second-generation immigrants born in Spain in the control group. In any case, we are unable to identify second generation Muslims in the birth records data.

Table 2.1: Live Births in Catalonia

Mother's country of origin	Freq.	Percent	Cum.
Not Muslim	46,388	61.79	61.79
Muslim	28,682	38.21	100.00
Total live births	75,070	100.00	

Elaboration: The Authors.
Source: National Registry of live births in Spain (INE, 2018).

directly received the attacks, Barcelona and Cambrils. We also include the municipality of Ripoll in this group of treated cities, as the perpetrators of the attacks were Muslim residents (and raised) in this particular city, which is also likely to be a strong source of Islamophobia for Muslim residents in this municipality. The non-affected cities group consist of all other municipalities in Catalonia that did not receive the attacks (and were not home of the perpetrators). For example, if a mother's city of residence is Cambrils, this live birth is considered as in the "treated cities group".

2.3.3 Descriptive Statistics

Table 2.2 shows summary statistics for the sample analyzed. We split the sample by treated vs untreated cities, Muslim vs non-Muslim mothers in the pre- and post-periods. For instance, we see that for babies born from Muslim mothers before the attacks, mean birth weight was 3,336.69 g; the share of low birth weight newborns was 5.60%; the share of pre-term deliveries, complications and c-sections was 4.24%, 10.51% and 24.86%, respectively.

Table 2.2: Descriptive Statistics. Means

Variables	Muslim		Non-Muslim		Treated cities		Non-treated cities	
	Before	After	Before	After	Before	After	Before	After
Newborns outcomes								
Birth weight (in grams)	3,336.69	3,329.08	3,287.02	3,297.76	3,276.09	3,283.84	3,314.53	3,318.65
LBW indicator	0.0560	0.0579	0.0677	0.0641	0.0691	0.0652	0.0615	0.0606
Female indicator	0.4790	0.4885	0.4889	0.4874	0.4862	0.4836	0.4844	0.4890
Childbirth characteristics								
Share of pre-terms	0.0424	0.0391	0.0631	0.0524	0.0561	0.0522	0.0549	0.0451
Share of complications	0.1051	0.1010	0.1184	0.1106	0.1167	0.1101	0.1118	0.1063
Share of c-sections	0.2486	0.2283	0.2847	0.2809	0.2899	0.2952	0.2634	0.2482

Elaboration: The Authors.
Source: National Registry of live births in Spain (INE, 2018).

2.4 Empirical Strategy

As mentioned before, jihadist terrorist attacks have increased the rejection against the Muslim community. This increased rejection has been shown in the literature to be higher in affected places. In this sense, the underlying intuition is that the attacks should primarily affect Muslim women living in the affected cities and should have no effect on other groups of women. At this point, we must acknowledge that one of the control groups (Muslim women living in non-affected cities) might also be affected to a certain extent by the increase in Islamophobia. However, due to the larger distance to the place of the attacks, the impact should be smaller than for our treated group of Muslim mothers in affected cities.

To assess the impact of the attacks on health at birth outcomes, we rely on a difference in differences in differences (DDD) approach. We compare differences across mother's country of origin (first difference), and mother's municipality of residence (second difference), before versus after the attacks. The first difference compares Muslims vs non-Muslims mothers (excluding Spanish-born mothers), and the second difference compares affected vs non-affected cities.

Our baseline specification is a DDD regression relating health at birth outcomes for newborn *i*, whose mother's municipality of residence is *m*, conceived in month *t*, year *y*, and whose mother's country of origin is *c*, as follows,

$$\begin{aligned} Y_{imtyc} = {} & \alpha + \beta_1 TreatedCity_m + \beta_2 Post_{ty} + \beta_3 Muslim_c \\ & + \beta_4 TreatedCity_m * Post_{ty} + \beta_5 TreatedCity_m * Muslim_c \\ & + \beta_6 Muslim_c * Post_{ty} + \theta TreatedCity_m * Post_{ty} * Muslim_c \\ & + \delta_m + \rho_t + \gamma_y + \tau_c + \epsilon_{imtyc} \end{aligned} \tag{2.1}$$

where Y_{imtyc} represents a birth outcome for a live birth conceived prior to the attacks. As outcomes we use: birth weight (in grams), low birth weight, prematurity, complications (during labor), c-section, and following Brown (2014), we use whether the newborn is a female since intrauterine exposure to stress may reduce male births (Catalano et al., 2006).

$TreatedCity_m$ takes the value of one when mother's municipality of residence is Barcelona, Cambrils or Ripoll, $Post_{ty}$ is an indicator of the occurrence of the attacks (1 if child was born after the August 2017 attacks and conceived up to right before the attacks), and $Muslim_c$ is a dummy variable taking the value of 1 for mothers born in a Muslim country and zero for mothers born in a different country (except Spain). Thus, the treatment group includes all mothers whose country of origin is a Muslim one and are residents of one of the affected cities. The coefficient of interest is $\boldsymbol{\theta}$, which captures the effect of the Catalonia terrorist attacks on newborns from Muslim mothers living in the treated cities. We include month ρ_t, year γ_y of conception fixed effects, to account for potential common time shocks affecting the outcomes, mother's country of birth τ_c fixed effects, and mother's municipality of residence δ_m fixed effects. By doing this, we use variation at the local level, rather than province level, as not everyone is equally affected within a province. We cluster standard errors at the mother's municipality of residence, since this is the level at which the effect takes place, accounting for any unobserved common group effects (Wooldridge, 2003; Cameron et al., 2008).

For the estimates in which we distinguish the trimester of pregnancy at the time of the attacks, we run the same specification as in Equation (2.1), but classify the exposure to the attacks (event) according to the trimester of gestation in which the woman was. Thus, in this case we have three coefficients of interest, θ_1, θ_2, and θ_3 that represent the effect (on birth outcomes) of being exposed to the attacks during the first, second or third trimester of pregnancy, respectively, for Muslim mothers living in treated cities.

The difference in difference estimator provides an unbiased estimate of the treatment if, in the absence of the treatment, the outcomes in the two groups (control and treatment) would have followed parallel trends, before and after the attacks. Our identification strategy is based on several pillars. First, the source of stress coming from the attacks is completely exogenous. Although, the threat of attacks was, in general, expected across western countries, the exact location, timing, and severity of a terrorist attack would have been impossible to predict by the general population living in Spain. Second, even though Muslim and non-Muslim mothers might differ in terms of observable characteristics, we do not expect these differences to change as a result of the attacks. In addition, we are comparing across immigrant population (excluding natives). Third, the DDD model allows us to control for two kinds of potentially confounding trends: changes affecting health at birth outcomes of Muslim newborns, unrelated to the attacks, and changes affecting health at birth outcomes of all Muslim and non-Muslim babies living in the treated cities. Fourth, to support our econometric strategy we perform several robustness checks. For example, we allow pregnant women whose city of residence is not affected to also be exposed to the attacks by using linear distance to the closest affected municipality; we re-estimate Equation (2.1) excluding Barcelona, Ripoll, and Cambrils cities (one at a time), to test whether our results are mostly driven by one city. Finally, we perform a placebo test using a fake treatment group. It has been established in the literature that women may change their behaviour when faced to potential threats to their integrity. However, as we mentioned before, the timing of any attack is unpredictable and, therefore, the possibility that women change their place of residence is very unlikely in our analysis (in addition, our post-event period includes only 1 year). In any case, to ensure the reliability of our model, it is important to analyze if there have been major changes to Muslim immigration after the attacks. In other words, we should examine whether Muslim women leave or emigrate, which could lead to selective selection in our sample. On this point, we provide evidence supporting that this is not the case; thus, first, we do not see an (absolute) increase in the number of Muslim women leaving the country (emigration). Second, we do not see that net emigration slowed down during the period of analysis (2015 to 2018). In the Appendix, in Table 2A.2, we provide data on emigration and net immigration of Muslim[13] women obtained from the National Institute of Statistics in Spain.

One might also be concerned that other channels rather (or in addition to) than stress may be the main driving forces, for instance lack of healthcare. However, on this point, in Spain's context universal health coverage is guaranteed to all pregnant women, regardless of immigration status. For example, in our sample 99.38% of births took place in hospitals. In addition, the Catalonia attacks did not affect the provision of health services at any level of care. Therefore, it seems plausible that the main driving force in our context is the stress affecting Muslim mothers. In the following sections, we will come back to these other alternative channels.

2.5 Effects on Health at Birth

2.5.1 Baseline Results

Table 2.3 shows the estimates of Equation (2.1). We report results for a battery of health at birth outcomes: birth weight, share of low birth weight newborns, share of pre-term childbirths, whether there were complications during labor, share of female babies, and share of c-sections. All regressions include mother's municipality of residence fixed effects, month and year of conception fixed effects, and mother's country of origin fixed effects.[14] Standard errors are clustered at the mother's municipality of residence level.

Starting with newborns' health outcomes, we get the expected direction in all the coefficients. We see a decrease in birth weight of 12.89 g (though not significant), and a significant (at a 1% level) increase in the share of low birth weight babies

[13]In here, a women is considered Muslim using the same definition of our baseline equation (Equation (2.1)).

[14]Regarding mothers' country of origin FE, given the sample size, one may be concerned about the number of observations per country across treatment and control cities and over time, which is important in order to understand which mothers are contributing to identification. In order to show that our results are robust to this potential issue, we re-estimate Equation (2.1) excluding these FE's. Results are very similar and are available upon request.

Table 2.3: Baseline Estimates Birth Outcomes

	(1)	(2)	(3)	(4)	(5)	(6)
	Birth Weight	**LBW**	**Prematurity**	**Complications**	**Female**	**C-Section**
DDD City*Muslim*Post	**-12.8937**	**0.0160*** **	**0.0058**	**0.0147** **	**0.0233***	**0.0122**
	(13.2413)	(0.0059)	(0.0046)	(0.0073)	(0.0118)	(0.0092)
Interaction City*Post	-2.9874	-0.0036	0.0085**	-0.0020	-0.0106	0.0155**
	(6.6793)	(0.0047)	(0.0034)	(0.0057)	(0.0068)	(0.0060)
Interaction City*Muslim	4.2772	0.0007	-0.0010	0.0075	-0.0145**	-0.0074
	(14.3387)	(0.0041)	(0.0034)	(0.0046)	(0.0057)	(0.0062)
Interaction Muslim*Post	-15.8251	0.0018	0.0077*	0.0012	0.0052	-0.0155*
	(10.6027)	(0.0056)	(0.0042)	(0.0069)	(0.0103)	(0.0090)
Mean dependent variable	3,282.19	0.0673	0.0427	0.1129	0.4705	0.2736
% Effect	**-0.39%**	**23.77%**	**13.58%**	**13.02%**	**4.95%**	**4.46%**
Municipality of residence FE	x	x	x	x	x	x
Month, year of conception FE	x	x	x	x	x	x
Mother's country origin FE	x	x	x	x	x	x
Observations	65,519	65,519	75,070	75,070	75,070	75,070

Notes: OLS estimates of Equation (2.1). Outcomes for the live births from mothers whose region of residence is Catalonia (includes the provinces of Girone, Barcelona, Lleida and Tarragona). Standard errors are clustered at the mother's municipality of residence level, 121 clusters. The row "mean dependent variable" shows the mean for the treated group before the Catalonia attacks (August 2017). ***p<0.01, **p<0.05, *p<0.10.

by 1.60 percentage points for our group of interest. Related to the average birth weight for immigrant Muslim mothers (living in the affected cities) before the attacks, which was 3,282.19 g, the effect is translated into a decrease of 0.39% and an increase of 23.77% in the share of low birth weight newborns.

With regards to childbirth characteristics, we get a positive coefficient for the share of pre-term newborns, though not significant. For the complications during labor, we find a significant increase of 1.47 percentage points which, compared to the pre-attacks mean, is translated into an increase of 13.58% in the share of complications for immigrant Muslim mothers living in affected cities. We also see a positive significant result for the share of female newborns (however, as we will show in the robustness check section - event study, we do not interpret the result of this outcome as causal due to the existence of a pre-trend). For instance, there is a 4.95% increase (with respect to the pre-attacks mean) in the share of female babies from (immigrant) Muslim mothers living in affected cities. Regarding c-sections, we do not find any significant result, even though the coefficient of interest has the expected (positive) sign.

Overall, results in Table 2.3 reveal that intrauterine exposure to stress faced by immigrant Muslim mothers living in the cities affected by the 2017 Catalonia terrorist attacks has detrimental effects on newborns' health. Something that is worth mentioning is that an attack such as the one analyzed here, might affect stress levels of all mothers in the region. This means that the control group (non-Muslim mothers) may not be totally unaffected by the event studied here. In any case, the negative effect we find is a lower bound. In addition, our results are supported by the medical literature that shows that increased levels of stress in pregnant women may lead to an unbalance in hormones that affect fetal grow and increases the probability of complications during labor, as well as birth weight (De Weerth and Buitelaar, 2005).

These findings are in line with other estimates available in the literature measuring the effects of terrorism on birth outcomes, that find negative effects on birth weight and an increased share of complications during delivery (Camacho, 2008; Mansour and Rees, 2012; Quintana-Domeque and Ródenas-Serrano, 2017). We find similar results in terms of magnitude, significance and direction of the estimated effects.

2.5.2 Trimester of Pregnancy

Following the literature on terrorism and health at birth, as well as the medical literature, we take into account the trimester of gestation in which the women was when the attacks took place (and therefore the exposure to stress). In Table 2.4 we show the estimates for the six different health at birth outcomes according to the trimester of exposure to stress.

From Table 2.4, we see that intrauterine exposure to stress throughout the three trimesters of gestation has detrimental effects on health at birth for newborns from (immigrant) Muslim mothers living in affected cities. In terms of statistical significance, from the F-statistics tests of equality of coefficients, we find that there is no evidence that the coefficients are the same across trimesters of pregnancy for all the outcomes, except for birth weight.

Regarding exposure during the first trimester, we find an increase in the share of low birth weight babies by 2.27 percentage points, an augmented share of preterm newborns (2.33 percentage points), a higher share of female newborns, and an increased share of deliveries through c-sections (2.93, and 2.34 percentage points, respectively). For the second trimester of gestation, we also find detrimental effects on low birth weight, an increase in the share of complications as well as in the female ratio by 2.04 and 3.455 percentage points, respectively. With respect to newborns' sex early in pregnancy, it has been shown that exposure to stress while in utero may be more dangerous for male fetuses. Thus, Sanders and Stoecker (2015) find that males are more vulnerable to side effects of maternal stress in utero, and Catalano et al. (2006) show evidence supporting male fetal loss as a consequence of exposure to stress. The conclusions of these papers are also in line with the medical literature that suggests that the probability of having a miscarriage is much higher early in the pregnancy. Finally, in the last trimester of pregnancy, we also

Table 2.4: Birth Outcomes and Trimester of Pregnancy

	(1)	(2)	(3)	(4)	(5)	(6)
	Birth Weight	**LBW**	**Prematurity**	**Complications**	**Female**	**C-Section**
DDD City*Muslim*Trim1	**-10.7676**	**0.0227*** **	**0.0233*** **	**-0.0026**	**0.0293** **	**0.0234** **
	(19.6477)	(0.0070)	(0.0045)	(0.0104)	(0.0146)	(0.0108)
DDD City*Muslim*Trim2	**-27.8500**	**0.0181***	**0.0013**	**0.0204** **	**0.0345** **	**-0.0184***
	(17.3327)	(0.0093)	(0.0087)	(0.0098)	(0.0150)	(0.0101)
DDD City*Muslim*Trim3	**-3.2481**	**0.0074**	**-0.0075**	**0.0268*** **	**0.0089**	**0.0252** **
	(13.2372)	(0.0060)	(0.0047)	(0.0085)	(0.0104)	(0.0104)
Interaction City*Event	-3.4581	-0.0034	0.0087**	-0.0020	-0.0108	0.0155**
	(6.7104)	(0.0048)	(0.0034)	(0.0057)	(0.0068)	(0.0059)
Interaction City*Muslim	4.6706	0.0006	-0.0011	0.0075	-0.0144**	-0.0074
	(14.3009)	(0.0041)	(0.0034)	(0.0046)	(0.0057)	(0.0062)
Interaction Muslim*Event	-15.8182	0.0019	0.0077*	0.0012	0.0052	-0.0155*
	(10.6170)	(0.0056)	(0.0042)	(0.0069)	(0.0103)	(0.0090)
Treated City	-54.7534***	0.0296***	-0.0062***	-0.0635***	-0.0254***	-0.0673***
	(7.3064)	(0.0022)	(0.0019)	(0.0027)	(0.0035)	(0.0041)
Muslim	-50.8355	-0.0015	0.0456**	-0.0562	0.1225*	-0.0619
	(60.3267)	(0.0464)	(0.0183)	(0.0403)	(0.0685)	(0.0890)
Mean dependent variable	3,282.19	0.0673	0.0427	0.1129	0.4705	0.2736
Municipality of residence FE	x	x	x	x	x	x
Month, year of conception FE	x	x	x	x	x	x
Mother's country origin FE	x	x	x	x	x	x
F-test						
Ho: θ_{trim1}= θ_{trim2}= θ_{trim3}	1.34	4.97***	58.50***	3.52**	2.99*	16.03***
p-value	0.2688	0.0084	0.0000	0.0328	0.0543	0.0000
Observations	65,519	65,519	75,070	75,070	75,070	75,070

Notes: OLS estimates considering the trimester of gestation. Outcomes for the live births from mothers whose region of residence is Catalonia (includes the provinces of Girone, Barcelona, Lleida and Tarragona). Standard errors are clustered at the mother's municipality of residence level, 121 clusters. ***p<0.01, **p<0.05, *p<0.10.

find an increase in the share of complications during delivery and a higher share of c-sections by 2.68 and 2.52 percentage points, respectively.

Our results are consistent with others in the literature that find that exposure to terrorism (particularly early in the pregnancy stage) is linked to negative health at birth outcomes (Camacho, 2008; Mansour and Rees, 2012; Quintana-Domeque and Ródenas-Serrano, 2017). However, we also find detrimental effects on the second and third trimester. This is in line with a number of studies that highlight that there is no consensus on whether we should expect prenatal exposure to stress to have a differential effects across trimesters (Mansour and Rees, 2012; Persson and Rossin-Slater, 2018).

2.6 Robustness Checks of Baseline Results

This section includes several robustness checks for our main results from Equation (2.1). First, we conduct an event study design to support the hypothesis of no differential pre-trends. Second, we examine whether our results hold when we allow pregnant women whose cities of residence are not affected to be also exposed to increased Islamophobia. Third, we explore whether our results are driven by one of the three cities. Finally, we perform a placebo test, in which we change the treatment group (Muslim mothers).

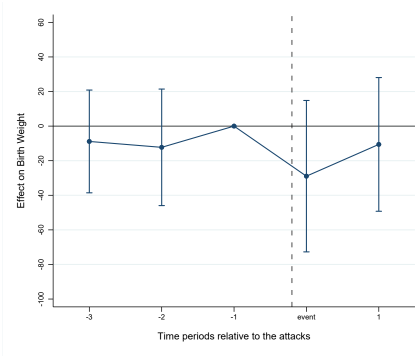
2.6.1 Event Study Estimates

To evaluate possible differential pre-trends between treatment and control groups, we conduct an event study version of Equation (2.1) for all the outcomes. For this, we generate three pre-attacks and two post-attacks periods (as dummy variables), both on an annual basis.[15] We set the period before the attacks $(t = -1)$ as the omitted category. In this specification, we also include month, year of conception fixed effects, municipality of residence fixed effects and mother's country of birth fixed effects. Figure 2.1 depicts the set of pre/post-attacks coefficients of the triple interaction (DDD) for the six health at birth outcomes, along with their 95 percent confidence intervals.

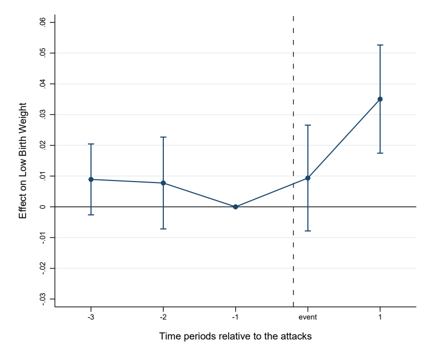
Overall, the set of pre-attack coefficients of the triple interaction are statistically indistinguishable from zero, with the exception of the very last period in the case of the share of female newborns. Regarding birth weight, the indicator of low birth weight, prematurity, complications during labor and c-sections, we do not find any significant coefficient for the triple interaction term in the pre-attack periods. In addition, we see that after the attacks there is an increase in detrimental health at birth outcomes, as indicated by the positive and statistically significant coefficients for the share of low birth weight newborns, prematurity, and the complications during labor indicator, which is in line with our baseline results. Regarding the results of the share of female newborns, due to the existence of one significant pre-attack coefficient, we do not interpret the results of this outcome as a causal impact of the attacks.

[15] As births may be seasonal, and we want to exploit all the time horizon we have available, we have constructed the pre/post-attacks periods using "year base" categories. Thus, we have three (annual) pre-attack time periods and two post-attack periods.

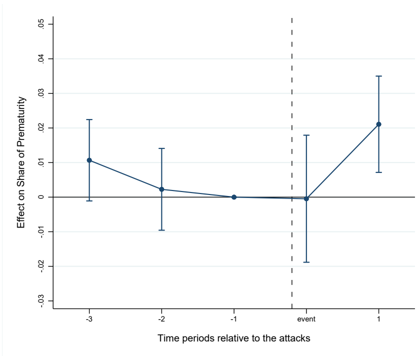
Figure 2.1: Event Study Graphs of Baseline Estimates



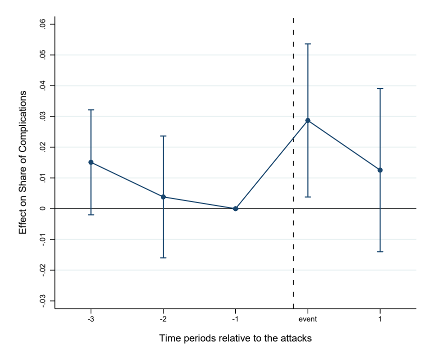
(a) Birth Weight



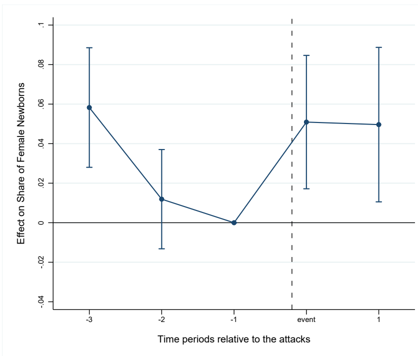
(b) Low Birth Weight



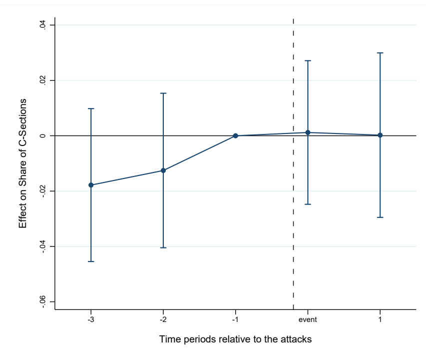
(c) Prematurity



(d) Complications



(e) Female



(f) C-Section

Notes: Figure shows the set of coefficients corresponding to the triple interaction term from the event study versions of Equation (2.1). Time periods are defined on an annual basis. Points estimates are shown along with their 95 percent confidence intervals. Reference category is one period prior to the attacks. Standard errors are clustered at the mothers' municipality of residence level, 121 clusters.

2.6.2 Allowing Other Cities to be Exposed

Our baseline specification establishes that only immigrant Muslim women living in affected cities are exposed to increased Islamophobia after the attacks and, consequently, to this source of stress. Nevertheless, it may be the case that Muslim women living in neighboring cities may also experience increased rejection and, therefore, this may also affect their newborns' health. To test whether this channel is affecting our results, we allow mothers whose municipality of residence is not one of the affected to be also exposed to stress coming from increased Islamophobia after a jihadist attack.

To do so, we re-estimate Equation (2.1) replacing the dummy $TreatedCity_m$ with a measure of distance to the treated cities. From here, we generate the pairwise interactions as well as the triple interaction. We use linear distances calculated using the latitudes and longitudes of every municipality in Catalonia. To every live birth, we assign the minimum distance between the mother's municipality of residence to Barcelona, Cambrils or Ripoll. In other words, we compute linear distances from mother's municipality to Cambrils, to Barcelona and to Ripoll and among these three, we choose the shortest distance. To model the distance, we follow the work done by Quintana-Domeque and Ródenas-Serrano (2017) and use a negative exponential form supporting the assumption that Muslim women will be more affected the closest they are to the places of the attacks. Therefore, the assumption that we are making is that stressful responses are likely to decay with physical distance from the site of the attack. Thus, as we define it, the variable distance can take the value of 0 (if the mother is residing in one of the affected municipalities: Barcelona, Cambrils or Ripoll) or the negative exponential distance between the mother's municipality of residence and the closest of the three attacked cities. Consistent with our hypothesis we find similar results, in terms of direction and magnitude, to our baseline estimates. Thus, exposure to the attacks in one of the affected municipalities (distance equals 0), translates into an increase in the probability of having detrimental health at birth. On the other hand, if the mother's municipality of residence is not one of the affected cities, then the increase in the probability of negative outcomes should be lower (compared to those living in Barcelona, Cambrils or Ripoll).

The results using this measure of distance are shown in Table 2A.3 of the Appendix. We can see that, overall, we obtain consistent results in terms of the direction of the estimated coefficients. For instance, we find significant detrimental effects in terms of birth weight, low birth weight, and the share of female newborns. At this point, it is important to notice that we are using a continuous measure of exposure for the affected cities rather than a binary measure, so that we are losing precision in some of our coefficients and the results are less significant. In any case, they point towards similar conclusions, in most of the cases, than our baseline findings.

2.6.3 Excluding Cities from the Treated Group

In this subsection we explore whether our results hold when we exclude from the sample the city of Barcelona, Ripoll, and Cambrils, one at a time. We re-estimate Equation (2.1) excluding, first, live births from Barcelona, then from Ripoll, and finally from Cambrils.

For the case of Barcelona, we might be worried that this city may be driving our results as it might have been more exposed to Islamophobia by the media, and also because it is the most populated city in Catalonia (specially if we compare it with Cambrils and Ripoll). To test this, we exclude all live birth from women living in Barcelona during the period of analysis. To support our empirical strategy, we should find statistically significant detrimental effects on health at birth outcomes. In Table 2A.4 of the Appendix, we report the results of the above specification. In general, we obtain negative effects on health outcomes. For instance, we see a large decrease in birth weight for newborns of Muslim mothers living in Cambrils and Ripoll after the attacks. We also find an increase in the share of c-sections by 6.05 percentage points. For the complications during labor indicator, low birth weight and prematurity, even if the coefficient is not significant, it has the expected sign. Overall, we see that the negative effects are not washed out when we exclude all live birth from women living in Barcelona despite the fact that we are eliminating a large number of observations in the treated group.

Second, we test the robustness of our findings by excluding Ripoll from the treated cities group. Ripoll is considered an affected municipality because the perpetrators of the attacks were all raised there. However, their treatment nature may not be the same as the other cities. Table 2A.5 in the Appendix shows the results of this exercise. We find the expected signs in all the coefficients of interest: we document an increase in the share of low birth weight babies by 1.49 percentage points and an increase in the share of complications and the female ratio by 1.34 and 2.13 percentage points, respectively.

Finally, as an additional robustness check, we do the same exercise excluding live births from Cambrils. Results of these estimates are shown in Table 2A.6. From here, we see that we get the expected signs for the coefficient of interest across all outcomes. In addition, we find detrimental statistically significant differences in the share of low birth weight newborns (1.63 percentage points), deliveries with complications (1.66 percentage points) and in the share of female newborns (2.80 percentage points). Therefore, we show that after excluding the observations from the city of Cambrils, the baseline effects are maintained.

2.6.4 Placebo

We conduct a placebo test to indirectly assess the common time trend assumption. We estimate Equation (2.1) using an alternative treatment group. We assign as placebo treatment units mothers born in Latin American countries and drop from the analysis those observations corresponding to Muslim mothers. The control group remains the same; live births from other immigrant mothers (excluding Spanish mothers). Because we expect that the increased stress due to Islamophobia affects mainly Muslim mothers living in the treated cities, we should not find an effect on Latin American pregnant women living in the same cities, otherwise our identification strategy would be challenged. Thus, we re-estimate Equation (2.1) using the group of Latin American immigrant women that had a live birth (conceived before the attacks) in Catalonia during the period of analysis as the treated group. Results of this exercise are shown in Table 2.5.

Table 2.5: Robustness Check - Placebo Test

	(1)	(2)	(3)	(4)	(5)	(6)
	Birth Weight	**LBW**	**Prematurity**	**Complications**	**Female**	**(C-Section)**
DDD City*Latin*Post	**10.0650**	**-0.0111**	**-0.0024**	**-0.0050**	**0.0088**	**-0.0237***
	(16.6348)	(0.0092)	(0.0074)	(0.0091)	(0.0147)	(0.0125)
Interaction City*Post	-9.3292	0.0028	0.0102*	0.0007	-0.0157	0.0289***
	(11.4172)	(0.0069)	(0.0057)	(0.0058)	(0.0120)	(0.0080)
Interaction City*Latin	-13.2252	0.0194***	0.0112***	0.0009	-0.0181**	-0.0004
	(8.7302)	(0.0038)	(0.0035)	(0.0056)	(0.0070)	(0.0070)
Interaction Latin*Post	-16.8169	0.0076	0.0111	0.0010	0.0043	0.0042
	(15.7449)	(0.0086)	(0.0069)	(0.0091)	(0.0144)	(0.0125)
Mean dependent variable	3,284.36	0.0719	0.0640	0.1325	0.4896	0.3186
Municipality of residence FE	x	x	x	x	x	x
Month, year of conception FE	x	x	x	x	x	x
Mother's country origin FE	x	x	x	x	x	x
Observations	42,031	42,031	46,388	46,388	46,388	46,388

Notes: OLS estimates of Equation (2.1). In the placebo estimation the treatment group is made up with mothers born in Latin American countries. Controls are mothers born in other foreign countries. Standard errors are clustered at the mother's municipality of residence level, 121 clusters. ***p<0.01, **p<0.05, *p<0.10.

Overall we find that none of the health at birth outcomes are statistically significant, except for c-sections (although it is negative: thus, it has the opposite sign compared to the baseline results). In addition, the DDD coefficients for all the outcomes (except the share of female) are in the opposite direction compared to the baseline results. Regarding the share of female, even if we find a positive coefficient, it is not statistically significant and much smaller than our baseline estimates. Thus, we conclude that there is no evidence of differential effects of the attacks on Latin American mothers living in affected cities, which supports our empirical strategy.

Altogether, we believe our identification strategy is robust to the several robustness tests that we have estimated. Our results hold when we allow mothers living close to affected cities to be exposed to the attacks-related stress; when excluding Barcelona, Ripoll, and Cambrils (one at a time) from the treated cities group. Our results are also robust to a placebo test, in which we replace the treatment group by one that should not be affected. Finally, the event study estimates add validity to our econometric strategy.

2.7 Assessing Mechanisms

Our main hypothesis is that Muslim women may be facing an increase in Islamophobia due to terrorist attacks perpetrated by jihadist groups, and this effect may be particularly salient in the cities where the attacks occur. We test this hypothesis using three alternatives. First, by examining whether there is an increase in the rejection against Muslims; second, we explore whether (reported) hate crimes increased as a consequence of the attacks. Finally, we assess the relation between the attacks and maternal stress, and whether stress may induce behavioral responses.

2.7.1 Rejection against Muslims

For the Islamophobia analysis, we use data from the Survey "Attitudes Towards Immigration" conducted by the *Centro de Investigaciones Sociológicas* of Spain for the years 2015, 2016 and 2017. The three waves of the survey are nationally representative and sample individuals aged 18 years and older that hold the Spanish nationality. It provides information on the perception of the Spanish population on immigration. The advantage of using this survey for our analysis is that the 2017 wave was conducted during the month of September, just one month after the attacks.

First, we assess whether there is an increase in rejection against the Muslim population. We use a question that asks Spaniards whether there is a specific group of immigrants for which they feel antipathy. The respondents can mention any group without being limited to pre-established categories. Using the responses to this question, we construct our first outcome variable, a dummy *"Muslim"* taking the value of 1 when respondents mention any group that refers to the Muslim population,[16] and zero otherwise. Second, we assess whether negative feelings about

[16] For example: Moroccan, Muslims, Islamists, "moro".

immigration have increased after the attacks in affected relative to non-affected areas. We rely on a question that asks Spanish individuals whether they consider immigration to be a positive or a negative issue for Spain. Using the responses to this question, we construct a dummy variable that is one when respondents declare immigration as "negative" for the country, and zero otherwise. Third, we explore any changes in hate speech[17] as a consequence of the attacks. We use information from a question asking Spaniards whether, in the last weeks, they have seen or heard any offensive comments against immigrants. As can be noted, just the first outcome allows us to directly identify rejection against the Muslim community due to restrictions in data availability as well as the contents of the questions asked in the survey. The other two outcomes are targeted to all immigrants, and we use them as an indirect measure of increased rejection against the Muslim community (as they are also included in the immigrant group).

To assess whether the level of rejection increased in the affected cities after the attacks, we need to look at the municipality level. However, in this particular dataset the lowest geographical level available is the province and, thus, we construct a dummy variable taking the value of one for the provinces to which the affected municipalities belong (Girona, Tarragona and Barcelona), and zero otherwise. Therefore, for this subsection, the treated group is made up of the provinces that include the cities affected by the attacks.

To formally explore our hypothesis, we use a difference-in-differences approach to test whether there is an increase in the rejection against immigrants and Muslims in the affected provinces after the attacks in Catalonia (relative to non-affected provinces). To do so, we estimate the following equation,

$$\begin{aligned} Y_{ipt} = \alpha + \beta_1 Province_p + \beta_2 Post_t + \beta_3 Province_p * Post_t \\ + \delta_p + \gamma_t + \epsilon_{ipt} \end{aligned} \tag{2.2}$$

Where Y_{ipt} represents the "rejection against immigrant" outcome of individual *i* in province *p* in year *t*. $Province_p$ is the variable for the control - treatment group, and takes the value of one for the affected provinces, and zero otherwise. $Post_t$ is a dummy variable taking the value of one for the post attacks period, 2017, and zero otherwise. The coefficient of interest is β_3, which captures the effect of the terrorist attacks on the outcomes of interest. We include time fixed effects γ_t to account for potential common time shocks affecting the outcomes. We also include province

[17]Hate speech comprises any abusive or threatening speech or writing that expresses prejudice against a particular group, especially on the basis of race, religion, or sexual orientation (Definition taken from the Oxford Dictionary).

fixed effects δ_p to account for any time invariant shock across provinces. We cluster standard errors at the level of treatment (province). Probability weights are included to take into consideration the survey design.

Before moving on to the regression results, we first plot the raw trends of the three outcomes in Figure 2B.2 of the Appendix 2B. Panels (a), (b) and (c) show that, after the attacks, there is an increase in the Islamophobia-related outcome in the treated locations. In addition, trends between affected and non-affected provinces look parallel before the attacks. With respect to regression results, Table 2.6 presents the estimates of Equation (2.2). Column (1) shows an increase in the share of nationals feeling antipathy towards Muslims in the provinces of Barcelona, Tarragona and Girona (relative to other Spanish provinces) by 5.46 percentage points after the Catalonia 2017 attacks. This effect represents an increase of 22.84% with respect to the pre-attacks mean. This result supports the statement that, after a terrorist attack perpetrated by a jihadist group, the affected places may experience an increase in the rejection against the Muslim community.

Table 2.6: Effects of Catalonia 2017 Attacks on Perception towards Immigration

	(1)	**(2)**	**(3)**
	Rejection Muslim	**Neg feeling**	**Hate Speech**
DD Province*Post	**0.0546****	**0.0560****	**0.0524****
	(0.0261)	(0.0209)	(0.0258)
Treated Provinces	-0.0819***	0.1029***	0.0944***
	(0.0084)	(0.0092)	(0.0143)
Post (September 2017)	-0.0060	-0.0754***	-0.1911***
	(0.0131)	(0.0218)	(0.0343)
Mean dependent variable	0.2391	0.3943	0.4917
% Effect	**22.84%**	**14.20%**	**10.66%**
Province FE	x	x	x
Year FE	x	x	x
Observations	6,959	5,588	7,334

Notes: OLS estimates of Equation (2.2). Column (1) shows the result for the share of rejection against Muslim. Column (2) shows results for negative feeling towards immigration and Column (3) the share of hate speech against immigrants. Standard errors are clustered at the province level, 50 clusters. ***p<0.01, **p<0.05, *p<0.10.

From column (2), we infer that the attacks increased the share of Spaniards having a negative feeling about immigrants by 5.60 percentage points in affected provinces relative to the non-affected ones. This represents an increase of 14.20% with respect to the average pre-attacks share of respondents having a negative feeling about immigration (which was 39.43%). Column (3) shows that, regarding hate speech, the difference between affected provinces and the rest has increased by 5.24 percentage points after the attacks. This represents a positive variation of 10.66% with respect to the pre-attacks mean in the treated group (provinces).

As shown above, we have presented evidence supporting an increase in Islamophobia after the terrorist attacks, which is in line with our expectations of the channels affecting Muslim pregnant women. These results are consistent with other papers examining how hate crimes (against Muslims) increase after a terrorist attack perpetrated by a jihadist group, as described in Section 2.2 . Moreover, the general negative feeling about immigration also increases in treated provinces (compared to the control group), as well as hate speech against immigrants.

To support the robustness of the results of this subsection, we conduct a placebo test and perform an event study design. For the placebo test, we focus on the outcome of rejection against Muslims, as it is the only direct measure of Islamophobia available. We define a placebo outcome using the same question of the survey as in column (1) of Table 2.6, but now measuring the rejection of Spaniards against immigrants coming from Latin American countries. To do this, we construct a dummy variable taking the value of one when respondents declare they feel antipathy against an immigrant born in a Latin American country (and zero otherwise), and re-estimate Equation (2.2). Results of these estimates are shown in Table 2A.7 of the Appendix 2A. The intuition behind using a placebo outcome is to test an outcome that, theoretically, should not be affected by the attacks (treatment) (Yamamoto, 2016). In this case, we expect that Latin American immigrants should not be affected by the attacks in terms of increased rejection against them. From Table 2A.7, we see that the difference-in-differences coefficient is not significant and goes in the opposite direction, which supports our hypothesis and provides reliability on the robustness of the identification strategy. In addition, in Figure 2B.3 of the Appendix 2B, we provide the raw trends of this placebo outcome. As it can be seen, prior to the attacks, rejection against Latin American immigrants was higher in non-affected provinces and, after the attacks, the raw trend does not show an increase in affected provinces.

For the event study design, we use pre-attack years to provide evidence in support of the parallel trends assumption. Due to data availability, we only have 3 time periods to conduct the analysis. With this in mind, we replace the variable $Post_t$ of Equation (2.2) by year dummies and take as the omitted category the period before the attacks (t=-1). We show the results of this exercise in Figure 2B.4 in the Appendix 2B. Here, we plot the set of difference in differences coefficients along with their 95 confidence intervals. Overall, we find no evidence of differential pre-trends in any of the outcomes, which adds validity to the results of this subsection.

2.7.2 Hate Crimes

As mentioned throughout the paper, Islamophobia can be experienced in several ways. One of the elements that can capture Islamophobia is reported hate crimes. To assess whether hate crimes have increased as a consequence of jihadist terrorist attacks in affected cities, we use two sources of administrative data. First, we rely on data from the Catalan Police on reported hate crimes for the years 2015, 2016 and 2017. The data is aggregated at the month-year level for municipalities in Catalonia. It includes reported hate crimes regardless of their origin, this means that they can be due to Islamophobia, sexual orientation, ethnicity rejection, and others. Despite the fact that we are not able to isolate Islamophobia-related hate crimes, we can still use the data to estimate an indirect measure of those types of hate crimes. In addition, we can look at the most local level, municipality, which fits very well with our econometric identification.

We formally test whether the number of hate crimes increased after the attacks in affected relative to non-affected cities by estimating a difference-in-differences (DD) model that compares treated cities (Barcelona, Cambrils and Ripoll) to non-affected cities in Catalonia before and after the attacks. To do this, we estimate the following equation,

$$\begin{aligned} Y_{imt} = \alpha + \beta_1 TreatedCity_i + \beta_2 Post_{mt} + \beta_3 TreatedCity_i * Post_{mt} \\ + \delta_i + \gamma_m + \rho_t + \epsilon_{imt} \end{aligned} \tag{2.3}$$

Where Y_{imt} is the natural logarithm of the total number of reported hate crimes in municipality *i*, in month *m* in year *t*. $TreatedCity_i$ is the dummy for the control-treatment group taking the value of one for affected municipalities, and zero otherwise; $Post_{mt}$ is an indicator taking the value of one for the post attacks period, and zero otherwise. β_3 is our coefficient of interest and captures the effects of the attacks on local hate crimes. We also include municipality fixed effects δ_i, month γ_m and year ρ_t fixed effects. We cluster standard errors at the municipality level.

Table 2.7, panel (A) presents the estimates of Equation (2.3). We find a significant increase of (approximately) 26.19% in local hate crimes in affected cities after the attacks compared to non-affected cities, which supports our hypothesis of increased Islamophobia after a jihadist attack. To test the robustness of these results, we first perform a placebo outcome test to show that our results are not driven by a general rise in crimes, and second, and an event study model of Equation (2.3).

Regarding the placebo, we test whether the increase in reported hate crimes can be explained by a general increase in crimes across municipalities in Catalonia. For

Table 2.7: Hate Crimes and Placebo Tests

	(A)	(B)
	Actual treated group	**Placebo outcome**
	(ln) Hate crime	**(ln) Falsification crime**
DD Treated City*Post	**0.2619***	**-0.1241**
	(0.1478)	(0.1040)
Treated City	0.1814***	0.5043***
	(0.0205)	(0.0140)
Post attacks	0.1690***	-0.0517
	(0.0350)	(0.0696)
Municipality FE	x	x
Month FE	x	x
Year FE	x	x
Observations	7,308	4,060
Number of clusters	203	60

Notes: Panel (A) presents the estimates of Equation (2.3) where the outcome variable is the natural logarithm of the total number of reported hate crimes in the Catalonia Region using the actual treated units. Robust standard errors in parentheses clustered at the municipality level. The data is aggregated at the municipality level. Panel (B) shows results of Equation (2.3) where the outcome is the natural logarithm of falsification crimes reported in Catalonia. In this estimates, standard errors are clustered at the "Área básica policial" level. The data is aggregated at the "basic police area" level where each area includes several municipalities. ***p<0.01, **p<0.05, *p<0.10.

this, we use falsification crimes as a placebo outcome.[18] The underlying intuition is that we should not expect a positive significant result in the DD coefficient. To formally test this, we re-estimate Equation (2.3) where the outcome is the natural logarithm of the total number of reported falsification crimes in the police area i, in month m in year t. Results are shown in Table 2.7, panel (B). We find a non-significant reduction in this type of crimes. This estimate supports our hypothesis of increased hate crimes as a plausible consequence of Islamophobia, rather than a general increase in crimes in affected cities after the attacks. With respect to the event study specification, we find one significant coefficient in the very last pre-attacks period, as shown in the Appendix 2B in Figure 2B.5. However, we must keep in mind that this outcome includes all types of hate crimes.

As a second exercise to see whether hate crimes have changed after the attacks, we use administrative records provided by the "Ministerio del Interior", which is

[18] As mentioned in previous sections, falsification crimes include documents forgery, identity usurpation and other related crimes.

the institution in charge of compiling all criminal records in the country. This administrative data set includes all reported hate crimes in Spain between 2015 and 2018. These records have the advantage that the nationality of the victim is reported. Thus, we are able to identify hate crimes against individuals that have a Muslim nationality. We define a victim as Muslim in the same way than in our baseline specification. In this data set the unit of observation is the hate crime for which we have information on the province, time of occurrence at the year level, and nationality of the victim. Despite the fact that the data is at the province level instead of at the municipality level, we believe that this exercise adds important evidence on the increase in hate crimes against Muslims after the attacks.

We apply a triple difference in differences econometric strategy to see whether hate crime rate changes after the attacks. In here, we compare in the first difference Muslim victims vs. other immigrant victims (excluding Spanish individuals), in the second difference affected provinces vs non-affected provinces, before vs after the attacks. Formally, we estimate a regression of the form:

$$\begin{aligned} Y_{pnt} &= \alpha + \beta_1 TreatedProv_p + \beta_2 Post_t + \beta_3 Muslim_n \\ &+ \beta_4 TreatedProv_p * Post_t + \beta_5 TreatedProv_p * Muslim_n \\ &+ \beta_6 Muslim_n * Post_t + \theta TreatedProv_p * Post_t * Muslim_n \\ &+ \delta_p + \rho_n + \gamma_t + \epsilon_{pnt} \end{aligned} \tag{2.4}$$

where Y_{pnt}, the outcome of interest, represents hate crime rate (per 100,000 inhabitants) against immigrants of nationality n, in province p and year t. $TreatedProv_p$ takes the value of one for the affected provinces (Barcelona, Girona, and Tarragona) and zero otherwise. $Post_t$ is a dummy variable that takes the value of one after the attacks and zero before. $Muslim_n$ takes the value of one when the victim is of a Muslim nationality and zero for all other immigrant nationalities. In this specification, as in the rest of the paper, we do not consider native individuals. The coefficient from the triple interaction θ shows whether the hate crime rate has changed for Muslims in affected provinces after (relative to before) the attacks. We include province δ_p, nationality of the victim γ_n, and time ρ_t fixed effects. Standard errors are clustered at the province level.

Results from the exercise described above are shown in Table 2.8. Here, we find a statistically significant increase in the hate crime rate against victims of Muslim nationalities in affected provinces after the attacks, as shown by the coefficient of the triple interaction term. Compared to the mean before attacks, this result corresponds to a 28.07% increase in the hate crime rate for our group of interest. The results of this section provide additional evidence in support of an increase in anti-Muslim hate crimes associated with the attacks. Finally, to add validity to this result, we perform an event study design which is shown in Figure 2B.6 of Appendix 2B.

Table 2.8: Hate Crimes and Nationality of the Victim

	Triple Difference in differences
	Hate crime rate (per 100,000 inhab.)
DDD Treated Prov*Muslim*Post	**0.0096****
	(0.0046)
Interaction Treated Prov*Post	0.0043***
	(0.0009)
Interaction Treated Prov*Muslim	0.0299***
	(0.0058)
Interaction Muslim*Post	0.0014
	(0.0020)
Mean dependent variable	0.0342
% Effect	**28.07%**
Province FE	x
Nationality of the victim FE	x
Time FE	x
Observations	19,136
Number of clusters	52

Notes: Table shows estimates of the triple difference in differences Equation (2.4) where the outcome is the hate crime rate against all immigrant nationalities. Robust standard errors in parentheses clustered at the province level. ***p<0.01, **p<0.05, *p<0.10.

2.7.3 Terrorist Attacks and Stress

In the previous sections we presented evidence supporting an increase in Islamophobia. As we stated before, this increased Islamophobia may be a source of stress affecting pregnant Muslim women. Due to data limitations, we are not able to identify any stress-related outcomes directly from the birth records data. However, we can still use other sources of data to assess whether Muslim women living in affected cities are facing increased levels of stress relative to other women. For this, we use the Catalan Health Survey (ESCA), conducted by the Health Department in Catalonia, for the years (cross-sections) 2015, 2016, 2017 and 2018. The survey compiles information about health, lifestyle and use of health services of the population living in Catalonia.[19] It is region-representative and samples individuals at

[19]Unfortunately, ESCA does not provide information on prenatal health care use.

the "basic health area" (AGA) level.[20] Every AGA is composed by one or more municipalities.

ESCA provides several variables related to health issues. For our purpose, we restrict the sample to women of childbearing age (between 15 and 45 years old) and use four questions related to stress. First, we utilize a binary measure of self assessed health, which takes the value of one when the respondent declares having bad or very bad health, and zero otherwise. In the survey individuals are also asked about emotional well-being measured through stress-related questions. Thus, we use these questions and construct three dummy variables corresponding to one when respondents declare they have felt relaxed, self-confident, and happy during the last two weeks; and zero otherwise. In addition, we perform a first component of a Principal Components Analysis (PCA) of the three above described variables. It is constructed in a way that the lower the index the worst is the stress condition of the women. We name this outcome as "emotional well-being index".

We formally test whether Muslim women living in affected cities experienced increased levels of stress after the attacks using the same specification as in Equation (2.1), where now the outcomes of interest are; self-assessed health, the three stress indicators and the emotional well-being index. As in the baseline estimates, we compare Muslim vs non-Muslim immigrant mothers (excluding Spanish women). We also include month and year fixed effects, women's country of birth fixed effects and AGA fixed effects. Standard errors are clustered at the AGA level. To support our hypothesis, we should find a positive coefficient for the self-assess health outcome and a negative coefficient for the stress indicators, which supports the idea that Muslim women living in the affected cities experience a decrease in emotional well-being.

Table 2.9 shows that being a Muslim women living in affected cities is associated with a decrease in the probability of being happy by 49.84 percentage points after the attacks. For the other outcomes, even if we do not find a statistically significant result, we get the expected signs. These results are consistent with stress affecting mental wellness, which is an important channel through which the fetus can be affected while in utero.

To address possible concerns regarding different trends between Muslim and non-Muslim women before the attacks, and also to say more about the impact after the attacks, we estimate an event study version of Equation (2.1) for the five stress markers outcomes. We use three pre-attacks and one post-attacks periods, where the reference category is the period before the attacks (t=-1). Overall, we do not find evidence of pre-trends in any of the outcomes examined. We find a significant

[20] AGA is the territorial health government in which the Catalan health system is organized.

Table 2.9: Terrorist Attacks and Stress

	(1)	(2)	(3)	(4)	(5)
	Bad Health	**Relaxed**	**Self-confident**	**Happiness**	**Emotional Well-being Index**
DDD City*Muslim*Post	**0.0622**	**-0.2082**	**-0.3399**	**-0.4984***	**-1.3353**
	(0.2210)	(0.1510)	(0.3781)	(0.2647)	(0.9060)
Interaction City*Post	-0.1095*	-0.0560	0.0680	-0.0685	-0.0360
	(0.0555)	(0.1108)	(0.0765)	(0.0899)	(0.2730)
Interaction City*Muslim	-0.0218	-0.0431	0.0727	0.0287	0.1143
	(0.0426)	(0.0743)	(0.0735)	(0.0971)	(0.2513)
Interaction Muslim*Post	0.0879	-0.0133	-0.0426	-0.1127	-0.2272
	(0.1021)	(0.1011)	(0.0735)	(0.0859)	(0.2492)
Mean dependent variable	0.0488	0.5161	0.8709	0.8387	-0.1661
Month FE	x	x	x	x	x
Year FE	x	x	x	x	x
Municipality (AGA) FE	x	x	x	x	x
Mother's country of birth FE	x	x	x	x	x
Observations	946	877	877	877	877

Notes: Sample includes years 2015 to 2018 of the Catalan Health Interview Survey provided by the Catalan Health Department. Women of childbearing age (15-45 years old). Estimation using weights according to survey design. Standard errors clustered at the "Area de region sanitaria" in Catalonia level, 43 clusters. ***p<0.01, **p<0.05, *p<0.10.

result for the "feeling relaxed" outcome but it is in the very last period. In addition, regarding the emotional well-being index, we see a significant and important drop in the first period which points to a decrease in emotional well-being for our group of interest right after the attacks. Results of the event studies are shown in Figure 2B.7 of Appendix 2B.

According to our hypothesis, the attacks should mainly induce stress in Muslim women living in affected municipalities, and should not have a significant impact on other groups of women. To test this, we conduct a placebo analysis in which we change the treated group by women born in Latin American countries (and exclude Muslim women from our sample). We re-estimate Equation (2.1) with the stress-related outcomes: bad health, relaxed, self-confident, happiness and the emotional well-being index. The results of this placebo estimates show that, for Latin American mothers, there is no significant association between the attacks and stress markers, except for the indicator of self-confidence but with the opposite sign. We present the results of this placebo test in Table 2A.8 of Appendix 2A.

So far, we have argued that the adverse health at birth consequences that we find are driven by maternal exposure to stress. However, this stress could induce behavioral responses and physical conditions in Muslim mothers which, in turn, may adversely affect newborn's health (Persson and Rossin-Slater, 2018). For example, it is possible that Muslim women respond to increased stress by taking up smoking or by changing their dietary habits. Those elements have been proven to increase the risk of developing diseases such as diabetes and, these conditions, have been documented to be a risk for the fetus. Moreover, high levels of stress have been linked to increased probabilities of developing high blood pressure which, in turn, increases the chances of having complications during labor, pre-term newborns and low birth weight babies. We study whether the Catalonia attacks are associated with the presence of high risk factors (hypertension and diabetes) and with being a smoker. In Table 2.10, we can see that the attacks are positively associated with being a smoker, and with the share of Muslim women (that live in the affected cities) having diabetes. For the hypertension indicator, although we do not find a statistically significant association, we get a positive sign in the coefficient of interest.

Table 2.10: Stress and Behavioral Responses

	(1)	(2)	(3)
	Smoking Status	**Hypertension**	**Diabetes**
DDD City*Muslim*Post	**0.2077***	**0.0183**	**0.2340****
	(0.1207)	(0.1258)	(0.1034)
Interaction City*Post	-0.1700***	0.0028	0.0022
	(0.0615)	(0.0517)	(0.0318)
Interaction City*Muslim	-0.0414	-0.1693*	-0.0194
	(0.0466)	(0.0932)	(0.0190)
Interaction Muslim*Post	-0.0213	0.2029**	-0.0072
	(0.0599)	(0.0819)	(0.0448)
Mean dependent variable	0.0189	0.0398	0.0348
Month FE	x	x	x
Year FE	x	x	x
Municipality (AGA) FE	x	x	x
Mothers' country of birth FE	x	x	x
Observations	877	946	946

Notes: Sample includes years 2015 to 2018 of the Catalan Health Interview Survey provided by the Catalan Health Department. Women of childbearing age (15-45 years old). Estimation using weights according to survey design. Standard errors clustered at the "Area de region sanitaria" in Catalonia level, 43 clusters. Smoking status equals 1 if daily smoker o occasional smoker. ***p<0.01, **p<0.05, *p<0.10.

To conclude this section, we have found evidence of an increased rejection against the Muslim community in affected places after the 2017 attacks, relative to non-exposed places. This represents a plausible channel of increased exposure to stress. In particular, an increase in the general negative feeling towards immigrants has been found in affected, compared to non-affected areas. We also find evidence in support of an increase in local hate crimes in exposed municipalities after the attacks. Moreover, we found that hate crime rate increases for Muslim victims in the affected provinces after the attacks. Finally, we find that being a Muslim women living in the target cities is associated with higher probabilities of reporting emotional stress. Previous evidence suggests that stress can induce behavioral and physical responses such as smoking and diabetes, which may potentially affect newborns' health. However, we recognize that there might be other potential alternative channels through which stress may affect Muslim women: for example, they could choose not to go out in public, or refuse to seek medical attention.[21] Moreover, increased hate crimes could lead to a reduction in labor supply by mothers or their husbands, opening the possibility for an income channel. Unfortunately, we are unable, due to data restrictions, to rule out these additional channels. In any case, we present robust evidence that Islamophobia-related stress is possibly one of the channels affecting health at birth.

2.8 Conclusions

Islamophobia is a phenomenon that has increased in the last decades since the 9/11 New York terrorist attacks perpetrated by jihadist groups. The usual target of this increased rejection are Muslim individuals, usually living in affected places. There are some papers that have found a link between Islamophobia and stress (Samari et al., 2018; Haque et al., 2019), which is particularly important for pregnant Muslim women as the evidence suggests that intra-uterine exposure to stress leads to detrimental health outcomes at birth. In turn, health at birth may influence adulthood outcomes such as earnings, educational attainment and well-being (Currie and Vogl, 2013). There are many factors that can affect health at birth: one of them is exposure to stress during pregnancy, that may have negative effects in terms of birth weight, complications during delivery, etc.

Regarding human capital, there are specific groups of individuals, in which immigrants are included, that are particularly vulnerable as a result of existing barriers to join the labor market, earn an equitable wage, or attend school, among others. If we add detrimental health at birth in the next generations to these already existing disadvantages, the gap between them and the host population becomes extremely difficult to eliminate.

In this sense, assessing the extent to which exogenous sources of stress, such as

[21] In this regards, as ESCA does not provide any information on prenatal care, we have used as indirect measures outcome variables such as (i) visits to the specialist, and (ii) visits to the gynecologist to see whether there are differences in health care use in our group of interest after the attacks. Estimates from this exercise do not show any significant result in terms of any change in health care use. Results of this exercise are available upon request.

those coming from terrorist attacks, affect health at birth of the immigrant Muslim community becomes important. It is relevant to understand if they have been object of rejection, discrimination and hate speech (Union of Muslim Communities in Spain, 2018) after a terrorist attack perpetrated by a jihadist group, as these types of situations have been proven to negatively affect pregnant women through stress.

In this chapter, we evaluate the effects of the Catalonia 2017 jihadist terrorist attacks on health at birth for newborns whose mothers are born in a Muslim country and reside in the affected cities. Using a triple difference in differences approach, we find that the attacks increase the share of newborns with low birth weight and increase the share of deliveries with complications for the population under study. We also present evidence regarding the timing of the attacks and the trimester of gestation. We find that exposure to stress, coming from the attacks, during the first trimester of pregnancy is linked to worse health at birth outcomes for our group of interest. This conclusion is consistent with other research assessing terrorist attacks and health at birth. In addition, we also find detrimental effects on the second and third trimester of gestation.

We also show that a plausible mechanism through which the attacks affect the group of interest might be maternal stress. This stress comes from increased rejection against the Muslim community -Islamophobia-, particularly after the attacks. We measure this increased rejection making use of survey data which analyzes attitudes towards immigrants in Spain. We find that the 2017 attacks increase the rejection against Muslims in affected provinces, increase the negative feeling towards immigrants and also contribute to an increase in the proportion of hate speech against immigrant groups in the provinces where the attacks took place. In addition, we provide an indirect measure of Islamophobia; an increase in the number of local hate crimes in target cities (relative to non-attacked cities) after the attacks. We go one step further and show that hate crime rates against Muslim victims increase after the attacks in affected areas, relative to other groups of the immigrant population. Regarding maternal stress, we present evidence suggesting an increase in the share of (immigrant) Muslim women living in affected cities that experience stress after the attacks. Moreover, we find an association between the attacks and stress-induced conditions that may potentially affect health at birth, such as smoking and diabetes. However, we recognize that there might be other potential channels affecting health at birth that we are not able to check due to data constraints. For instance, changes in health care behaviors and labor market results after terrorist attacks or exposure to discrimination. In this line, Johnston and Lordan (2012) find that anti-Muslim sentiment is associated with a decrease in part-time employment. Likewise, it may be the case that Muslim women avoid or skip prenatal care as a result of increased Islamophobia. Due to these limitations, we cannot attribute the negative effect on health at birth exclusively to the increase in maternal stress. Nonetheless, our results are in line with the medical and economic literature that shows a positive relation between stress and detrimental health at birth, as well as research linking Islamophobia and stress.

Our estimates are robust to a set of potential threats allowing us to draw causal conclusions regarding the effects of the 2017 attacks on health at birth. However, there are some other limitations in our study, coming mostly from restrictions in data availability. First, there is limited information regarding the socio-economic status of the household or the specific mother, so that we cannot assess the potential existence of heterogeneous effects. Second, we do not have a direct measure of mother's health status, which can obviously play a role in newborn's health. In spite of these drawbacks we believe our results are robust based on our identification strategy.

Finally, our results are of high policy relevance to encourage future actions towards a more inclusive society, fighting against discrimination towards a group of individuals that not only has a great presence in Spain, 4% of the total population, but also across other European countries as well as in The United States of America.

Appendix 2A: Additional Tables

Table 2A.1: Muslim Immigrants Distribution Across Countries

Country of birth	Freq.	Percent	Cum.	% of population that is Muslim
Algeria	483	1.68	1.68	98.00
Gambia	964	3.36	5.04	95.30
Guinea	260	0.91	5.95	84.20
Guinea-Bissau	41	0.14	6.09	42.80
Mali	427	1.49	7.58	92.40
Morocco	21,338	74.40	81.98	99.90
Mauritian	61	0.21	82.19	99.20
Nigeria	544	1.90	84.09	47.90
Senegal	1,180	4.11	88.20	95.90
Bangladesh	596	2.08	90.28	90.40
Pakistan	2,703	9.42	99.70	96.40
Syria	85	0.30	100.00	92.80
Total	28,682	100.00		

Elaboration: The Authors.
Source: National Registry of live births in Spain (INE, 2018). For the share of Muslim population in each country we refer to the World Population Review (2010). `https://worldpopulationreview.com/countries/muslim-majority-countries/`, and `file:///C:/Users/Grace/Downloads/FutureGlobalMuslimPopulation-WebPDF-Feb10.pdf`.

Table 2A.2: Emigration and Net Immigration of Muslim Women

Panel A: Emigration (from SPAIN)
With destination to countries with a majority of Muslim population
Per semester, years 2015 to 2018.
Women between 15 and 49 years old – All nationalities

	2015-I	2015-II	2016-I	2016-II	2017-I	2017-II	2018-I	2018-II
# Migratory movements	1,855	2,091	2,087	2,236	2,765	2,663	2,094	1,899

Panel B: Emigration (from SPAIN) to foreign countries
Per semester, years 2015 to 2018.
Women between 15 and 49 years old – born in Muslim countries

	2015-I	2015-II	2016-I	2016-II	2017-I	2017-II	2018-I	2018-II
# Migratory movements	3,832	4,315	4,061	4,210	4,256	4,239	3,918	3,446

Panel C: Net immigration (abroad)
Per semester, years 2015 to 2018.
Women between 15 and 49 years old – born in Muslim countries

	2015-I	2015-II	2016-I	2016-II	2017-I	2017-II	2018-I	2018-II
# Migratory movements	1,775	1,814	2,580	3,163	3,419	5,057	6,094	10,573

Elaboration: The Authors.
Source: Emigration and immigration flows taken from the National Institution of Statistics of Spain (INE). The unit of observation is the number of migratory movements.

Table 2A.3: Robustness Check - Measure of Distance to Treated Cities

	(1)	(2)	(3)	(4)	(5)	(6)
	Birth Weight	**LBW**	**Prematurity**	**Complications**	**Female**	**C-Section**
DDD Dist*Muslim*Post	**-73.6575***	**0.0521*****	**0.0041**	**-0.0401**	**0.0588***	**-0.0558**
	(40.0898)	(0.0188)	(0.0130)	(0.0371)	(0.0341)	(0.0424)
Interaction Distance*Post	-20.4512	0.0013	0.0194*	0.0053	0.0029	0.0342
	(30.7420)	(0.0156)	(0.0110)	(0.0201)	(0.0262)	(0.0259)
Interaction Distance*Muslim	13.9920	-0.0040	0.0031	0.0548***	-0.0184	0.0015
	(35.3630)	(0.0148)	(0.0097)	(0.0176)	(0.0261)	(0.0245)
Interaction Muslim*Post	37.8409	-0.0350**	0.0053	0.0351	-0.0340	0.0291
	(32.8891)	(0.0163)	(0.0114)	(0.0288)	(0.0266)	(0.0354)
Mean dependent variable	3,282.19	0.0673	0.0427	0.1129	0.4705	0.2736
Municipality of residence FE	x	x	x	x	x	x
Month, year of conception FE	x	x	x	x	x	x
Mother's country origin FE	x	x	x	x	x	x
Observations	65,519	65,519	75,070	75,070	75,070	75,070

Notes: OLS estimates. Outcomes for the live births from mothers whose region of residence is Catalonia. Standard errors are clustered at the mother's municipality of residence level, 121 clusters. The variable "Distance" is modelled as a negative exponential form of the minimum distance to Barcelona, Cambrils or Ripoll. ***p<0.01, **p<0.05, *p<0.10.

Table 2A.4: Robustness Check - Excluding Barcelona City

	(1)	(2)	(3)	(4)	(5)	(6)
	Birth Weight	**LBW**	**Prematurity**	**Complications**	**Female**	**C-Section**
DDD City*Muslim*Post	**-139.5538*****	**0.0163**	**0.0283**	**0.0312**	**-0.0859***	**0.0605*****
	(11.9017)	(0.0301)	(0.0254)	(0.0469)	(0.0449)	(0.0182)
Interaction City*Post	-6.4718	0.0196	0.0221	-0.0413***	0.0765**	-0.0253
	(12.2431)	(0.0172)	(0.0162)	(0.0103)	(0.0342)	(0.0376)
Interaction City*Muslim	233.7940***	-0.0414***	-0.0423***	-0.0064	-0.0132	-0.0025
	(31.3878)	(0.0034)	(0.0073)	(0.0118)	(0.0196)	(0.0088)
Interaction Muslim*Post	-15.9175	0.0018	0.0079*	0.0013	0.0060	-0.0156*
	(10.5626)	(0.0055)	(0.0042)	(0.0069)	(0.0103)	(0.0090)
Mean dependent variable	3,476.82	0.0336	0.0144	0.1223	0.4820	0.2590
Municipality of residence FE	x	x	x	x	x	x
Month, year of conception FE	x	x	x	x	x	x
Mother's country origin FE	x	x	x	x	x	x
Observations	47,621	47,621	54,582	54,582	54,582	54,582

Notes: OLS estimates of Equation (2.1). Outcomes for the live births from mothers whose region of residence is Catalonia excluding those living in Barcelona City (21,565 observations). Standard errors are clustered at the mother's municipality of residence level, 120 clusters ***p<0.01, **p<0.05, *p<0.10.

Table 2A.5: Robustness Check - Excluding Ripoll City

	(1)	(2)	(3)	(4)	(5)	(6)
	Birth Weight	**LBW**	**Prematurity**	**Complications**	**Female**	**C-Section**
DDD City*Muslim*Post	**-10.2867**	**0.0149*** **	**0.0046**	**0.0134***	**0.0213***	**0.0115**
	(12.0062)	(0.0057)	(0.0042)	(0.0074)	(0.0123)	(0.0091)
Interaction City*Post	-2.8852	-0.0034	0.0087**	-0.0018	-0.0111	0.0152**
	(6.6855)	(0.0047)	(0.0034)	(0.0057)	(0.0067)	(0.0060)
Interaction City*Muslim	3.1384	0.0011	-0.0004	0.0074	-0.0136**	-0.0071
	(13.8459)	(0.0039)	(0.0033)	(0.0046)	(0.0056)	(0.0062)
Interaction Muslim*Post	-15.8402	0.0018	0.0077*	0.0012	0.0052	-0.0154*
	(10.6046)	(0.0056)	(0.0042)	(0.0069)	(0.0103)	(0.0090)
Mean dependent variable	3,281.53	0.0675	0.0429	0.1122	0.4714	0.2726
Municipality of residence FE	x	x	x	x	x	x
Month, year of conception FE	x	x	x	x	x	x
Mother's country origin FE	x	x	x	x	x	x
Observations	65,447	65,447	74,982	74,982	74,982	74,982

Notes: OLS estimates of Equation (2.1). Outcomes for the live births from mothers whose region of residence is Catalonia excluding those living in the municipality of Ripoll (90 observations). Standard errors are clustered at the mother's municipality of residence level, 120 clusters. ***p<0.01, **p<0.05, *p<0.10.

Table 2A.6: Robustness Check - Excluding Cambrils City

	(1)	(2)	(3)	(4)	(5)	(6)
	Birth Weight	**LBW**	**Prematurity**	**Complications**	**Female**	**C-Section**
DDD City*Muslim*Post	**-9.2203**	**0.0163*** **	**0.0053**	**0.0166****	**0.0280*** **	**0.0118**
	(11.4679)	(0.0059)	(0.0045)	(0.0071)	(0.0104)	(0.0091)
Interaction City*Post	-2.8910	-0.0041	0.0081**	-0.0015	-0.0116*	0.0164***
	(6.6735)	(0.0047)	(0.0034)	(0.0057)	(0.0066)	(0.0058)
Interaction City*Muslim	-4.3075	0.0023	0.0002	0.0077	-0.0150**	-0.0079
	(8.1318)	(0.0034)	(0.0031)	(0.0047)	(0.0058)	(0.0062)
Interaction Muslim*Post	-15.8887	0.0019	0.0077*	0.0012	0.0052	-0.0155*
	(10.6016)	(0.0056)	(0.0042)	(0.0069)	(0.0103)	(0.0090)
Mean dependent variable	3,273.16	0.0688	0.0437	0.1131	0.4691	0.2752
Municipality of residence FE	x	x	x	x	x	x
Month, year of conception FE	x	x	x	x	x	x
Mother's country origin FE	x	x	x	x	x	x
Observations	65,180	65,180	74,710	74,710	74,710	74,710

Notes: OLS estimates of Equation (2.1). Outcomes for the live births from mothers whose region of residence is Catalonia excluding those living in the municipality of Cambrils (372 observations). Standard errors are clustered at the mother's municipality of residence level, 120 clusters. ***p<0.01, **p<0.05, *p<0.10.

Table 2A.7: Placebo Outcome

	Rejection Latin American
DD Province*Post	**-0.0033**
	(0.0066)
Treated Provinces	0.0266***
	(0.0024)
Post (September 2017)	0.0156***
	(0.0051)
Mean dependent variable	0.0267
Province FE	x
Year FE	x
Observations	6,959

Notes: The outcome variable is made up by the rejection against immigrants coming from Latin American countries. Robust standard errors in parentheses clustered at the province level (50 clusters). ***p<0.01, **p<0.05, *p<0.10.

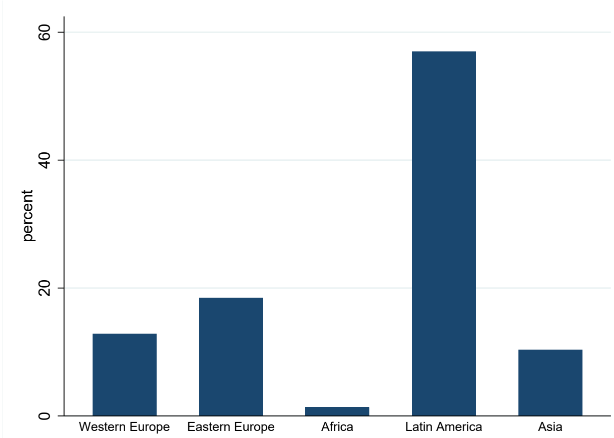
Table 2A.8: Terrorist Attacks and Stress - Placebo

	(1)	**(2)**	**(3)**	**(4)**	**(5)**
	Bad Health	**Relaxed**	**Self-confident**	**Happiness**	**Emotional Well-being Index**
DDD City*Latin*Post	**-0.1849**	**-0.0589**	**0.2478****	**0.2274**	**0.6324**
	(0.1144)	(0.2886)	(0.0980)	(0.1536)	(0.3805)
Interaction City*Post	0.0096	-0.0197	-0.1028	-0.2470**	-0.5051
	(0.0986)	(0.2411)	(0.0972)	(0.1053)	(0.3148)
Interaction City*Latin	0.1679	-0.1288	-0.1031	-0.1676	-0.4756
	(0.1013)	(0.1425)	(0.0965)	(0.1022)	(0.2964)
Interaction Latin*Post	0.1564*	-0.1336	-0.1966***	-0.2182***	-0.6855***
	(0.0833)	(0.1202)	(0.0550)	(0.0621)	(0.1963)
Mean dependent variable	0.1573	0.5287	0.8506	0.8621	-0.1544
Month FE	x	x	x	x	x
Year FE	x	x	x	x	x
Municipality (AGA) FE	x	x	x	x	x
Mother's country of birth FE	x	x	x	x	x
Observations	686	665	665	665	663

Notes: Sample includes years 2015 to 2018 of the Catalan Health Interview Survey provided by the Catalan Health Department. Women of childbearing age (15-45 years old). The treatment group is made up with mothers born in Latin American countries. Controls are mothers born in other foreign countries. Estimation using weights according to survey design. Standard errors clustered at the "Area de region sanitaria" in Catalonia level, 43 clusters. ***p<0.01, **p<0.05, *p<0.10.

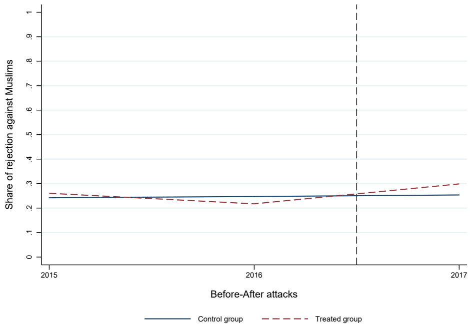
Appendix 2B: Additional Figures

Figure 2B.1: Non-Muslim Immigrants Distribution across Countries

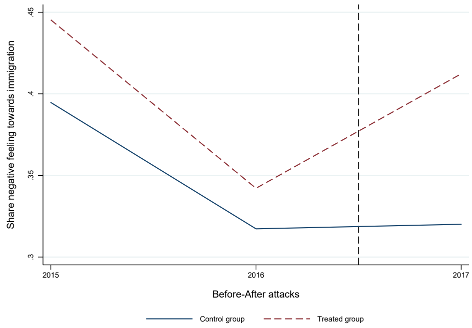


Notes: Own elaboration using the National registry of live births in Spain provided by the INE, years 2015 - 2018. Every bar represents the percent that represents each region within the control group

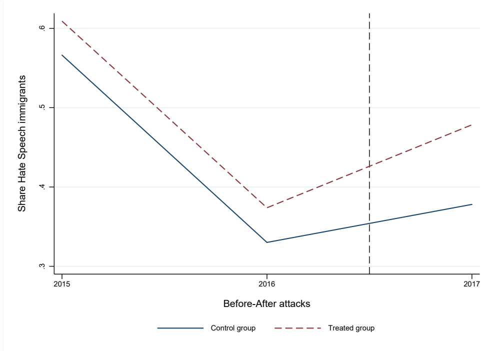
Figure 2B.2: Raw Trends for Measures of Islamophobia



(a) Rejection against Muslims



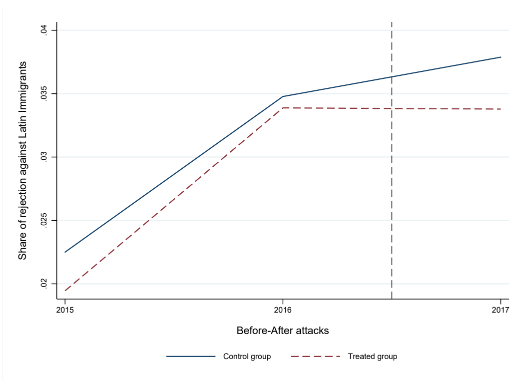
(b) Negative feeling immigration



(c) Hate speech

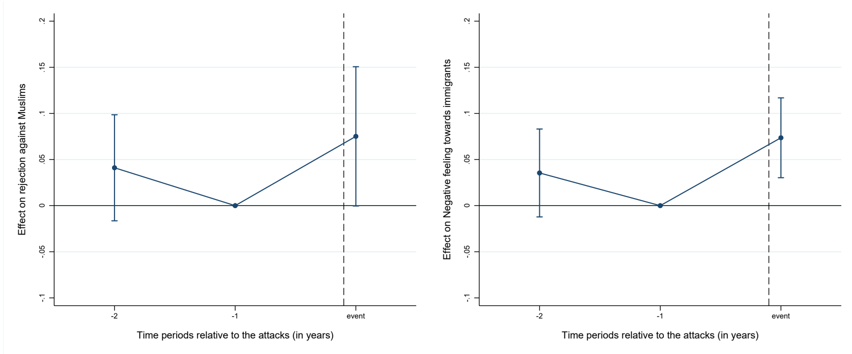
Notes: Figure shows the raw trends of the Islamophobia-related outcomes. Panel (a) depicts trends for the rejection against Muslim, panel (b) for the negative feeling towards immigrants, and panel (c) for hate speech against immigrants, in all cases comparing affected vs non-affected provinces. Authors' calculations based on data from the "Centro de Investigaciones Sociológicas" of Spain, for years 2015-2017.

Figure 2B.3: Raw Trends Rejection against Latin American

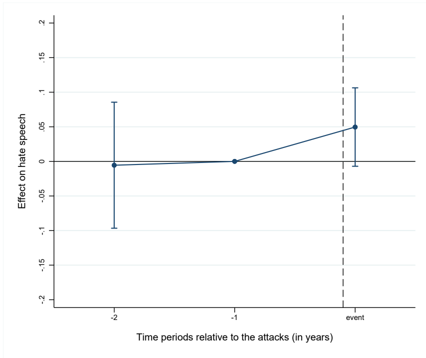


Notes: Figure shows the raw trends of the rejection against Latin American comparing affected vs non-affected provinces. Authors' calculations based on data from the "Centro de Investigaciones Sociológicas" of Spain, for years 2015-2017.

Figure 2B.4: Event Study Graphs of Rejection against Immigrants



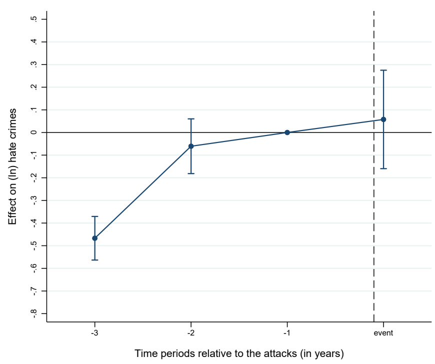
(a) Rejection against Muslims (b) Negative feeling immigration



(c) Hate speech

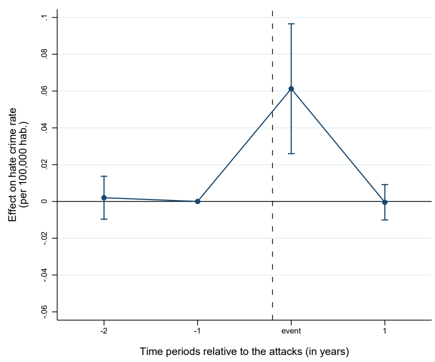
Notes: Figure shows the set of coefficients corresponding to the difference in differences terms from the event study version of Equation (2.2). Time periods are defined on an annual basis. Point estimates are shown along with their 95 percent confidence intervals. Reference category is one period prior to the event date. Standard errors are clustered at the province level, 50 clusters. Authors' calculations based on data from the "Centro de Investigaciones Sociológicas" of Spain, for years 2015-2017.

Figure 2B.5: Hate Crimes - Event Study



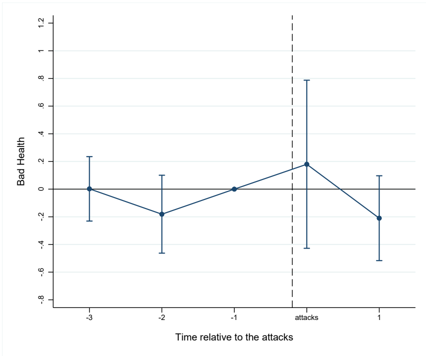
Notes: Figure shows the set of coefficients corresponding to the difference in differences terms from the event study version of Equation (2.3). Sample includes years 2015 to 2017. Time periods are defined on an annual basis. Point estimates are shown along with their 95 percent confidence intervals. Reference category is one period prior to the event date. Standard errors clustered at the municipality level, 203 clusters. Authors' calculations based on data from Catalan Police Department.

Figure 2B.6: Hate Crimes and Nationality of the Victim - Event Study

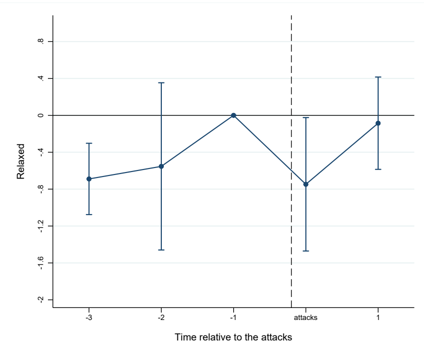


Notes: Figure shows the set of coefficients corresponding to the difference in differences terms from the event study version of Equation (2.4). Sample includes years 2015 to 2018. Time periods are defined on an annual basis. Point estimates are shown along with their 95 percent confidence intervals. Reference category is one period prior to the event date. Standard errors clustered at the province level, 52 clusters. Authors' calculations based on data from Catalan Police Department.

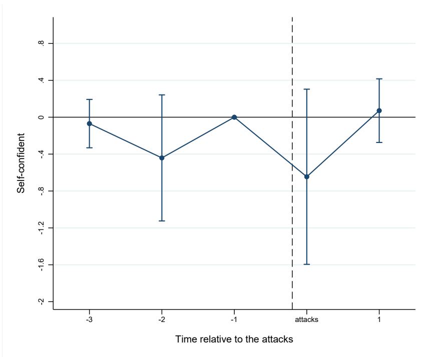
Figure 2B.7: Terrorist Attacks and Stress - Event Study Graphs



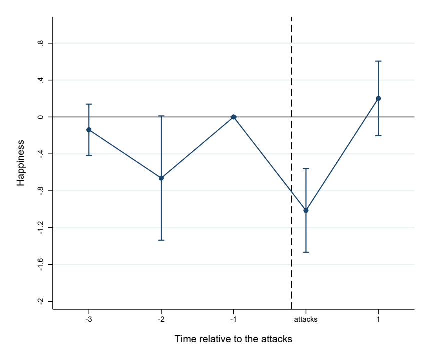
(a) Bad Health



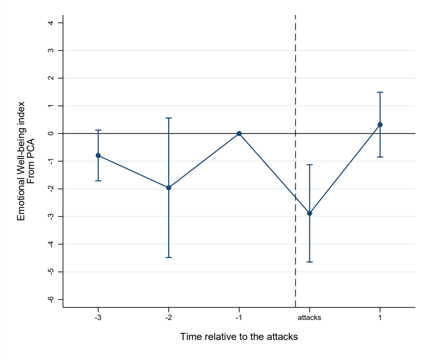
(b) Relaxed



(c) Self-confidence



(d) Happiness



(e) Emotional well-being index

Notes: Figure shows the set of coefficients corresponding to the triple interaction term from the event study versions of Equation (2.1) using as outcomes the stress markers shown in Table 2.9. Time periods are defined on an annual basis. Point estimates are shown along with their 95 percent confidence intervals. Reference category is one period prior to the event date. Sample includes years 2015 to 2018. Immigrant women of childbearing age (15-45 years old). Estimation using weights according to survey design. Standard errors are clustered at the "Area de region sanitaria" in Catalonia level, 43 clusters. Authors' calculations based on data from Catalan Health Department, Catalan Health Interview Survey.

3 Drug Tenancy Thresholds, Consumption and Newborn's Health

3.1 Introduction

The use/abuse of substances that generate dependency has become a severe worldwide health problem (World Health Organization, 2021). According to the World Health Organization, during year 2020, around 275 million people, which represents around 6% of the world's population between 15-64 years old, have used an illicit drug at least once in the previous year. Moreover, a non-negligible number of people (around 36 million, 13% of drug users) suffer from a drug use disorder or dependency becoming it in a serious public health issue (World Health Organization, 2021). Health effects of drug use are well documented, for instance, neurological impairments and increased risks of certain infectious diseases, among others (UNODC, 2018). Drug consumption among women of child bearing age is particularly problematic because drugs may cross the placenta and damage the fetus. In this chapter, I explore the link between maternal drug use and newborn's health, by focusing on a policy change in Ecuador aimed at decriminalizing drug consumers, which seems to have contributed to an increase in drug use.

The medical literature shows that there is an association between drug use during pregnancy and poor health at birth. For example, intrauterine exposure to drugs is linked to low birth weight, very low birth weight, preterm births, congenital malformations, feeding problems with the newborn, higher probabilities of being admitted to the neonatal intensive care unit, and higher child mortality (Grant et al., 2018; Srikartika and O'leary, 2015; Simmons and Austin, 2022; Haight et al., 2018; Finnegan et al., 1977; Creanga et al., 2012; Umer et al., 2022; Bailey and Diaz-Barbosa, 2018; Oro and Dixon, 1987; Behnke et al., 2013).

Given the socioeconomic burden behind drugs, countries have implemented several actions to deal with them, ranging from communication campaigns to marijuana liberalization policies. Ecuador, a South American country, offers and excellent case study. In year 2013, the government to recognize the health needs of drug consumers, as well as, to distinguish them from traffickers, enacted a resolution that decriminalize drug consumption using as means a "table" that established the maximum permissible amounts of possession for personal consumption of substances subject to control. This table contains thresholds of drug tenancy for marijuana, cocaine, heroin, and other synthetic drugs. The policy does not make illicit drugs to be legal but allows the police force to objectively distinguish between a consumer and a trafficker. Data from the National Survey on Drug Use and Consumption (2016) shows that annual prevalence of marijuana consumption among teenagers was around 9.6%, for cocaine, base paste and heroin was 2.3%, 2% and 2.5%, respectively, with evidence of a dynamic of territorial agglomeration (Secretaría Técnica de Prevención Integral de Drogas, 2017). Likewise, the survey shows that average age of first drug use (any illicit drug) is around 14.62 years old.

The aim of this chapter is to analyze whether the implementation of the table of maximum permissible amounts of possession for personal consumption of substances subject to control in Ecuador has an impact on newborn's health, through an increase in drug consumption among women of reproductive age. For this, I start the analysis by creating a measure of geographical exposure to drugs at the provincial level, which generates provinces with higher and lower intensities in terms of drug exposure. Then, I exploit the regulatory change explained in the previous paragraph. I use the geographic variation in drug exposure in a difference-in-differences framework to compare, before and after 2013, provinces in Ecuador that are more exposed to drugs to those less exposed.

A key assumption in any difference-in-differences setting is the parallel trends assumption. In this setting it means that provinces with a lower treatment intensity are a good counterfactual for provinces with a higher treatment intensity. To add evidence supporting this assumption, I perform a set of robustness checks in the following sections.

To measure health at birth, I use administrative data provided by the official agency in charge of the statistics in the country for two years before and two years after the policy change. I explore changes in average birth weight, low birth weight, very low birth weight and preterm births. I find that provinces with a higher drug exposure have a higher share of newborns with low birth weight and very low birth weight. I also find that average birth weight is less in more exposed provinces. For example, increasing the initial exposure to drugs by 1 percentage points decreases average birth weight by 33.35 g. It also increases the share of low birth weight live births by 1.43 percentage points. This is equivalent to an increase of approximately 16.26% in the share of low birth weight when compared to the baseline.

The results also show that newborns living in more exposed provinces have higher rates of neonatal hospitalizations. For instance, I find an increase in the hospitalization rate (per 1,000 live births) of around 17.72 for conditions related to maternal drug use. In addition, these neonates are expected to have a rate 1.77 times greater for drug-related hospitalizations. However, I do not find any significant difference in the neonatal hospitalization mortality rate.

I also find that the detrimental health at birth outcomes observed is very likely to come from increased women's drug use. Using data from hospital discharges and outpatient visits, I find that women of reproductive age living in more exposed provinces show higher drug-related hospitalization rates. For instance, an increase of 1 pp in the initial drug intensity exposure is linked to an increase of around 0.0885 drug-related hospitalizations per 10,000 women. On the contrary, I do not find any significant difference in prenatal care use or in women's fertility.

This chapter contributes to a growing literature evaluating the effect of drug decriminalization policies on health-related outcomes. On this area, research is restricted to marijuana liberalization policies and to the opioid crisis in the U.S. For instance, Williams and Bretteville-Jensen (2014) using data from Australia, found that the impact of decriminalization, in terms of increases in cannabis use, is concentrated amongst minors. Laqueur et al. (2020) find no evidence of an impact on cannabis use or on the perceived risk of use among Uruguayan teenagers. Kim (2021) find strong evidence supporting that policies that limit access to legal opioids in the U.S. have increased heroin death rates. Wen et al. (2015) find that medical marijuana laws in the U.S. increase the probability of current marijuana use, regular marijuana use and marijuana abuse among those aged 21 or above.

More closely related to this chapter, Meinhofer et al. (2021) study how marijuana liberalization policies affect perinatal health in the U.S. The authors find an increase in maternal marijuana use but no effect on newborn's health. My work is distinguishable in that: (i) I exploit a policy targeted to multiple types of illicit drugs whose aim was not to make drug consumption legal, and (ii) in a developing country setting which, to my knowledge, has not been studied before. (iii) The methodological approach is different as I evaluate a national policy exploiting geographical variation on pre-existing drug exposure intensities. Finally, I make use of highly reliable administrative data on healthcare use, which allows me to avoid issues that may arise in survey data.

This chapter also contributes to the literature on the determinants of child and adulthood outcomes. Health at birth has been found to be a significant determinant of child development, education attainment, and future earnings (Almond et al., 2005; Lindeboom et al., 2010; Almond and Currie, 2011; Scholte et al., 2015).

Assessing the effects of drug use is important not only under the consumers' health perspective, but also because of potential negative spillovers on the society.

For instance, it may affect local crime, in which drug circulation might play an important role (Dell, 2015; Adda et al., 2014; Gavrilova et al., 2019; Lindo and Padilla-Romo, 2018). Drug use may have significant negative impacts on educational performance (Michaelsen and Salardi, 2020). In addition, drug consumption is also associated with risky behaviors, especially among younger groups. On this line, according to Rashad and Kaestner (2004), substance use is positively associated with adolescent sexual behaviors such as early initiation of sexual intercourse, and multiple sexual partners. Finally, all these undesirable consequences may be transferred to future populations perpetuating the negative effects.

The rest of the chapter proceeds as follows. Section 3.2 provides the context of the decriminalization policy in Ecuador and the link between maternal drug use and newborn's health. Section 3.3 describes the data sets and the computation of the geographical measure of drug intensity. Section 3.4 presents the identification and empirical strategy. In Section 3.5, I present the results on health at birth and women's substance use. In Section 3.6, I present a battery of robustness checks. Finally, Section 3.7 concludes with a discussion of the preliminary results.

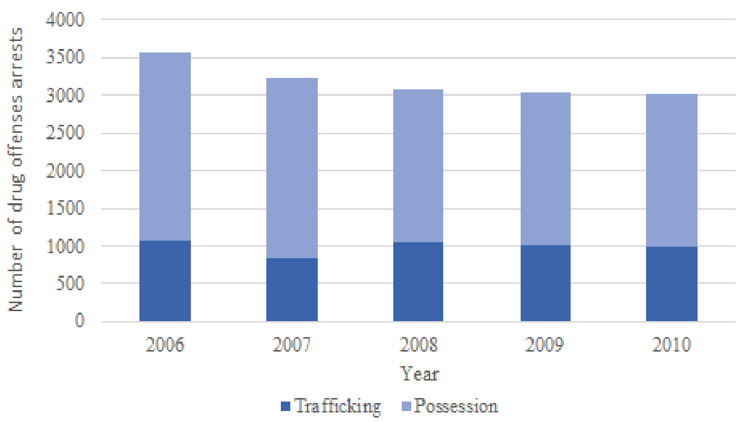
3.2 Background

3.2.1 Drug Consumption in Ecuador

Ecuador is a country located in South America in which illicit drugs are not legal. In year 2008, under the new Constitution of the Republic of Ecuador, addictions are declared as a public health issue. The new law established that the State is responsible for developing coordinated programs of information, prevention to control the consumption of alcohol, tobacco and narcotic and psychotropic substances. In addition, the State is also responsible for providing treatment and rehabilitation to occasional, regular and problematic users. Likewise, it was established that consumers may not be criminalized and that they should be considered as in need of medical support (Asamblea Constituyente Del Ecuador, 2008).

In spite of this, the enforcement of drug consumers' rights, as established in the Constitution, was not applied as there was not any legal document that set the parameters to objectively distinguish a consumer from a trafficker. This argument is in line with the numbers of arrests involving illicit drugs. Figure 3.1 shows drug offenses arrests from year 2006 to 2010 at the national level. We can see that, overall, the proportion corresponding to drug possession is larger than those of trafficking for all years. For instance, for year 2010, possession arrests accounted for almost the 70% of all drug offenses. Likewise, for year 2014, there were around 25 thousand people incarcerated, from which approximately a 26% corresponds to drug offenses.

Figure 3.1: Drug Offenses Arrests - Ecuador



Notes: Own elaboration using data from the National Antinarcotics Office of the National Police Institution.

It was not until year 2013 when the government through the National Drug Secretary[1] to recognize the health needs of drug consumers, as well as, to distinguish them from traffickers, issued a resolution that established the maximum permissible amounts of possession for personal consumption of substances subject to control, popularly known as "the table". This table contains thresholds of drug tenancy for marijuana, cocaine, heroin, and other synthetic drugs. Thus, the objective of the "table" is to guide judges so as not to criminalize people who use drugs, complying with the provisions of the Constitution regarding the non-criminalization of consumption. In this setting, it is important to emphasize that the policy does not make illicit drugs to be legal in the country but allows the police force to objectively distinguish between a consumer and a trafficker. Table 3.1 shows the thresholds that were approved in year 2013.

Table 3.1: Table of Maximum Permissible Amounts of Possession for Personal Consumption

Substance	**Quantity (in grams)**
Marijuana	10
Cocaine base paste	2
Cocaine hydrochloride	1
Heroin	0.10
MDA	0.05
MDMA (Extasis)	0.02
Amphetamines	0.04

Elaboration: The Author.
Source: Consejo Nacional de Sustancias Estupefacientes y psicotrópicas (CONSEP) Resolución 001 CONSEP – CD – 2013.

The resolution by which the drug consumption table was issued did not establish any parameters in terms of the frequency with which a person can carry these amounts without being criminalized, nor on combinations of drug quantities.

Consumption of substances that generate dependency is a socioeconomic phenomenon that has become an important public health issue in Ecuador. According to Naranjo and Jaramillo (2017) for year 2015, the estimated annual cost of the drug problem was around US$ 214,029,337.75, which corresponds to US$ 13 per capita,[2] or 0,2% of the gross domestic product. Panel (a) of Figure 3C.1 in Appendix 3C provides some descriptive evidence of how after year 2013 the number of hospital discharges (individuals between 15-45 years old) related to illicit drug use increases. Likewise, panel (b) shows the trends of drug-related hospitalizations

[1] This institution, formerly called "Consejo Nacional de Control de Sustancias Estupefacientes y Psicotrópicas", was eliminated and merged with the Ministry of Health in year 2018.

[2] Based on a total population of 16.144.363 inhabitants.

in neonates,[3] again after year 2013, there is an important increase in absolute numbers.

But how could a policy that is not intended to criminalize drug users affect drug use? There are several factors that may influence the initiation into harmful substance use. The United Nations Office on Drugs and Crime (UNODC, 2018) emphasizes that drug availability as well as the perception of risk of drug use play a key role. On this, an important channel through which the policy might induce an increase in drug use is the reduction in the perception of risk of drug use and a potential increase in supply (ease of access). To measure changes in risk perception is difficult, mainly due to the lack of comparable surveys assessing this dimension. However, an increase in drug consumption should be reflected in some outcomes such as criminal activity and health care use. In the upcoming sections, I provide evidence supporting an increase in health care use related to drug use.

3.2.2 Drug Use during Pregnancy and Newborn's Health

The underlying idea is that the policy implemented to facilitate the distinction between traffickers and consumers may have favored the increase in drug use. The increase in drug use in women of reproductive age may be harmful for the developing fetus.

Maternal substance use is associated with adverse perinatal health outcomes, which includes small for gestational age (Srikartika and O'leary, 2015), preterm births, low birth weight (Bailey and Diaz-Barbosa, 2018), congenital malformations (O'Leary et al., 2012), complications during childbirth (Haight et al., 2018), among others. In addition, all these perinatal health conditions may also increase the likelihood of admissions in the neonatal intensive care unit.

Almost all drugs that generate dependency have the capacity to easily cross the placenta during pregnancy, which exposes the fetus to these substances (Grant et al., 2018; van Hoogdalem et al., 2021). The transport of these substances, from mother to fetus, usually occurs around the fifth week of embryonic life. However, even before the placenta is formed, any substance can negatively affect the development of fetus' organs.

Negative effects of maternal drug use may vary depending on the type of substance that the pregnant woman consumes and the timing of consumption (Behnke et al., 2013). For instance, marijuana use may be linked to cognitive problems of newborn (Meinhofer et al., 2021). Cocaine use is associated with miscarriages, preterm births, premature placental abruptions, stillbirths and low birth weight (Oro

[3] In Section 3.3, I provide a detailed explanation of how the variable drug-related hospitalizations is constructed.

and Dixon, 1987). Methamphetamines may be associated with stillbirths and preterm births, feeding problems, irritability. Heroin and other opiates may cause the abstinence syndrome in neonates, feeding problems, prematurity and congenital malformations (Haight et al., 2018).

Drug use during pregnancy may affect newborn's health not only through the mother-fetus transmission of the substance itself, but also through maternal health and behaviors. For example, women using cocaine are more likely to suffer from anorexia and other feeding disorders. In addition, drug use is associated with an increase in risky behaviours, which may increase the probabilities of infections that can be transmitted to the fetus (Rashad and Kaestner, 2004).

Finally, drug use during pregnancy might also have long-term impacts (beyond health) on children. Behnke et al. (2013) find an increase in the risk of developing behavioral problems, impaired cognition and poorer performance in school.

3.3 Data and Variable Definitions

3.3.1 The Geographical Drug Exposure Index

To assess the impact of maternal drug consumption on newborn's health, I exploit the timing and geographic variation of the intensity of drug exposure across provinces before the policy implementation. To formally calculate this intensity, I develop a composite index at the province level for year 2012, which I will call from now on as "the *index*".

To calculate the *index*, I partially follow Secretaría Técnica de Prevención Integral de Drogas (2017) who designed the "Drug Territorial Concentration Index" (Índice de Concentración Territorial de Drogas, in Spanish) for the year 2016 in Ecuador. In this work, the national authorities developed an index whose aim was to identify the highest incidence territories for policy design. This index considers four dimensions including drug supply and demand, population size and socioeconomic conditions.[4] Due to lack of data before year 2014 for some of the individual indicators proposed by the official authorities, I use other alternative variables. In addition, the methodology used to compute this index has not been publicly released.

To compute the *index*, I use several individual indicators obtained from three different data sources. First, I use information from the IV National (Ecuadorian) Survey of Drug Consumption (2012) from which I generate the annual prevalence of

[4] To have more information regarding the index constructed by the national authorities, the reader can enter: `http://sisco.copolad.eu/web/uploads/documentos/PLAN_NACIONAL_DE_PREVENCION_INTEGRAL_Y_CONTROL_DEL_FENOMENO_SOCIO_ECONOMICO_DE_LAS_DROGAS_2017-2021.pdf`

(illicit) drug use, and the frequency of its use. Second, using administrative records provided by the National Institute of Statistics (INEC) on the universe of hospital discharges for year 2012, I calculate the drug-related hospitalization rate. Third, I use data provided by The National Police of Ecuador on the universe of arrests, and calculate the drug trafficking arrest rate for year 2012. All individual indicators are weighted by the exposed population.

Smith et al. (2015) remark that to combine individual indicators into a single index, one needs to keep in mind that there might be some problems associated with the accurate construction of the index: the variables (i) are measured on different scales, (ii) have different distributions, (iii) may or may not apply to the same individual, and (iv) measure, to different degrees, the underlying idea of the index. To overcome these issues, factor analysis may be a good alternative to obtain weights for the set of individual indicators in order to combine them into a single index (Commission, 2008; Smith et al., 2015). Therefore, I use the common factor analysis Maximum Likelihood technique to generate the *index*.

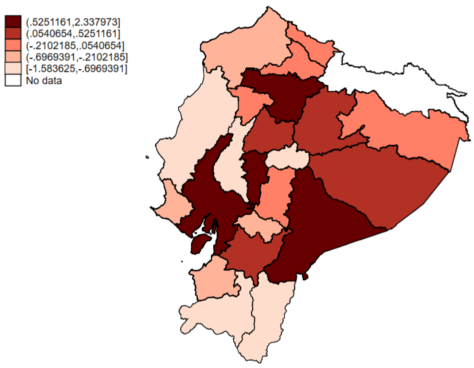
Because the purpose of this *index* is to use indicators that are individually representative of the drug exposure per province, the analysis should focus on exploring a one-common factor model. The idea behind a one-common-factor model, in this setting, is that there is an underlying factor at the province level that makes these four different measures likely to exist together in the same province. Therefore, the indicators measure, with different levels of accuracy, the underlying factor. The methodology assumes that although the measurement is imperfect, the indicators that are most highly correlated with the underlying factor will also be highly correlated with the other indicators. Factor analysis generates a set of weights for the individual indicators, which are combined to make the factor score, or, in other words, the *index*. During the computation of the *index*, I did not find any significant second common factor. I tested the existence of a second common factor using the Eigen values, which is the standard criteria in this methodology.

The *index* is a continuous number that ranges from -1.5836 to 2.3379 where higher values denote stronger drug exposures. [5] The *index* varies significantly across provinces as shown in Figure 3.2. For instance, provinces to both, the northern and southern borders are in the third and second quintiles of the distribution, respectively. Provinces such as Guayas, Pichincha, Bolívar and Morona Santiago, are the "top" in terms of higher values of the *index*.

On the contrary, Tungurahua, Loja and Zamora are at the bottom of the drug exposure index distribution. Details on the computation of the index including the geographical distribution of the individual indicators are shown in Appendix 3A.

[5] For ease of interpretation in the estimates, I re-scale the index between zero and one.

Figure 3.2: Drug Exposure Index (2012) - Ecuador



Notes: Figure shows geographic variation at the province level for the drug exposure *index*. Light colors correspond to lower levels of drug exposure. The Galapagos Islands are excluded from the analysis because there is no information on the indicators to build the *index*.

3.3.2 Main Data Sources

This chapter uses data from four administrative sources. National Births Records data and Neonatal Hospital Discharges provide the outcome variables for newborn's health. For women drug use, I use data from hospital discharges and from outpatient visits.

Birth Records

I use data on the universe of live births between the years 2011 to 2015 provided by the INEC. In this dataset the unit of observation is the live birth. I consider live births to women between 13 and 45 years old. I also exclude observations with birth weight less than 500 g and multiple births.

For every live birth, there is information on weight (in grams), date of birth, gender, type of delivery and gestational length (in weeks). From here, I construct the main outcomes of interest: average birth weight (in grams); an indicator of low birth weight (LBW) that takes the value of 1 for weight below 2500 g, and 0 otherwise; very low birth weight (VLBW), which takes the value of 1 if the weight is below 1,500 g, and 0 otherwise. I also consider prematurity with a dummy variable that takes the value of 1 when the delivery occurs before the 37th week of gestation. I follow Balsa and Triunfo (2022) and distinguish between LBW and VLBW because VLBW births represent an important proportion of newborn's health care costs. In addition, there is medical literature showing that prenatal substance abuse is associated with extreme outcomes such as very low birth weight (Behnke et al., 2013; Oro and Dixon, 1987). Finally, I also consider the type of delivery with a dummy variable that takes the value of 1 for cesarean sections and 0 otherwise. The database also has information on mother's characteristics such as age and province of residence. I use mother's province of residence to match with the drug intensity exposure index.

To classify live births into exposed and non-exposed to the policy implementation, I use reported gestational length and the date of birth to calculate the (estimated) conception date. For this, I follow the approach employed by Quintana-Domeque and Ródenas-Serrano (2017) and convert gestational age (in weeks) into months dividing by 4. Then, I substract this number from the month of birth to estimate the conception date.

Following common practice in the health economics literature, I conduct the analysis around the conception date because the actual date of birth may be endogenous. With this definition, the sample is made up of 943,237 live births conceived between April 2010 and March 2015.

Neonatal Hospital Discharge Data

To measure health outcomes during the neonatal period, I use hospital discharge records from years 2011 to 2015 provided by the INEC. This dataset comprises the universe of overnight hospitalizations of neonates (less than 29 days of birth) in public and private hospitals. The data includes patient's length of stay, date of admission and discharge, main diagnosis (ICD-10 diagnostic code), condition of discharge (death or alive), patient's age (in days), date of birth, gender and province of residence.

To generate the neonatal health outcomes, I consider only conditions originated in the perinatal period and congenital malformations. As with the birth records data, I conduct the analysis around the estimated conception date,[6] which sums up 93,147 hospital discharges. The outcomes of interest are: (i) the neonatal hospitalization rate, which is the total number of hospitalizations per 1,000 live births; (ii) the neonatal hospital death rate per 1,000 live births and (iii) the drug-related neonatal hospitalization rate per 1,000 live births.

To build the last outcome, I follow the work done by Meinhofer et al. (2021) who based their election on literature that establishes an association between maternal drug use and detrimental neonatal health (Behnke et al., 2013). These conditions

[6]In this dataset there is no information about gestational length. Therefore, to calculate the estimated conception date, I assume all cases correspond to pregnancies of 38 weeks. I support this assumption on the results of column 4 of Table 3.3, which shows no statistically significant coefficient for pre-term births.

include, for example, exposure to noxious substances via placenta or breast milk, neonatal withdrawal symptoms from maternal use of drugs, slow fetal growth, feeding problems, congenital malformations, neonatal jaundice due to drugs transmitted from mother, and disorders related to low birth weight. Table 3B.1 in the Appendix 3B details the conditions along with their corresponding ICD-10 diagnostic codes.

I compute all outcomes by collapsing hospital discharges at the month, year of conception and neonate's province of residence level. The unit of observation is the number of hospitalizations (drug-related hospitalizations) in each year, month of conception and province of residence. Then, to calculate the rates, I divide each observation by the number of live births in the corresponding province and month-year. I obtain the number of live births from the Birth Registries published by the INEC. Finally, I merge this data with the drug exposure index, which is also at the province level.

Adults Health Care Use Data

The underlying idea is that newborn's health might be negatively affected as a consequence of an increase in the use/abuse of drugs by mothers. To examine changes in consumption, I use two dataset: (i) the universe of hospital discharges from years 2011 to 2015 and (ii) outpatient health care use from years 2010 to 2015. Both data sets are provided by the INEC.

The first dataset comprises all hospitalizations across the country in both private and public institutions. In here, the unit of observation is an individual's hospital discharge. For every unit of observation, there is information at the patient's level on sex, age, and province of residence. Healthcare information includes the date of occurrence, number of days of hospital stay, and the primary ICD-10 diagnostic and procedure codes associated with the discharge. I use patient's province of residence to match observations with the *index* and other variables at the province level. From here, I generate a female sample, which includes women between 13 and 45 years old. The outcomes of interest are: (i) the drug-hospitalization rate per 10,000 at risk population,[7] (ii) the proportion of drug related hospitalizations (relative to the total number of hospitalizations),[8] and (iii) the number of complications during childbirth per 1,000 live births. I define a drug-related hospitalization according to the ICD-10 diagnostic codes. All conditions involving any substance use or abuse that might cause dependency are considered. A detailed list with these conditions is shown in Table 3B.2 in the Appendix 3B.

[7]To calculate this rate, I use census data on the population per province and year provided by the Ecuadorian National Institute of Statistics (INEC).

[8]For this outcome, I exclude from the denominator the hospitalizations referring to natural childbirth without complications.

With respect to outpatient health care use, the dataset includes the universe of visits provided by mental health care practitioners of public institutions. The unit of observation is the health care center, for which there is information on the number of visits and the province where the medical unit is located at the year level. As with the other samples, I restrict the analysis to women between 15 and 49 years old. The outcomes of interest are the total number of visits, and the visits rate per 10,000 at risk population. I collapse the data at the province and year level and merge it with the *index* data.

3.3.3 Descriptive Statistics

Table 3.2 presents descriptive statistics for the outcome variables at the province, month-year level, except for the outpatient visits, which is at the year level. Panel A shows the mean and standard deviation for newborn's health outcomes. Average birth weight is around 3,089 g, the share of low birth weight newborns is around 10%, the share of very low birth weight is approximately 1%. A 6.23% of deliveries are before 37 weeks of gestation The proportion of c-sections is 44.87%.

With respect to the neonatal hospitalization data, the hospitalization rate for all medical conditions originated in the perinatal period and congenital malformations is 94.67 per 1,000 live births per province. The hospitalization rate for conditions related to maternal use of drugs of addiction is 34.39 per 1,000 live births. Likewise, the neonatal hospital death rate is 4.3042 per 1,000 live births.

Panel B shows summary statistics for women's hospitalizations caused by conditions related to substance use. The hospitalization rate per 10,000 women of childbearing age is 0.1271. The proportion of substance use hospitalizations with respect to all-causes hospitalizations is 0.36%. Regarding outpatient visits for causes related to mental and behavioral disorders, the yearly rate per 10,000 women per province is 277.19.

In Panel C, I present descriptive information regarding mother's characteristics derived from the lived births data. Average mother's age is 25.67 years old. The share of low educated mothers, defined as those with less than completed secondary education, is 32.72%. Finally, the share of single mothers is approximately 20%.

Table 3.2: Descriptive Statistics

	Observations	**Mean**	**Std.Dev**
Panel A: Newborn's health outcomes			
Live births data			
Birth weight (g)	1,320	3,089.611	69.9375
Low birth weight	1,320	0.0994	0.0236
Very low birth weight	1,320	0.0079	0.0055
Preterm	1,320	0.0623	0.0201
C-section	1,320	0.4487	0.1848
Neonatal hospitalization data			
Hospitalization rate per 1,000 live births (all causes)	1,320	94.6708	58.8937
Drug-related hospitalization rate per 1,000 live births	1,320	34.3915	29.9188
(#) Drug-related hospitalizations	1,320	73.6087	71.6296
Neonatal hospital death rate per 1,000 live births	1,320	4.3042	3.8804
Panel B: Women's substance use			
Hospital discharge data			
Substance use hospitalization rate per 10,000 women	1.320	0.1271	0.1126
Proportion of substance use hospitalizations	1,320	0.0036	0.0029
Complications childbirth rate per 1,000 live births	1,320	0.1436	0.0891
Mental health outpatient visits data			
Outpatient visit rate per 10,000 women	132	277.1929	161.6139
Panel C: Mother's characterization			
Live births data			
Average mother's age	1,320	25.6767	0.7574
Share of low educated mothers	1,320	0.3272	0.0614
Single mother indicator	1,320	0.1826	0.0816

Notes: All variables show provincial averages at the month-year level, except for the outpatient visit rate, which is at the year level. A mother is considered as low educated if her higher degree is less than secondary education. All variables are generated from the dataset described in Section 3.3.

3.4 Empirical Strategy

To assess whether maternal drug use negatively affects newborn's health, I exploit the timing and geographic variation of the intensity of drug exposure across provinces. For this, I use the *index* to account for differences across provinces and compare them before versus after the enacting of the policy.

Newborn's exposure to maternal drug use is determined by month-year of conception and the mother's province of residence. Because the policy was enacted in year 2013, all newborns conceived from April 2012 onwards are therefore exposed (to some degree). In this setting, I compare newborns conceived before April 2012 (non-affected), with live births conceived after April 2012 (affected).

The policy affected all provinces, however the initial drug exposure intensity induced by the policy varies across provinces because some provinces were, from before, more exposed to drugs than others. To capture these different intensities, I use the geographical index developed in Section 3.3.1. The *index* is continuous, therefore, the treatment variable is also continuous, which exploits a different treatment intensity across provinces. Using a continuous over a dichotomous measure has some advantages. First, I can make use of all the variation in the drug *index*, which is less restrictive and does not impose to considered treated units (in a lower degree) as untreated units. Second, one does not need to rely on arbitrary assumptions to categorize a group of observations as treated or as untreated (Altman and Royston, 2006; Schmidheiny and Siegloch, 2022).

The baseline specification is a difference in differences (DiD) at the province level, which exploits the different intensity in drug exposure before versus after the policy implementation. I collapse both live births and neonatal hospitalizations data at the province, month and year of conception, and estimate by Ordinary Least Squares (OLS) an equation of the form:

$$Y_{pmt} = \alpha + \beta_1 Index_p * Post_{mt} + \beta_2 X'_{pmt} + \delta_p + \rho_m + \gamma_t + \epsilon_{pmt}, \tag{3.1}$$

where Y_{pmt}, is a newborn's health outcome in province p, month m, year t. $Index_p$ is a continuous variable that measures the drug exposure in province p before the policy implementation. $Post_{mt}$ is a dummy that takes the value of one for live births (conception) affected by the policy and zero for those unaffected.

X'_{pmt} represents a set of control variables. For regressions on health at birth it corresponds to average mother's age in province p, month m, year t. For regressions on neonatal hospitalization data it accounts for average neonate's age (in days) and gender. The parameter δ_p represents 22 provinces fixed effects, ρ_m and γ_t account for month and year of conception fixed effects, respectively. ϵ_{ipqt} is the error term. To compute standard errors, I cluster them at the treatment level, that is the province, accounting for any unobserved common group effects. As there are only 22 provinces in the analysis, I report wild bootstrap standard errors to address the "few clusters" issue (MacKinnon and Webb, 2018; Roodman et al., 2019; Cameron et al., 2008). As the outcomes are evaluated as provincial averages, all regressions are weighted with the number of live births per unit of time-province.

β_1 is the coefficient of interest and measures the effect of increasing the geographic (at the province level) intensity of drug exposure on newborn's health after the policy implementation. The estimated effect is an intention-to-treat (ITT) effect.

The DiD estimator, to provide unbiased estimates of the treatment, is based on the assumption that in the absence of the treatment, the outcomes in the two groups (control and treated) would have followed parallel trends. In a setting where the treatment variable is continuous, an additional assumption may be required, which is called the "strong parallel trends assumption" (Callaway et al., 2021). In the design of this chapter, it means that provinces with a lower treatment intensity are a good counterfactual for provinces with a higher treatment intensity. As with the standard DiD design, this assumption cannot be directly tested, however in the robustness check section of this chapter, I provide several tests to support this assumption.

A threat to identification is potential changes in women's fertility. For instance, an increase in drug use may be also correlated with more risky behaviors resulting in an increase in the number of pregnancies. To discard this channel, I re-do Equation (3.1) using as dependent variables: (i) the natural logarithm of live births, (ii) the number of live births over women of reproductive age and (iii) the natural logarithm of the previous outcome. I present the results of this exercise in Section 3.6.

A common issue with the enactment of laws is that there can be an anticipation effect. To rule out this concern, I perform an event study version of Equation (3.1), in which I replace $Index_p * Post_{mt}$ with a set of interactions between the *index* (treatment variable) and pre and post-policy dummies. The event study also allows to add evidence on the plausibility of no differential pre-trends across provinces. In addition, it allows to examine the dynamics of the estimated effect. In Section 3.6, I show the results of the event study estimation.

An additional concern may be that newborns are experiencing detrimental health at birth because mothers are not taking antenatal care rather than an increase in drug consumption itself. To rule out this channel, in Section 3.6, I show evidence supporting no differential changes in prenatal care use.

3.5 Results

3.5.1 Newborn's Health Results

Table 3.3 shows the results of Equation (3.1) for newborn's health outcomes. Panel A presents estimates for health at birth such as birth weight, share low birth weight, share of very low birth weight, share of prematurity and share of c-sections. Panel B shows outcomes for neonatal hospitalizations, which includes the hospitalization rate (all causes) per 1,000 live births, drug-related hospitalization rate (per 1,000 live births), the hospital death rate (per 1,000 live births). I also present results for the # of drug-related hospitalizations, which is estimated using a negative binomial model and expressed as incidence rate ratios. All regressions include mother's province of residence fixed effects and month, year of conception fixed effects. The regressions of Panel A control for average mother's age,[9] while those of Panel B control for average neonate's age and gender.

Overall, Table 3.3 reveals that exposure to maternal drugs of abuse may be linked to detrimental health at birth. On average, live births in a province that was more drug-intensive prior to the policy, have higher probabilities to experience low birth weight and very low birth weight. In addition, average birth weight decreases. For example, increasing the initial exposure to drugs (measured by the *index*) by 1 percentage points decreases average birth weight by 33.35 g. It also increases the share of low birth weight live births by 1.43 percentage points, and the probability of very low birth weight by 0.37 pp. Related to the averages before the policy, this represents a decrease in birth weight of around 1%, and increase of approximately 16.26% and 82.22% in LWB and VLWB, respectively. There is no significant effect for pre-term deliveries or c-sections.

With respect to hospitalization outcomes, neonates residing in provinces that faced a greater exposure to drugs have, overall, higher hospitalization rates, specially those related to maternal drug use. There is an increase of around 17.73 drug-related hospitalizations per 1,000 live births per month. In addition, from the estimated negative binomial model, these neonates are expected to have a rate 1.77 times greater for drug-related hospitalizations. Related to pre-policy averages, there is an increase of around 31.92% in the drug-related hospitalization rate for newborns below 29 days of birth. Though all the estimated coefficients have the expected sign, I do not find any significant difference in all-causes hospitalization rate, or in the neonatal hospitalization mortality rate.

[9] As a robustness check, I also control for the share of low educated and single mothers, separately, at the province level. Results from this exercise are consistent with those of Table 3.3 and are available upon request.

Table 3.3: Newborns' Health Results

	Panel A: Health at birth outcomes					Panel B: Neonatal hospitalization outcomes			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Birth weight	**LBW**	**VLBW**	**Preterm**	**C-section**	**Hosp rate**	**Drug hosp rate**	**nbreg # drug hosp (IRR)**	**Hospital death rate**
ITT	-33.3558***	0.0143***	0.0037**	0.0082	0.0306	13.9405	17.7290*	1.7745**	1.5811
P-value	[0.0020]	[0.020]	[0.0130]	[0.1071]	[0.3564]	[0.2432]	[0.0571]	[0.0200]	[0.5746]
Pre-policy mean	3,118.75	0.0879	0.0045	0.0540	0.2964	147.439	55.5455	20.4394	5.1596
Mother's controls	yes	yes	yes	yes	yes	no	no	no	no
Patient's controls	no	no	no	no	no	yes	yes	yes	yes
Province FE	yes	yes	yes	yes	yes	yes	yes	yes	yes
Month of conception FE	yes	yes	yes	yes	yes	yes	yes	yes	yes
Year of conception FE	yes	yes	yes	yes	yes	yes	yes	yes	yes
Observations	1,320	1,320	1,320	1,320	1,320	1,319	1,319	1,319	1,319
Clusters	22	22	22	22	22	22	22	22	22

Notes: Columns 1 to 7 and 9 report OLS estimates of Equation (3.1). Column 8 is estimated using a negative binomial model and reports the incidence rate ratio. Regressions for health at birth outcomes adjust for average mother's age. Regressions for hospitalization outcomes adjust for average newborns' age (in days) and gender. Hospitalizations and mortality rate are per 1,000 live births. All regressions are aggregated at the province level and weighted with the number of live births in a province-unit of time. Post-estimation wild bootstrap p-values in square parenthesis for the small number of clusters. ***p<0.01, **p<0.05, *p<0.10.

The results from Table 3.3 show that newborns more intensively exposed to drugs are experiencing negative issues with birth weight but not with preterm birth. This result can be explained from the medical literature that has found a strong correlation between maternal drug use and low birth weight. For instance, Hwang et al. (2017) find that infants born to mothers with substance use disorders are at higher risk for adverse health outcomes in the perinatal period in terms of LBW. Likewise, Bailey and Diaz-Barbosa (2018), based on observational studies, conclude that babies born to mothers who abuse illegal substances are at increased risk of LBW. In addition, Umer et al. (2022) find that in-utero exposure to illegal substances is associated with low birth weight among rural newborns. Most of the literature finds associations between maternal drug use an preterm birth, however, I do not find any statistically significant difference in this outcome. This suggests that the effect seen in the indicators of low birth weight may not come from short gestational length-related complications.

With respect to hospitalization outcomes, I find an increase in the absolute number of drug-related hospitalizations and in the rate per 1,000 live births. The medical diagnoses that are included in this definition are those related with low birth weight, newborns affected by noxious influences transmitted via placenta, and related. These results are in line with those of the Panel A, in the sense that worse health at birth might induce higher probabilities of neonatal hospitalizations. Altogether, the results might imply a higher healthcare use. The medical literature on the relation between in-utero exposure to drugs and neonatal hospitalization supports these findings. For example, exposed neonates are more likely to have cardiac, respiratory, neurologic, infectious, hematologic, and feeding/nutrition problems, which leads to higher hospitalization rates (Hwang et al., 2017; Creanga et al., 2012).

The medical literature also finds that neonates born to mothers with substance use disorders are more likely to be re-admitted to hospitals Hwang et al. (2017). Unfortunately, due to data limitations, I am not able to distinguish which hospital discharge corresponds to a relapse and which to a completely new hospitalization. In this sense, one should keep in mind this point when interpreting the results of this subsection.

Finally, the lack of an effect for the neonatal hospital death rate, which was 5,16 per 1,000 live births before the policy implementation, is compatible with the absence of a significant result for preterm births. According to the World Health Organization, preterm birth complications are the leading cause of death among children under 5 years of age, moreover, studies under the medical literature find a positive association between maternal drug use and neonatal mortality rate, but mainly due preterm birth complications (such as immature immune systems) (Finnegan et al., 1977; Perin et al., 2022).

3.5.2 Women's Substance Use Disorders

The main hypothesis of this chapter is that newborns to mothers who are residents of provinces that faced a greater exposure to drugs may have worst health at birth outcomes. To support this statement, there should also be an increase in maternal drug use. In this section, I indirectly test it using two different data sources as described in Section 3.3.

To assess whether maternal drug use have changed, I need to look at consumption among pregnant women. However, there is no data for the country that registers maternal drug use. Despite this drawback, I can still indirectly test it by analysing changes in drug use in women of childbearing age. For this, I re-estimate Equation (3.1) for two samples: (i) hospital discharges, and (ii) outpatients visits.

For the first sample, the outcomes of interest are the drug-related hospitalization rate per 10,000 women, the proportion of drug-related hospitalizations relative to total female hospitalizations, and the rate of childbirth complications per 1,000 live births. All regressions include the same fixed effects as in Equation (3.1), standard errors are also clustered at the province level and estimated using the wild bootstrapping procedure. The variable $Post_{mt}$ takes the value of 1 for hospitalizations from January 2013 onwards, and zero otherwise. The vector of control variables includes average patient's age and a categorical variable which represents quintiles of the official poverty index at the patient's province of residence level.[10] As in the

[10]The poverty index is calculated by the National Institute of Statistics (INEC). It is a multidimensional poverty index that reflects the proportion of people who are multidimensionally poor in

newborn's outcomes, I conduct the analysis at the province level.

The outcome of interest in the second sample is the number of mental health visits per 10,000 women per province per year. Unlike the hospital discharge outcomes, this data is aggregated at the year level. For this reason, I add the year 2010 to the analysis to have more information regarding the pre-policy period. This regression includes province and year fixed effects and the poverty index categorical variable. The regression is weighted with the number of women of reproductive age in a province-year. The variable $Post_{mt}$ of Equation (3.1) is defined as in the women's hospital discharge sample.

Panel A of Table 3.4 presents the results for the hospital discharge outcomes. Women living in provinces with an initial larger exposure to drugs, on average, show larger chances to be hospitalized for causes related to drug use. For instance, an increase of 1 pp in the initial drug intensity exposure is linked to an increase of around 0.0885 drug-related hospitalizations per 10,000 women, after the enacting of the policy. Likewise, there is also significant increases in the proportion of drug-related hospitalizations and in complications during delivery. In the interpretation of the estimated coefficients, one needs to keep in mind that these outcomes reflect intensive substance use (cases that need to be hospitalized), and that few cases of drug use may require a hospital admission. Nevertheless, the medical literature shows that some types of illegal substances such as cocaine, heroine, and other synthetics are very harmful for the developing fetus even in small quantities (Finnegan et al., 1977; Creanga et al., 2012; Umer et al., 2022; Bailey and Diaz-Barbosa, 2018).

Research on the association between maternal drug use and health outcomes finds that women who use/abuse drugs are at a higher risk of complications during childbirth (Haight et al., 2018). Examples are: higher probability of preterm labor, failed induction of labor, obstructed labor due to malformations of fetus, intrapartum hemorrhage, umbilical cord complications, retained placenta and membranes, abnormality in fetal heart rate and rhythm increases. In column 3 of Table 3.4, I show that the rate of this type of complications increases in provinces that are initially more exposed to illicit drugs. In this section of the analysis, I focus on women's outcomes, however, father's drug consumption may also be important because father's lifestyle has a great social and psychological influence on the life of the mother and baby. In Table 3B.3 of the Appendix 3B, I show the results for drug-related hospitalization outcomes for men between 15 and 45 years old. They reveal a significant increase in illegal substance use-related hospital discharges.

Panel B of Table 3.4 shows the results for mental health visits per 10,000 women.

each province. More information can be found at www.ecuadorencifras.gob.ec

Table 3.4: Women's Substance Use

	Panel A: Hospital discharges outcomes			Panel B: Outpatient visits outcomes
	(1)	(2)	(3)	(4)
	Hospitalization rate per 10,000 women	Proportion of drug-related hospitalizations	Complications rate per 1,000 live births	Outpatient visits per 10,000 women
ITT	0.0885**	0.0029**	0.0741*	123.3849**
P-value	[0.0190]	[0.0190]	[0.0971]	[0.0120]
Pre-policy mean	0.1391	0.0038	0. 2503	251.2597
Socioeconomic controls	yes	yes	yes	yes
Province FE	yes	yes	yes	yes
Month FE	yes	yes	yes	no
Year FE	yes	yes	yes	yes
Observations	1,320	1,320	1,320	132
Clusters	22	22	22	22

Notes: Regressions from columns 1 and 2 are weighted with the number of total hospitalization for women in a province-unit of time. The regression of column 3 is weighted with the number of live births in a prov-unit of time, and column 4 with the number of women of childbearing age. Socioeconomic controls for columns 1 to 5 include average patient's age and a categorical variable which represents quintiles of the official poverty index at the patient's province of residence level. For column 4 only the quintiles of the poverty index. Post-estimation wild bootstrap p-values in square parenthesis for the small number of clusters. ***p<0.01, **p<0.05, *p<0.10.

This variable includes all conditions categorized in the group F of the ICD-10 classification. Although this variable does not allow to separate mental and behavioral disorders due to psychoactive substance use from other mental health issues (for example depression, schizophrenia, mood disorders, etc.), it still provides an indirect measure of the substance use at the outpatient level. Some mental health issues and substance use disorders may occur together. Examples are depression, anxiety disorders, schizophrenia, and personality disorders (Santo Jr et al., 2022b,a).

Women receiving health care in provinces initially more exposed to drugs have a higher mental health care visits rate, after the policy implementation. On average, 123.38 (per 10,000 women) more visits per year per province in face of an increase of 1 pp in the drug index exposure. To add validity to the results of this subsection, in Figure 3C.2 of the Appendix 3C, I provide event studies plots for the four outcomes. I use one period before the policy implementation as the reference category. Due to the structure of the data, the event studies estimates for women's hospital discharges is aggregated at the quarter level, and for the outpatient outcome at the year level. Overall, I do not find any suspicious of pre-trends in any of the outcomes. Moreover, the pre-policy coefficients are close to zero, and of smaller size compared to those of the post-policy period.

3.5.3 Mother's Characterization

In this section, I investigate whether the characteristics of mothers conceiving in initially more exposed provinces differ from those of their counterparts. Formally, I re-estimate Equation (3.1) using as outcomes average mother's age, the share of mothers below 19 years old, the share of low educated mothers, and the share of single mothers in province p, month m, year t. I include month, year of conception fixed effects and province fixed effects. Standard errors are clustered at the province level.

Table 3.5 shows that there are no significant differences in mothers characteristics, at least in terms of the outcomes analyzed. In spite of this, the coefficients are of the expected sign. For instance, an increase of 1 pp in the drug exposure index is associated with a decrease in average maternal age. The same is observed in the proportion of mothers under 19 years of age.

Table 3.5: Mothers' Characteristics

	(1)	**(2)**	**(3)**	**(4)**
	Age	**Below 19 y/o**	**Low educ**	**Single**
Continous DiD	-0.0189	0.0023	0.0123	0.0271
P-value	[0.7497]	[0.6116]	[0.2282]	[0.6136]
Pre-policy mean	25.5691	0.2003	0.3418	0.1289
Province FE	yes	yes	yes	yes
Month of conception FE	yes	yes	yes	yes
Year of conception FE	yes	yes	yes	yes
Observations	1,320	1,320	1,320	1320
Clusters	22	22	22	22

Notes: Data for pregnancies conceived between April 2010 and March 2015. All regressions are aggregated at the province level and weighted with the number of live births in a prov-unit of time. Post-estimation wild bootstrap p-values in square parenthesis for the small number of clusters. ***p<0.01, **p<0.05, *p<0.10.

3.6 Robustness Checks

This section reports several robustness checks for newborn's health results presented in Table 3.3. I start conducting an event study version of Equation (3.1) to support the hypothesis of no differential pre-trends. Then, I assess whether there are changes in fertility responses and in prenatal care use. Third, I perform a falsification test in which I change the policy implementation date. Finally, I propose an alternative specification of the treatment variable.

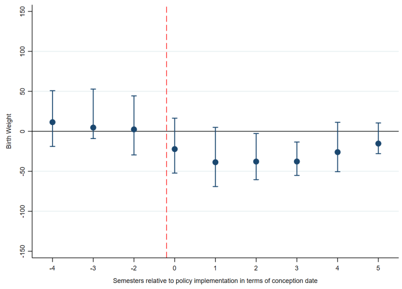
3.6.1 Event Study Estimates

The underlying assumption of Equation (3.1) under a DiD framework is that newborn's health outcomes across provinces should have evolved in parallel in absence of the policy. If this is not the case, then the estimated coefficient could be capturing differences in trends rather than the actual effect of the policy. To support the plausibility of the parallel trends assumption, I replace the variable $Post_{mt}$ in the interaction term of Equation (3.1) by a set of dummy variables for every pre-post policy period. I generate four pre-policy and six post-policy periods on a semi-annual basis.[11] I set the reference category to one period prior to the policy ($t = -1$). Figure 3.3 depicts the set of pre-post policy coefficients of the DiD estimator with their 95% confidence intervals.

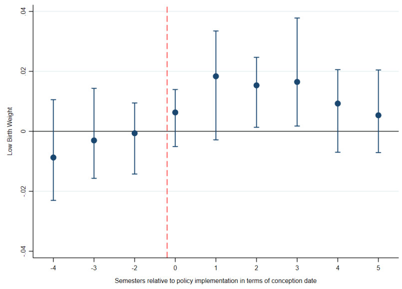
Overall, the set of pre-policy coefficients are statistically insignificant and of a small magnitude compared to post-policy coefficients. For preterm births, I find a statistically significant coefficient in the pre-policy period but it is on the very first period (first semester of year 2011). With respect to average birth weight, the share of low birth weight, very low birth weight, and c-sections, I do not find any significant coefficient (at the 95% level) for the interaction terms in the pre-policy period, which is compatible with no differential pre-trends. Moreover, the set of pre-policy coefficients are, overall, close to zero, which adds evidence supporting no anticipatory effect. Finally, in the post-policy implementation (located to the right of the dotted line in the graph) period there are statistically significant coefficients, which points to detrimental health at birth in provinces with an initial greater exposure to drugs.

[11] Because births may be seasonal, I have constructed the pre-post policy periods on a semi-annual basis, which results in 4 pre and 6 post policy periods.

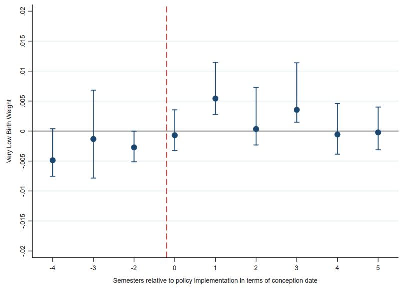
Figure 3.3: Event Study Plots of Newborn's Health Outcomes



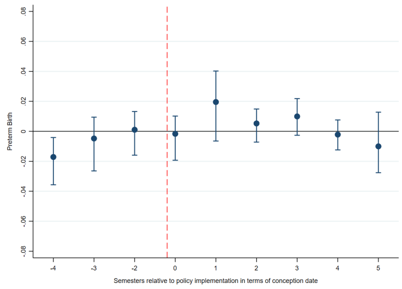
(a) Birth Weight



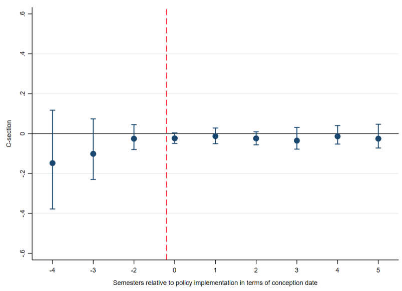
(b) Low Birth Weight



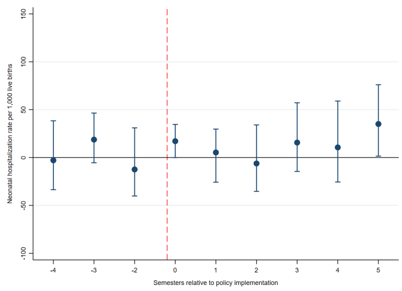
(c) Very Low Birth Weight



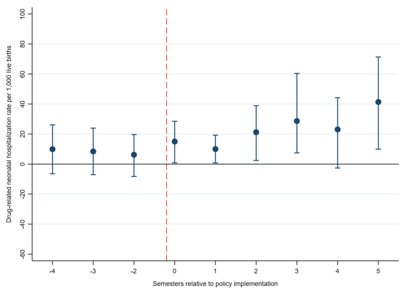
(d) Preterm Births



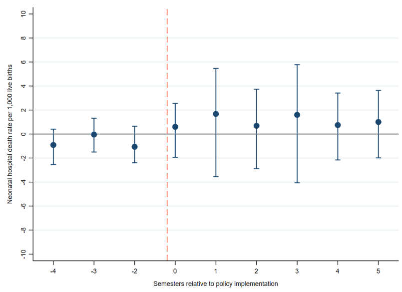
(e) C-Section



(f) Hospitalization rate



(g) Drug-hospitalizations rate



(h) Hospital deaths rate

Notes: Figure shows the set of coefficients corresponding to the interaction term from the event study versions of Equation (3.1). Time periods are defined on an semi-annual basis. Points estimates are shown along with their 95 percent confidence intervals. Reference category is one period prior to the policy and is omitted from the graphs. Confidence intervals are estimated by wild-bootstrap cluster method with 999 repetitions.

3.6.2 Fertility Responses and Prenatal Care

Exposure to drugs may influence maternal fertility because it might induce women to practice risky sexual behaviours. Baggio et al. (2020) in an analysis for marijuana use in the U.S. find an increase in sexual activity, a reduction in contraceptive use, and an increase in the number of births. The authors attribute the findings to behavioral responses, which includes increased attention to the immediate hedonic effects of sexual contact, increased sexual frequency (Sun and Eisenberg, 2017), and a reduction in the perception of risks of unprotected sexual intercourse. I evaluate whether this might be the case in the setting of this chapter. To do this, I re-estimate Equation (3.1) using the number of live births as outcome variable. I employ three different transformations of the number of live births, its natural logarithm, the rate per 10,000 women of childbearing age, and the natural logarithm of the rate per 10,000 women. The regressions include average mother's age (at the province level) as control variable. I conduct the analysis around the estimated conception date as in the baseline specification. I report p-values computed from a wild bootstrapped procedure with clusters at the province level. Table 3.6 shows the results of this exercise.

Table 3.6: Fertility Changes and Prenatal Care

	(1)	(2)	(3)	(4)
	ln(births) rate per	**Births per 10,000 women**	**ln(births per 10,000 women)**	**Prenatal care**
ITT	-0.0450	-1.4427	-0.0437	0.0126
P-value	[0.3413]	[0.3413]	[0.3123]	[0.1692]
Mother's controls	yes	yes	yes	yes
Province FE	yes	yes	yes	yes
Month of conception FE	yes	yes	yes	yes
Year of conception FE	yes	yes	yes	yes
Observations	1,320	1,320	1,320	1,320
Clusters	22	22	22	22

Notes: Regressions for live births conceived between April 2010 and march 2015. Regressions of columns 1 to 3 are weighted with the number of women of childbearing age. Outcome 4 is weighted with the number of live births. Outcome 2 corresponds to the ratio between total number of live births and the total number of women of childbearing age (per 10,000 women). Outcome 3 is the natural logarithm of outcome 2. All regressions include average maternal age per province and month-year as control. Post-estimation wild bootstrap p-values in square parenthesis for the small number of clusters. ***p<0.01, **p<0.05, *p<0.10.

Overall, I do not find evidence that supports any compositional effects in live births. For instance, outcome (1) of Table 3.6 shows that there is no significant difference in the (ln) of live births per province and month-year. Likewise, when I analyze the birth rate per 10,000 women and its natural logarithm, I do not find any statistically significant result.

Prenatal care is positively associated with birth outcomes (Balsa and Triunfo, 2015; Currie and Grogger, 2002). In the context of maternal drug use, there are studies that find a negative association between drug consumption and adequate prenatal care. For instance, Shieh and Kravitz (2006) find that pregnant cocaine and heroin users are more likely to initiate prenatal care later than other women. Simmons and Austin (2022) find that pregnant women who report to have used illicit drugs are more likely to initiate prenatal care later and less likely to receive adequate prenatal care. In Column 4 of Table 3.6, I assess whether this is also happening in the context of this chapter. For this, I re-estimate Equation (3.1) using as outcome the share of women receiving prenatal care in province p, month m and year t, variable that I construct from the live births data. The regression includes mother's characteristics as control variables, and is weighted with the number of live births.[12] The results show that there are not any statistically significant changes in the share of pregnant women receiving prenatal care.

3.6.3 Falsification Test

Live births conceived before April 2012 were not affected by the policy at any degree. Therefore, I should not expect to find any detrimental health at birth outcomes for these newborns, otherwise there might be other factors besides the policy that are affecting high, medium and low drug-intensity exposure provinces in terms of negative newborn's health. To rule out this concern, I re-estimate Equation (3.1) considering only live births that were conceived before April 2012, and assuming that the policy took place in January 2012 (instead of 2013). Panel A of Table 3B.4 in Appendix 3B shows the results of this exercise.

Overall, all the coefficients show no statistically significant relation with the placebo policy date. In addition, the placebo estimated coefficients, $\beta 1's$ are smaller than those estimated for the actual policy date. For instance, for the average birth weight, the placebo date finds a negative coefficient of around 7 g versus 33 g that I find for the actual policy implementation date.

[12]Due to data limitations, I am only able to assess the outcome shown in column 4 of Table 3.6. However, it would be ideal to evaluate other outcomes such as the number of visits per trimester of pregnancy and at least 9 prenatal care visits.

3.6.4 Other Treatment Definition

In this robustness check, I dichotomize the treatment variable (the *index*) and generate a control and a treatment group. From the original *index* variable, I generate five quintiles and assign to the treated group those provinces that belong to the fifth, four and third quintiles of the *index* distribution. Provinces that are in the second and first quintiles are, therefore, assigned to the control group. Formally, in Equation (3.1), I replace $Index_p$ by $Treat_p$, which is a dummy variable that takes the value of one for provinces in the treated group and zero for those in the control group. Control variables and fixed effects are the same as those used in Table 3.3. Standard errors are also clustered at the province level.

Panel B of Table 3B.4 in the Appendix 3B shows that the results of the binary treatment are consistent with those of the continuous treatment in terms of the direction of the estimated coefficients. This set of results also points to detrimental newborn's health across the outcomes of interest. In addition, I find statistically significant differences (at the 95% confidence level) in terms of average birth weight, the share of low birth weight, and the share of very low birth weight newborns. For example, in provinces with a prior high drug intensity exposure, average birth weight is, on average, 25.67 g less compared to provinces with a lower drug intensity exposure. Using a binary treatment definition implies that units with small intensities become part of the control group although they also respond to the policy change. On this, Schmidheiny and Siegloch (2022) argue that depending on the elasticity of the treatment effect with respect to the reform intensity, it is possible that the estimates obtained from the dichotomization can be larger, smaller or identical to the continuous treatment.

3.7 Conclusions

Drug use is considered a worldwide public health issue that causes important socioeconomic losses. According to the World Health Organization (2021), for year 2019, drug use disorders resulted in 18 million years of healthy life lost. Likewise, drug consumption does not only affect the primary consumer, but can be easily transmitted from mother to the fetus through the placenta. Intrauterine exposure to substances of abuse is linked to detrimental health at birth outcomes (Bailey and Diaz-Barbosa, 2018; Haight et al., 2018).

As a public health issue, governments need to be able to identify consumers who are in need of health care to offer them the required treatment. In an attempt to recognize the health needs of drug consumers, as well as, to distinguish them from traffickers, in year 2013, Ecuador issued a regulation that established maximum permissible amounts of possession for personal consumption of substances subject to control by means of a table that contains thresholds of drug tenancy for marijuana, cocaine, heroin, and other synthetic drugs.

In this chapter, I study how the above mentioned regulation might affect health at birth by means of an unexpected increase in drug use among women of reproductive age. While the policy was designed not to incarcerate consumers who actually need health care, it seems that it contributed to an increase in illicit drug use. Using a difference-in-differences approach that compares provinces initially more exposed to drugs to those less exposed, before and after the regulatory change, I find, as preliminary results, a reduction in average birth weight, an increase in the share of low and very low birth weight newborns. I also find that newborns whose mothers reside in more drug-exposed provinces have higher hospitalization rates to conditions related to maternal drug use. However, I do not find any difference in the neonatal hospital death rate.

I also show that in provinces with an initial higher illicit drug exposure, there is an increase in the women's hospitalization rate to conditions directly related to drug use. Moreover, there is evidence supporting an increase in average outpatient visits related to mental health issues, which are common comorbidities of drug use. These results are compatible with an increase in the consumption of illicit substances. However, this might also reflect an increase in hospital-seeking behaviour, expansion of health care targeted to drug conditions, or more reporting of drug diagnoses. While I cannot rule out these channels, the results of neonatal hospitalizations to conditions related to maternal drug use support an increase in women's drug use.

Negative consequences of maternal drug use on newborn's health may come not only from mother's consumption itself but from other related factors that can play a role. For instance, drug use is linked to risky behaviours (Rashad and Kaestner, 2004; Baggio et al., 2020), which increases the probabilities of infections that can be transmitted to the fetus. Consumption of some type of illicit drugs is also associated with feeding problems of the mother, which can decrease the nutrients that the fetus needs for adequate development. Moreover, environmental factors around drug consumption may also affect newborn's health. While I am not able to rule out all these factors, I provide evidence supporting no significant changes in women's fertility, and in the probability of receiving prenatal care. Despite the existence of these additional channels, they are all somehow related to the negative consequences that drug consumption may have on individuals and the society.

The main results on newborn's health, while preliminary, are robust to a set of tests. However, there are some other limitations, which mainly come from restrictions in data availability. First, I do not have information on mother's health status, which may influence newborn's health. Second, women's substance use results reflect intensive drug use that requires hospitalizations, and therefore not all cases of drug use are considered, which depending on the type of substance, may have important negative consequences on the fetus. I try to overcome this limitation by exploring mental-health care outpatient visits. While this last measure may not be the most accurate, it helps to provide some insights about less intensive drug use. In spite of these limitations, I believe the results are robust and provide valuable information on the unintended consequences of a regulatory change aimed at not incarcerating drug consumers.

While the results of this chapter are still preliminary, the findings seem to suggest that the increase in drug use, through increased availability, is possibly one of the channels affecting newborn's health living in provinces more exposed to drugs.

Finally, the (preliminary) results of this chapter shed light on the unintended consequences that regulatory change might have had on newborn's health. While considering drug consumers as individuals in need of health care is an important step to tackle the negative effects of illicit drug use, effective policies to discourage consumption should be placed in the agenda.

Appendix 3A: Computation of the Geographical Drug Exposure Index

1. Definition

The “*index*” is a measure that aims to help to identify the high incidence zones regarding the illicit drug problem at the province level. The *index* is built from four individual indicators that capture information from both the supply and demand side.

2. The individual indicators

2.1 Annual prevalence of drug use

Measures the annual prevalence of illicit drug use per province i (in percentages). The source is the IV National Survey of Drug Use conducted in year 2012. It is calculated as follows:

$$\frac{\text{\# of respondents who declare to have consumed any illicit drug in the previous year in province } i}{\text{Population at risk in province } i}$$

2.2 Frequent use of illicit drugs

Measures the percentage of individuals who declare a frequent use of any type of illicit drugs. Source is the "IV National Survey of Drug Use" conducted in year 2012. Frequent use is defined as "the substance is continued to be used after having experimented and its consumption is integrated into the usual lifestyle".

2.3 Drug-related hospitalization rate

Measures the hospitalization rate due to (illicit) drug consumption per 10.000 inhabitants between 13 and 45 years old per province i. I use administrative data from the “Instituto Nacional de Estadísticas y Censos” on the universe of hospitalizations in the country, and census data to obtain the population (denominator), both from year 2012. The indicator is calculated as follows:

$$\frac{\text{\# of drug related hospitalizations in province } i}{\text{Population at risk in province } i}$$

The numerator includes the following medical diagnoses, according to the ICD-10 classification:

Table 3A.1: ICD-10 Diagnostics Used in the Individual Indicator

ICD-10	Description
F11	Mental and behavioral disorders due to use of opioids
F12	Mental and behavioral disorders due to use of cannabinoids
F13	Mental and behavioral disorders due to use of sedative hypnotics
F14	Mental and behavioral disorders due to use of cocaines
F15	Mental and behavioral disorders due to use of other stimulants, including caffeine
F16	Mental and behavioral disorders due to use of hallucinogens
F18	Mental and behavioral disorders due to use of volatile solvents
F19	Mental and behavioral disorders due to multiple drug use and use of other psychoactive

Elaboration: The Author.
Source: ICD-10 available at `https://eciemaps.mscbs.gob.es/ecieMaps/browser/index_9_mc.html`

2.4 Drug trafficking arrest rate

This indicator measures the arrest rate due to drug trafficking per 10.000 inhabitants per province i. To construct this index, I use administrative records from the "National Police" institution in Ecuador on the universe of arrests during year 2012. For population data, I use Census data from the "Instituto Nacional de Estadísticas y Censos". Thus, the indicator is calculated as follows:

$$\frac{\text{\# of drug trafficking arrests in province } i}{\text{Population in province } i}$$

3. Methodology for index calculation

I use the common factor analysis Maximum Likelihood technique to generate the *index*. This method allows to combine individual indicators that may: (i) be measured on different scales, (ii) have different distributions, (iii) may or may not apply to the same individual, and (iv) measure, to different degrees, the underlying idea of the index imperfectly (Smith et al., 2015). As the purpose of this index is to use indicators that are individually representative of the drug problem per province, the analysis involves exploring a one-common factor model. If during the computation, a second factor arises, it would suggest the existence of a new domain. I ruled out this possibility using the standard test of examination of the Eigen values.

The process of combining the four indicators using a one-common factor analysis is as follows:

1. All indicators were standardized to a normal distribution with zero mean and a standard deviation of one.

2. The standardized scores were factor analyzed using the Maximum Likelihood

method. From this process, I obtained a set of weights for each indicator. Weights are shown in Table 3A.2.

Table 3A.2: Weights for Individual Indicators

Indicator	**Weight**
Annual prevalence of drug use	0.28040
Frequent use of illicit drugs	0.17339
Drug-related hospitalization rate	0.45093
Drug trafficking arrest rate	0.19596

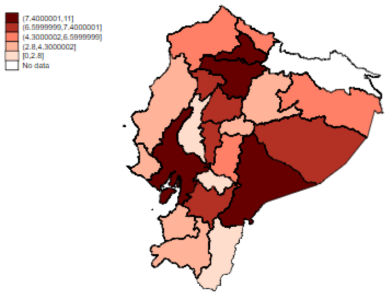
Elaboration: The Author.
Source: Weights derived after performing a (one) factor common analysis.

3. The indicators were combined using the weights presented in Table 3A.2.
4. To facilitate interpretation, the factor scores can be transformed into an index that has a maximum value of 1 and a minimum of 0.

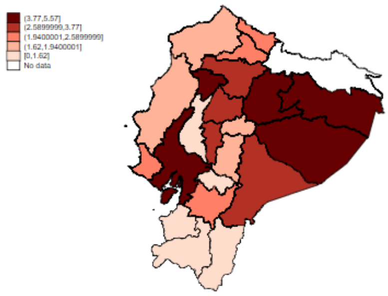
4. Index visualization

Figure 3A.4 shows the geographical variation of the individual indicators. Overall, there is a similar pattern across the four indicators, which supports the combination of them into a single index.

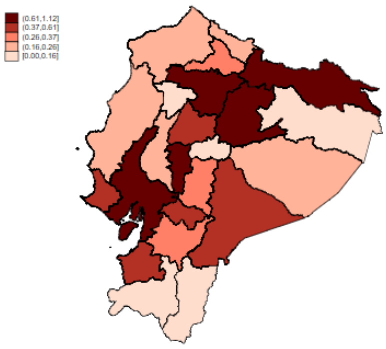
Figure 3A.4: Geographical Variation of Individual Indicators



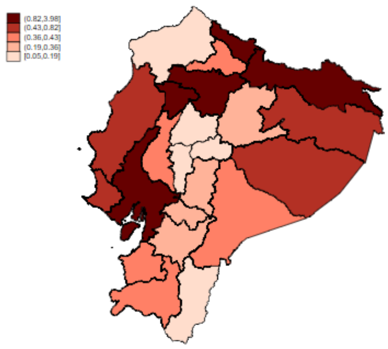
(i) Annual prevalence of drug use



(j) Frequent use of illicit drugs



(k) Drug-related hospitalization rate



(l) Drug trafficking arrest rate

Notes: Figure shows the geographical distribution of the individual indicators of the composite index. All indicators are calculated using data from year 2012.

Appendix 3B: Additional Tables

Table 3B.1: International Classification of Diseases Codes for Neonatal Hospitalization Outcomes

ICD-10	Description
P04	Fetus and newborn affected by noxious influences transmitted via placenta or breast milk
P07	Disorders related to short gestation and low birth weight, not elsewhere classified
P27	Chronic respiratory disease originating in the perinatal period
P96	Other conditions originating in the perinatal period
P92	Feeding problems of newborn
Q00-Q99	Congenital malformations, deformations and chromosomal abnormalities
F18	Mental and behavioral disorders due to use of volatile solvents
F10-F19	Mental and behavioural disorders due to psychoactive substance use
P961	Neonatal withdrawal symptoms from maternal use of drugs of addiction
P584	Neonatal jaundice due to drugs or toxins transmitted from mother or given to newborn

Elaboration: The Author. Source: ICD-10 available at `https://eciemaps.mscbs.gob.es/ecieMaps/browser/index_9_mc.html`

Table 3B.2: International Classification of Diseases Codes for Women Drug-Related Hospitalizations

ICD-10	Description
F11	Mental and behavioral disorders due to use of opioids
F12	Mental and behavioral disorders due to use of cannabinoids
F13	Mental and behavioral disorders due to use of sedative hypnotics
F14	Mental and behavioral disorders due to use of cocaines
F15	Mental and behavioral disorders due to use of other stimulants, including caffeine
F16	Mental and behavioral disorders due to use of hallucinogens
F18	Mental and behavioral disorders due to use of volatile solvents
F19	Mental and behavioral disorders due to multiple drug use and use of other psychoactive
F55	Abuse of non-dependence-producing substances
O35.5	Maternal care for (suspected) damage to fetus by drugs
R78	Findings of drugs and other substances, not normally found in blood
T40	Poisoning by narcotics and psychodysleptics [hallucinogens]
T43	Poisoning by psychotropic drugs, not elsewhere classified
G405	Special epileptic syndromes (drugs)
K292	Alcoholic gastritis

Elaboration: The Author. Source: ICD-10 available at `https://eciemaps.mscbs.gob.es/ecieMaps/browser/index_9_mc.html`

Table 3B.3: Men's Substance Use Hospital Discharges

	(1)	**(2)**
	Hospitalization rate per 10,000 men	**Proportion of drug-related hospitalizations**
ITT	0.1076***	0.0036***
P-value	[0.0000]	[0.0000]
Socioeconomic controls	yes	yes
Province FE	yes	yes
Month FE	yes	yes
Year FE	yes	yes
Observations	1,320	1,320
Clusters	22	22

Notes: Data from drug related hospitalizations for men between 15 and 45 years old from years 2011 to 2015. Regressions are weighted with the number of total hospitalization for men in a prov- unit of time. Socioeconomic controls include average patient's age and a categorical variable which represents quintiles of the official poverty index at the patient's province of residence level. Post-estimation wild bootstrap p-values in square parenthesis for the small number of clusters. ***p<0.01, **p<0.05, *p<0.10.

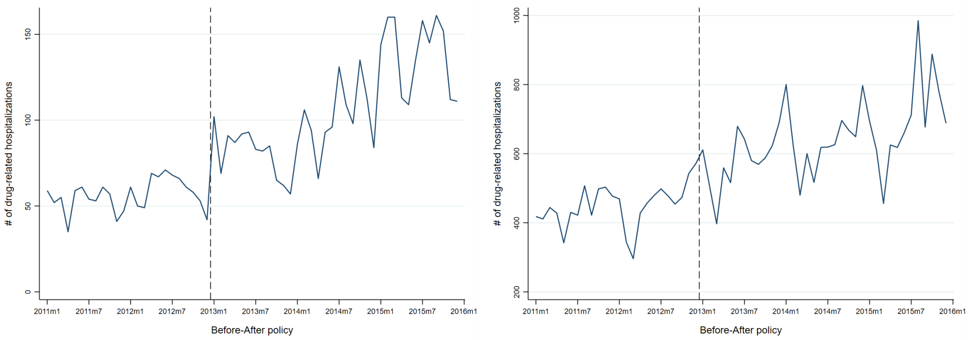
Table 3B.4: Robustness Checks

	Health at birth outcomes					**Neonatal hospitalization outcomes**			
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**
	Birth weight	**LBW**	**VLBW**	**Preterm**	**C-section**	**Hosp rate**	**Drug hosp rate**	**nbreg # drug hosp (IRR)**	**Hospital death rate**
Panel A: Falsification tests									
Policy January 2012 (ITT)	-7.0372	0.0027	0.0031	0.0129	0.1118	-11.7383	-6.0726	0.7828	0.0939
P-value	[0.5285]	[0.6036]	[0.1652]	[0.1822]	[0.4464]	[0.2933]	[0.2783]	[0.4234]	[0.9029]
Observations	528	528	528	528	528	527	527	527	527
Panel B: Binary Treatment									
Binary DiD (ITT)	-25.6755***	0.0106**	0.0030**	0.0069*	0.0004	7.9734	10.1291	1.2922	0.5071
P-value	[0.0030]	[0.0170]	[0.0130]	[0.0841]	[0.9890]	[0.4895]	[0.1251]	[0.1101]	[0.7177]
Observations	1,320	1,320	1,320	1,320	1,320	1,319	1,319	1,319	1,319
Mother's controls	yes	yes	yes	yes	yes	no	no	no	no
Patient's controls	no	no	no	no	no	yes	yes	yes	yes
Province FE	yes	yes	yes	yes	yes	yes	yes	yes	yes
Month of conception FE	yes	yes	yes	yes	yes	yes	yes	yes	yes
Year of conception FE	yes	yes	yes	yes	yes	yes	yes	yes	yes
Clusters	22	22	22	22	22	22	22	22	22

Notes: Panel A presents regressions assuming that the policy took place one year prior to the actual date – January 2012. Panel B shows regressions for a dichotomous treatment definition. All regressions for health at birth outcomes adjust for average mother's age. Regressions for hospitalization outcomes adjust for average newborns' age (in days) and gender. Hospitalizations and mortality rate are per 1,000 live births. All regressions are aggregated at the province level and weighted with the number of live births in a province-unit of time. Post-estimation wild bootstrap p-values in square parenthesis for the small number of clusters. ***p<0.01, **p<0.05, *p<0.10.

Appendix 3C: Additional Figures

Figure 3C.1: Trends in Drug-Related Hospital Discharges

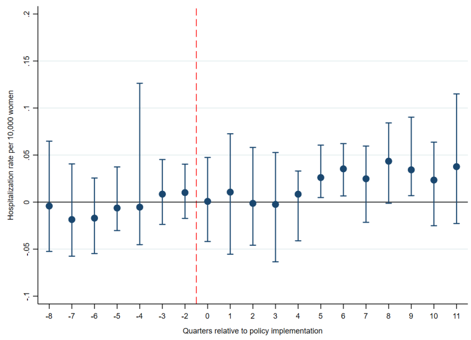


(a) Adults sample

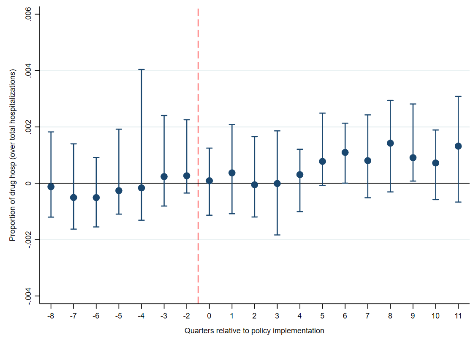
(b) Neonates sample

Notes: Own elaboration using data from Hospital Discharges years 2011-2015 provided by the INEC. The graph shows the evolution of the number of drug-related hospitalizations for women and men between 15-45 years old.

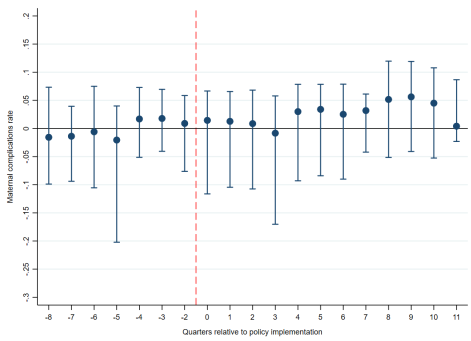
Figure 3C.2: Event Study Plots of Women's Substance Use



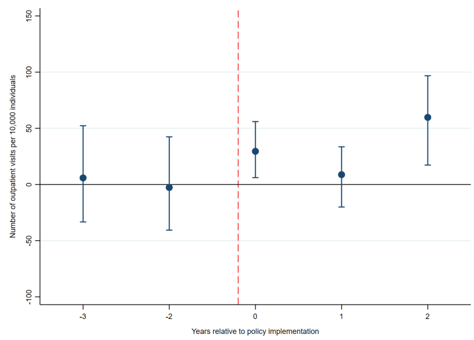
(a) Hospitalization rate per 10,000 women



(b) Proportion of drug-related hospitalizations



(c) Maternal complications rate



(d) Outpatient visits rate

Notes: Figure shows the set of coefficients corresponding to the interaction term from the event study versions of Equation (3.1) for women's substance use. For outcomes (a), (b) and (c) the time periods are defined on a quarterly basis, for outcome (d) on an yearly base. Points estimates are shown along with their 95 percent confidence intervals. Reference category is one period prior to the policy and is omitted from the graphs. Confidence intervals are estimated by wild-bootstrap cluster method with 999 repetitions.

4 Job Competition in Civil Servant Public Examinations and Sick Leave Behavior

4.1 Introduction

In several countries the entry system to access permanent civil servant positions relies on the traditional process of formal public examinations,[1] which application is often justified as a mechanism to guarantee the objectivity and transparency of the process (Bagüés, 2005). Through public examinations doctors, teachers, judges and other positions are selected to provide public service. An example of countries in which public examinations is the main system are Portugal, Italy, Belgium, Greece, France, Spain and others in Latin America.

Civil servant positions are very attractive as they offer benefits that do not exist in the private sector (these benefits vary according to the legislation of the country). For instance, in Spain, the most relevant benefit is long term employment stability. This "secure job" condition may also be translated into economic stability for individuals. In addition, the public sector, in most of cases, offers better working conditions in terms of vacation days, labor conciliation policies and working hours. In this sense, the attractiveness of public service positions may encourage individuals to participate, situation that is not negligible. For example, in Spain in the year 2018, for a call of 4,725 positions there were more than 80,000 participants registered for taking the examination (Ministerio de Hacienda y Función Pública, 2018). In addition, according to the Spanish Labor Force Survey, EPA in Spanish, during the second trimester of year 2010 the number of Spanish people preparing for a public examination was around 72,438.

To access civil servant positions, individuals have to take an examination called "oposición" in Spanish, or public examination in English, which is not an easy task. According to Bagüés (2005) preparing a public examination usually requires a lot

[1] There are two main systems to entry public service: (i) private-sector style recruitment, and (ii) traditional civil service recruitment with formal public examinations. However, the former is usually applied for non-permanent contracts.

of time, which may be more critical for individuals who are working and preparing an exam at the same time. There might be several possibilities that originate from workers who are faced with time constraints and the need to study. For example, this could cause workers to: (i) reduce the allocation of time for sleeping which might reduce bodily function, (ii) reduce the allocation of time for leisure increasing stress or anxiety, (iii) increase stress/anxiety related to high stakes civil service recruitment testing, situations that might increase the rate of sickness episodes. Therefore, we can frame the impact of the Spanish civil service recruitment process assessing whether civil servant positions announcements increase sickness episodes.

This chapter's aim, therefore, is to identify whether public examinations from new calls of civil servant positions have an impact on sickness absences. We make use of rich administrative dataset on the universe of sickness absences and civil servant positions offered in Spain from 2009 to 2015, and exploit the variation coming from the different number of positions announced (and with examinations) across Spain and time.

Research on public examinations as a selection process has been somehow explored in the physiological and personnel management literature, where the topic is analyzed in the aim to establish "best practices". However in the economic literature, research is still very scarce. For instance, Sundell (2012) studies what is the best way to recruit civil servants, he finds that when the risk for patronage is low, private sector style-recruitment can be attempted to find the best candidates, but when the risk instead is high, public examinations are preferable to prevent patronage and nepotism. Bagues and Esteve-Volart (2006), using administrative data from all the major public examinations in Spain, analyze how overconfidence affects the gender gap in promotion. They found that most of candidates are women and that conditional on experience, men have higher probabilities of passing the exams, however men are affected with greater overconfidence which leads to greater dropout rates compared to women. In other work, the same authors explore the potential existence of a "state nobility" in top civil service in Spain, for this, they use information on public examinations in the main Corps of the Spanish Administration and find that candidates that have a relative in the Corps are more likely to apply for these positions compared to non-relatives, and also to perform better in the exams (Bagues and Esteve-Volart, 2008).

As a first insight of the results of this chapter, we find that, overall, there is an increase in the absence rate (per 1,000 workers) in the public administration, education and health sectors several months before the public examinations. We find an increase by up to 5 spells per 1,000 workers per month per province. With respect to the intensive margin, we find that the average mean duration of the sickness spell may increase to up to 5 days in the more sensitive economic sector (the education sector, in our sample). In addition, we find mixed evidence on whether women(men) are more responsive than men(women) in terms on sick-related work absences. We also explore which type of medical conditions experience variation as a response to public examinations, and find that the effect is mostly driven by stress-related absences. Our results are robust to several checks, such as a placebo test, and changes in the treatment definition. We also provide some evidence which supports a negative relation between public examinations and public sector workers' health. We find an increase in the probability of general practitioner and specialist visits and in the number of specialist visits. We also find some significant differences between women and men in terms of self-assessed health.

This chapter contributes to the existing literature on the civil servants recruitment system in several ways. First, to the knowledge of the authors, this is the first attempt to investigate the causal relationship between public examinations and sickness absences. We also explore differences across economic sectors and demographic groups like gender. Second, we add evidence supporting that stress-related conditions seem to be the leading cause of the increase in sickness absences. Finally, we use two high-quality official administrative data set for our analysis. Sick leave data contains the universe of sickness spells occurred in Spain during the years 2009 to 2015. With respect to the public examination data, we built the dataset using official administrative records, and generated a novel data set containing the universe of public calls for years 2009 to 2015.

The results of this chapter are of important policy implications as public examinations might be generating negative externalities in terms of work absences. In addition, sickness absence represents an important cost to the social security system in Spain and also to the society through productivity losses. For example, in 2015, public spending on incapacity represented around 2.4% of the GDP (OECD Data, 2015), being this number above the average of OECD countries (1.9%). In terms of absence from work due to illness, Spain registered an average of 11.3 days per year per worker (OECD Stat, 2015). Our results are also of external validity as several countries across the globe use public examinations as a civil servant selection system.

The rest of the chapter proceeds as follows. Section 4.2 provides the institutional background on public examinations and on the sick leave system in Spain. Section 4.3 describes the data sets and shows some descriptive statistics. Section 4.4 presents the empirical strategy. Section 4.5 presents the results. In Section 4.6, we explore the composition of sick leave spells regarding the most common illnesses used in medical certificates. Section 4.7 presents a set of robustness checks. In Section 4.8, we provide evidence on public sector workers' health. Finally, Section 4.9 concludes with a discussion of the results.

4.2 Institutional Background

4.2.1 Public Examinations in Spain

Public examinations is the selection method used to gain access to a variety of permanent positions in the civil service in several countries. Some examples are countries in Latin America, Asia and Europe such as Portugal, Italy, Belgium, Greece and Spain, etc.

In Spain, according to the Boletín Oficial del Estado (BOE) (2019), the process to enter civil service is made through three type of systems: (i) public examinations, which consist of one or more passing tests to assess applicant's knowledge in the relevant area, (ii) open competition based on merits and qualifications of applicants, and (iii) open competition-public examinations, which is a combination of the other two. The public examination is the ordinary entry system, however, sometimes, due to the nature of the functions to be performed, the competition-public examination may be used and, exceptionally, the open competition. Regarding public examinations, they must include general and specific topics related to the functions of the position to be filled, they can also include psycho-technical "tests" to improve the objectivity of the selection process.

The process begins with the authorization, by the competing authorities, of the "public job offer", after this, the institutions are allowed to publish the calls in the official journals of the public administration.[2] There is no minimum nor maximum in the number of positions announced, thus, an institution can offer 1, 1,000 or more places (Boletín Oficial del Estado (BOE), 1984). The calls must contain all the requirements and characteristics of the positions. For instance, the number of positions, access system, number of examinations required, grading system, etc. Once the call is published, candidates have 20 days to submit their applications, after this period, the competing authority issues a resolution that contains the list of admitted and non-admitted candidates based on the accomplishment of the requirements established in the call. Only admitted participants can take the examination(s) which are formally announced indicating the exact place, date and hour of celebration. Candidates who pass the exam continue to the last stage of the process that consist on the exhaustive revision of the documents provided in support of their qualifications.[3] Finally, candidates who passed the exam(s) and meet all the requirements are declared permanent civil servants and are assigned to their places of work.

Civil servant positions can be offered at all administrative levels: (i) The General

[2]The Official Journals refers to "Boletin Oficial del Estado" which is at the national level; and to all the locals and regional journals.

[3]This refers to academic, professional experience, courses, and any other requirement according to what was established in the official call.

State Administration that comprises all jobs whose scope of action is at the national level; (ii) Regional Level "Comunidades Autónomas", that includes most of the education and health economic sector positions; and (iii) the local administration institutions, usually municipalities and others that operate at the local level.

Looking more into the exams, they usually involve a huge load of material to study, which includes general and job-specific questions. Therefore, in order to pass a public examination, candidates must acquire very-specific knowledge, which may not be very useful for other positions in the labor market (Bagues and Esteve-Volart, 2006). To have an idea of the amount of material to study, for a call in year 2015 of 44 positions for administrative assistants in the region of Galicia, applicants had to learn 30 topics for the examination, which range from the the Spanish Constitution of 1978 to specific administrative procedures (Diario Oficial de Galicia, 2015).

Anecdotal evidence of candidates preparing public examinations indicates that 6 out of 10 combine work with exam study (Ecoaula.es, 2022). Likewise, participants begin their preparation several months before the official examination date. However, the study habits become more demanding as the exam date approaches. For example, on average, more than two-thirds dedicate between 5 and 7 days per week to study, and during the months previous to the exams the daily studying time can increase up to 10 hours (Opositatest, 2019). This behaviour is compatible with increased probability of reducing the allocation of time for sleeping or leisure, increased stress and anxiety which might increase sickness absence rates in those candidates that are also working.

4.2.2 Sick Leave System in Spain

Spain has a social security system that has several benefits for workers. One of them is paid sick leave, which has the objective to compensate the income loss as consequence of temporary (common) illness, professional illness or accident. Common illness are considered all medical conditions that are independent of work activity, whereas professional illness refers to those acquired as a result of the work performed in the institution. Working accidents are any sudden events that occurs due to or during work, and that produce in the employee an organic injury, mental disturbance, a disability or death (Ministerio de Trabajo y Economía Social, 2023).

As explained by Marie and Vall Castelló (2020), to access sick leave benefits derived from a common illness, individuals have to be working, and contributing to social security for at least 180 days of the last 5 years before the beginning of the condition that generated the sick leave. Regarding the benefit amount, the system distinguishes between private and public employees. For private workers, when the origin is a common illness, individuals do not receive any compensation during the first 3 days of the absence, from the 4th to the 20th day they are entitled to the 60% of previous wage, and from the 21st day onwards to the the 75% (Ministerio de Trabajo y Economía Social, 2023). As seen, both public and private employees have the benefit of paid sick leave, though with different schemes, they are very similar in terms of coverage.

4.3 Data and Descriptive Analysis

4.3.1 Sick Leave Data

We use administrative data provided by the social security administration of Spain for years 2009 to 2015 on the universe of sickness absence certified by a physician due to common illness. This data is at the individual (worker) level and contains information on month-year of birth, gender, start and end date of the sickness, the specific disease causing the sickness leave according to the International Classification of Diseases tenth revision (ICD-10), the province of residence, and the economic activity of the firm where the individual works (according to the CNAE classification[4]). With this information, we are able to obtain the outcomes of interest: (i) the total number of spells, and (ii) the mean duration of sick leave spells. At this point, it is worth mentioning that working with the full cases of sick leave episodes has advantages over self-reported data. For example, one limitation of survey data is measurement error, that in this case, can lead to an under/over reporting of the number of lost work days. The data we have available does not suffer from this problem.

To give answer to our question, from the total number of sickness absences, we only keep sick leave episodes coming from workers registered under the "general contribution scheme" and exclude other categories, for instance, those who are self-employed.[5] We also focus on workers between 16 and 65 years old, which leaves us with a total of 25,586,406 sick leave spells. Due to the richness of the data, we are also able to distinguish between men and women as well as the economic sector to which employees belong. With the information available, we are able to unequivocally distinguish between three sectors: (i) public administration and defense, which includes exclusively public employees; (ii) education, that includes all the staff of pre-primary, primary schools, high schools, tertiary education and other education-related workers; and (iii) health, including all hospital activity, long

[4]CNAE is the standard economic activity classification used in Spain. We have this code at the 5 digit level.

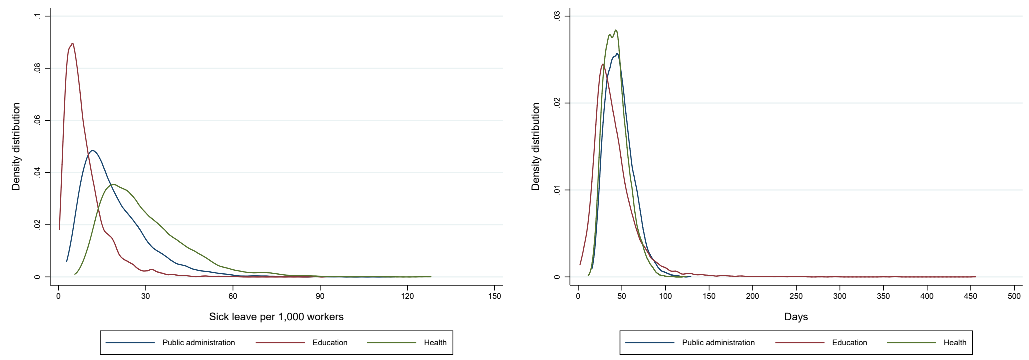
[5]The other categories correspond to domestic workers and workers in the mining and agricultural sector.

term care workers and other health support employees.

To construct the first outcome of interest, the total number of sickness spells, we need to take into account that unemployment in the Spanish labor market is very seasonable, and that the number of individuals employed has changed considerably during the period of analysis. For this reason, we collapse (aggregate) the data at the province month-year level and divide it by the total number of "employed" workers in each economic activity during the second trimester of the corresponding year. In this sense, the unit of observation is at the province-month level. We derive the number of "employed" workers for each sample from the Spanish Labor Force Survey. For the second outcome, (mean) duration of sick leave spells, as it is conditional on the existence of a sickness episode, we do not need to divide it by the corresponding working population.

In Figure 4.1, we plot the kernel density estimates for (a) the total number of sickness spells per 1,000 workers, and (b) the mean duration of sickness absences for the three economic activity sectors. From here, we see that the maximum height for the public administration sector is around 20 sick leave spells per 1,000 workers, for the education group is around 15 per 1,000 workers and for the health sector around 25 episodes per 1,000 workers. With respect to the average mean duration, the education sector reports the longest episodes compared to the other sectors.

Figure 4.1: Distribution of Sick Leave per Economic Sector



(a) # Sick leave per 1,000 workers

(b) Mean duration of sick leave

Notes: Own elaboration using administrative Spanish Social Security data which includes the universe of sickness absence of employees under the "general contribution scheme" for years 2009-2015. It also includes data from the Spanish Labor Force Survey on the total number of workers per economic sector activity.

In Table 4.1, we show descriptive information regarding the total number of spells and the mean duration of sick leave episode per economic sector, comparing women and men. As we see in panel (a), overall, women report more sickness episodes than men across the economic sectors under analysis. Regarding mean duration of sickness absences, individuals working in the public administration sector report the largest number (39.55 and 39.19 days form women and men, respectively).

Table 4.1: Descriptive Statistics of Sick Leave Spells by Economic Activity and Gender.

Panel A. Number of spells			
Economic sector	**All**	**Female**	**Male**
Public Administration	2,324,525	1,357,086	967,439
Education	1,080,380	848,166	232,214
Health services	3,642,089	2,974,555	667,534
Other sectors	18,539,412	9,139,130	9,400,282
All sectors	**25,586,406**	14,318,937	11,267,469
Panel B. Mean duration (in days)			
Economic sector	**All**	**Female**	**Male**
Public Administration	39.40	39.55	39.19
Education	33.26	33.29	32.12
Health services	35.91	35.96	35.70
Other sectors	36.85	36.93	36.76
All sectors	36.79	36.76	36.83

Notes: Own elaboration with data from the Spanish social security administration on the universe of sickness absence due to common illness, years 2009 to 2015. Average mean duration corresponds to a sickness spell that started in month, m, year t.

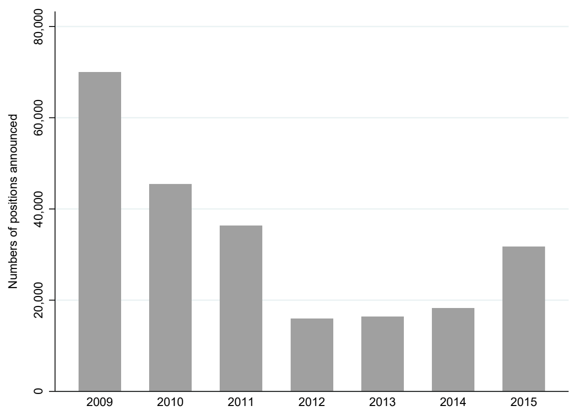
4.3.2 Public Examination Data

Regarding public examination data, we built this database using administrative records from the Ministry of Territorial Policy and Public Function in Spain, from 2009 until 2015. This database contains the full information on public employment calls of the General State Administration, the Autonomous Communities (from now on, regions), the European Union, International Organizations and Local Administration. The latter includes the corresponding calls of the provincial councils, provincial capitals, municipalities of more than 100,000 inhabitants and those who call at least 3 places, regardless of the number of inhabitants. The General State Administration includes positions offered by the ministries, the social security institution, armed forces, justice administration. The Autonomous Communities include positions in the education sector, health care institutions, justice administration and other administrative offices, and the Local level compiles all calls from the municipalities and other local administrations.

In this database, the unit of observation is the call for which we have information on the date of publication, whether the position is for the international, national, regional or local administration, the type of contract of the position, the minimum educational degree required, the convening institution, the number of positions offered, the region-province of workplace and the date of the examination. With this information, and according to the institution that offers the places, we assign the CNAE code (economic activity sector) to every call in the dataset. Thus, we generate three categories: public administration and defense, education, and health.[6]

The universe of positions includes permanent and temporal positions, and also those offered at the European Union level, which gives a total of 263,714 positions. However, for the purpose of our analysis, we only keep those offered inside Spain and those whose access is under the "public examination" system. Specifically, we exclude temporal positions as they do not require a public exam to access the place,[7] which leaves us with 234,123 positions offered across the country. Figure 4.2 shows these positions, from here, we can see that 2009 was the year with the the most number of places called, then, we see a decrease in the number, and a new increase from year 2014 onwards.

Figure 4.2: Number of Positions Announced in Spain



Notes: Own elaboration using administrative data from the Ministry of Territorial Policy and Public Function in Spain, for years 2009-2015. It includes all positions that require a passing exam as entry condition in which the place of work is Spain. Excludes all temporal contracts.

To test whether public examinations influence sick leave behavior, we need the examination date. However, due to data restrictions, we do not have this information for all the calls. In particular, this information is mainly missing for the local positions. In Table 4.2, we provide details on the percentage of missing data according to the administrative level that offers the civil servant positions.

We can see that for positions offered by the State Public Administration (National on the table), there is no missing information. For the regional ones, 442 positions have missing information on the examination date, which corresponds to a 0.38% while, for local calls, there is a 59.98% of missing dates. Overall, the missing

[6]The composition of the these three economic activities is the same as in the sick leave data.

[7]In any case, we have verified that all temporal positions excluded, indeed, did not have a passing exam as entry requirement

Table 4.2: Missing Examination Date

Adm. Level	# missing	# of positions offered	% missing
National	0	96,861	0.00
Regional	442	115,267	0.38
Local	13,193	21,995	59.98
Total	13,635	234,123	5.82

Elaboration: The Authors using data from the Ministry of Territorial Policy and Public Function in Spain, years 2009-2015.

information is 5.82% of the sample. In spite of the percentage seen in the local positions, we have to keep in mind that the unit of observation in the database is the call, which can include from 1 to *n* positions. This is important because, as we will explain in the empirical strategy section, it is not the same a call with one position than one with, for instance, 500 (or more) positions. In this sense, from the 13,193 positions with missing examination date, a 49.10% corresponds to calls in which just one position was announced, and a 92.57% to up to 10 work places.

Finally, to have an idea of the cross-sectional variation in the number of positions at the regional level, in Figure 4B.1 of the Appendix 4B, we depict the number of positions offered per 1,000 workers for the year 2009, which registered the maximum number of new workplaces. This map includes the positions for which we have information on the examination date. We excluded the two autonomous cities, Ceuta and Melilla, just for graphing purposes. In addition, Table 4A.1 of Appendix 4A shows the number of positions per year per region during the period of analysis.

4.4 Estimation Strategy and Identification

We are interested in the impact of public examinations on sick leave behavior. To do so, we focus on regional and local calls as in these ones we can identify the workplace at the province level.[8] Thus, we collapse and link sick leave with public examination data at the province, month and year of the examination level. We, therefore, get a balanced panel of provinces where the unit of observation is the province month-year, with one observation per province and month (for each year) counting the total number of spells, its mean duration, and number of positions announced. From here, we derive 3 samples: (i) the public administration sample, with data exclusively for individuals working in the public sector; (ii) the education sector, that contains all sick leave spells and civil servant positions of this sector, as previously defined; and (iii) the health sector. With this sample definition, we are

[8]For the national calls it is not possible to know the geographical workplace.

left with 123,627 positions distributed across the 52 provinces of Spain to identify the effect of new call for positions on sick leave.

Formally, we exploit variation on the number of positions having an exam in province p, month m, year t. We measure this impact through two outcomes of interest: (i) the total number of sickness absences (per 1,000 workers), and (ii) the (average) length of the sickness episode conditional on taking it. As individuals who are currently working may be affected in terms of increased health-related work absence, we expect to see an increase in the total number of sickness absences before the examination date. At this point, as the reader might wonder, and due to data constraints, we do not really know whether people participating in public examinations and taking sick leave are actually the same. Therefore, our estimates are an intent-to-treat-effect.

We continue our analysis using an event study design that allows us to see how sick leave behaves before the examination date. We estimate the following baseline equation:

$$Y_{pmt} = \alpha + \sum_{j=-6}^{4} D_{pmt} + \delta_p + \gamma_m + \rho_t + \epsilon_{pmt} \tag{4.1}$$

where Y_{pmt} is the outcome of interest in province *p*, month *m*, year *t*. Our independent variables of interest are the set of dummies (0-1) that capture how the competition generated by public examinations influences sick leave behavior months before the examination date. We include seven pre-examination dummies, to compare how the two measures of sick leave vary, one month, two months, up to seven months before the examination date. We use seven months before the exam taking into account the average time between the calls and the examinations. We also include province fixed effects (δ_p) for time-invariant heterogeneity at the province level, month and year effects (γ_m, ρ_t) for common shocks across provinces, and cluster standard errors at the province level. Given that the outcomes are evaluated at the province level, and that provinces have different population sizes, all regressions are weighted by the number of employees per province. We perform separate regressions for each sub-sample, the public administration, education, and health economic sectors.

Literature on event study designs suggests to set the reference period (normalization) to one period prior to the treatment (Alsan and Wanamaker, 2018; Freyaldenhoven et al., 2019; Fuest et al., 2018; Schmidheiny and Siegloch, 2022). In our setting, the "treatment" is the competition (or need to study for the exams) generated by public examinations. In this context, it is not straightforward to establish a normalization period, thus, we decide to set this reference category at the pre-7 (7 months before the exams) month taking into account the average time between the publication of the call and the examination date. In addition, we follow the methodology proposed by Schmidheiny and Siegloch (2022) and estimate the event study with binned endpoints for multiple events of identical size.

The next step to build our identification strategy is to decide which calls we are going to use in our event(treatment) definition. Thus, for our baseline estimates, we focus on the largest calls, which are those that have a number of positions equal or greater than the 99 percentile of the "positions" distribution.[9] The decision to focus on a large number of positions is because sick leave behavior might not respond to a small number of workplaces announced, as preparation requires big efforts in terms of study, and, more importantly, the perceived probability of "wining the place" is higher when more positions are available. Therefore, participants might have more incentives to increase the demand of time to study, which in turns, might affect health resulting in an increase in absence rates. In this way, our event variable is a dummy taking the value of 1 when the number of positions (in province p, month m, and year t) in the call is greater or equal to the 99 percentile of the positions distribution.[10] Figure 4B.2 of the Appendix 4B shows the distribution of large events per 1,000 workers split by economic sector.

Finally, we assign the workplace location at the province level for every call. In the case of local calls, we have the information to which province the position belongs (in other words the workplace province), so they are directly assigned to each province in the panel. However, for the regional calls, we know the "autonomous community" to which the workplace can be allocated, therefore we assign the number of positions in every regional call to each province that comprises the "autonomous community",[11] since individuals have higher probabilities of participating in calls that are within their region of residence.

[9] As a robustness check, in a following section, we conduct the baseline estimates using as treatment definition the calls with a number of positions equal or greater than the 95 percentile of the positions distribution.

[10] In the robustness check section, we present estimates where the event variable is constructed taking into account all the positions offered (not only the largest calls).

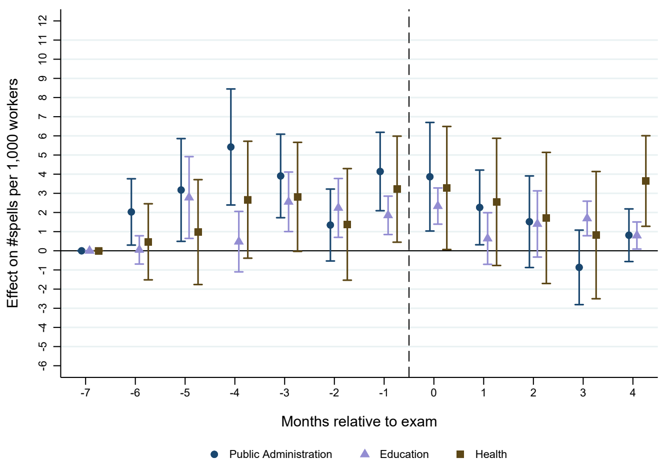
[11] For example, if there is a call of 100 positions in the health sector in the region of Catalonia, then we assign the 100 positions to the four province that comprise the region, that is Barcelona, Girone, Lleida and Tarragona.

4.5 Results

4.5.1 Effects on the Number of Sick Leave Spells

We start the analysis of the impact of public examinations on sick leave behavior by plotting the event study specification of Equation (4.1) for the first outcome, the number of spells per 1,000 workers (per economic sector). Figure 4.3 shows the point estimates along with their 90 percent confidence intervals for each sample according to the economic sector. In the graph, each series represents a different regression.

Figure 4.3: Effect of Public Examinations on Sick Leave Spells: Results by Economic Sector



Notes: Figure plots event study estimates and 90 percent confidence intervals of different specifications of Equation (4.1). Treatment definition is according to the 99 percentile of the positions distribution. Dependent variable is the total number of spells per 1,000 workers for each economic sector. The reference period is pre-7 (seven months before the examination). Standard errors are clustered at the province level.

Looking at the estimates of the public administration sector, we find a positive and significant increase in the number of spells per 1,000 workers several months before the public examinations. We see an increase of around 4 sickness spells (per 1,000 workers) in the same month of the exams. Looking backwards, we also find detrimental effects from one to six months before the tests. At first sight, the magnitude of the impact seems to be larger for this sector than those of the other economic sectors. We also see a decrease in the size of the coefficient after the examination date. In all the pre-examination dates, all coefficients are of the expected signs.

Moving to the education sector, we also find a significant increase in the number of spells (per 1,000 workers) five, three, two, one months before, and in the month of the examination. The largest coefficient corresponds to three months before the exams, which is around 3 additional health-spells per 1,000 workers per month and province. In addition, after the public exams take place, the absence rate decreases. With respect to the health & social services sector, for the calls with the largest number of positions, we find a significant increase in the outcome of interest one, and in the month of the public exams. In terms of magnitude, overall, we find the largest effects in the public administration and education sectors, which suggests that individuals in these sectors might be the most affected.

An interesting finding is that we still see some significant variation after the exams. This result is compatible with the "syndrome of the public examination participant" ("sindrome del opositor", in Spanish) which is related to lasting increased levels of anxiety and stress even after the exams. In table 4A.3 of the Appendix 4A, we report the estimated results of the coefficients plotted in Figure 4.3.

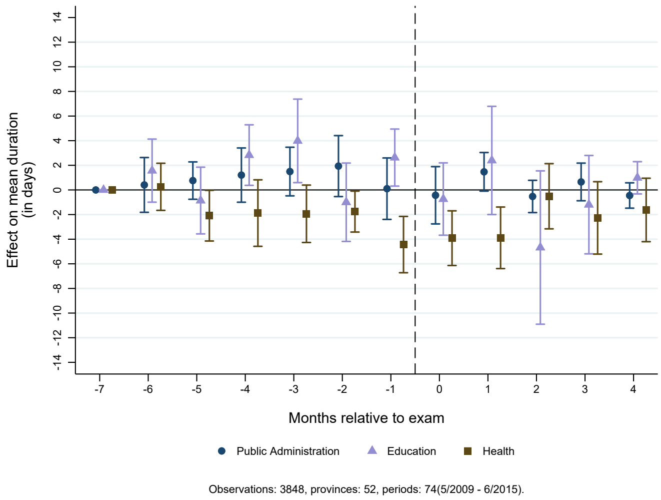
Our results point to an increase in the number of sick leave spells derived from the public examination system, which may be a (or one of the) negative consequence (externality) of the civil servant entry process. We find that in the public administration sector, sickness-related absences increase by up to 5 spells per 1,000 workers per month per province, in the education sector to up to 3 spells per (1,000) workers (per month per province), and in the health sector to a maximum of 3 spells per (1,000) workers (per month per province). With these results, one may also wonder whether stress/anxiety is playing a role behind the increase in sickness absences. We will discuss this possibility in the following sections of this chapter.

4.5.2 Effects on Length of Sick Leave Episodes

We continue exploring the impact of public examinations on sick leave behavior by analysing whether average mean duration of sickness episodes change before public exams. For this, we estimate Equation (4.1) where now the outcome of interest is the mean duration (in days) of sickness spells (per month, per province) conditional on having taken a sick leave. Figure 4.4 plots the average mean duration of sick leave spells for the three economic sectors. Likewise, in Table 4A.4 of the Appendix 4A, we show the estimated coefficients.

Beginning with the public administration sector, we do not see any significant variation in the average mean duration of sickness spells before the examination date. Turning to the education sector, we find that average mean duration increases several months before the exams. For instance, one month before the exams the mean duration of the sickness spells increases by around 2 days per month per province, and three months before, this measure increases by 3.9 days. The results for the education sector point to an overall increase in the average number of days that individuals take under sick leave, thought slightly significant. The strongest

Figure 4.4: Effect of Public Examinations on Sick Leave Duration: Mean Length in Days



Notes: Figure plots event study estimates with 90 percent confidence intervals of Equation (4.1). Treatment definition is according to the 99 percentile of the positions distribution. Dependent variable is the mean duration of sickness spells per economic sector. The reference period is pre-7 (seven months before the examination). Standard errors are clustered at the province level.

coefficient corresponds to three months before the exams which is compatible with the announcement of the process (call for positions).

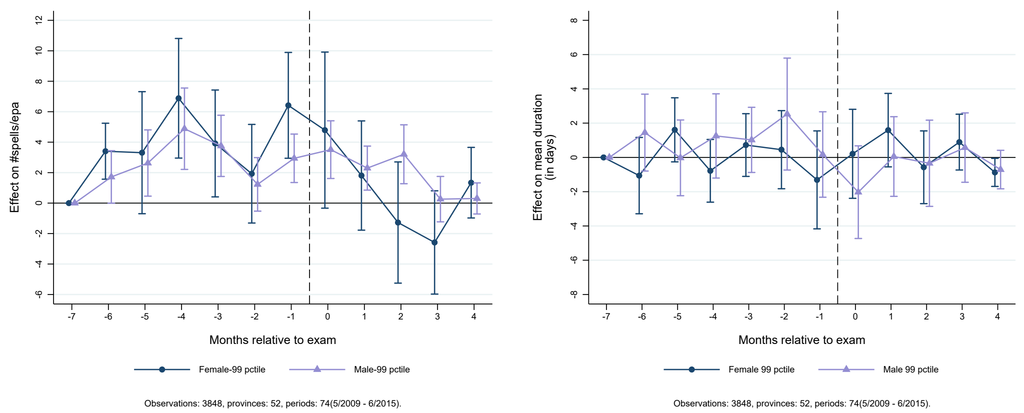
An interesting result is seen in the health & social services sector. Here, we find a statistically significant decrease in the mean duration of sickness spells. A possible explanation for these findings might be that the type of sickness absences occurring in this sector may be mainly coming from diseases that, on average, have a shorter mean duration such as digestive, respiratory or stress-related conditions in comparison with other conditions such as neoplasms or injuries that usually demand more days off. This issue will be further explored in upcoming sections.

Overall, we find evidence supporting that the civil servant recruitment process may be generating negative unintended effects in terms of sickness absences. We find that this effect is stronger for individuals working in the education sector. Sickness absences in the workplace might be linked to work time lost as well as to productivity losses. For example, Herrmann and Rockoff (2012), on a research focused on teachers in New York City, found that worker absences have large negative impacts in terms of productivity, mainly coming from employing a temporary substitute as well as from replacing a regular worker.

4.5.3 Gender Differences in Sick Leave

From the main results, one may also wonder whether there are differences in sick leave behaviour in terms of gender. To answer this question, within each economic sector, we split the sample between female and male, and re-estimate Equation (4.1) for the two samples: women and men workers using sick leave. Thus, the outcomes of interest are (i) the number of sickness absences per 1,000 female(male) workers in each economic sector,[12] and (ii) the mean duration of these sickness spells. We apply the same econometric strategy and define the treatment according to the 99 percentile of the positions distribution. We start with Figure 4.5 in which we plot the results for the public administration women/men employees.

Figure 4.5: Effect of Public Examinations on Sick Leave in the Public Administration Sector: Women vs Men



(a) Outcome: #Spells per 1,000 workers

(b) Outcome: Mean duration

Notes: Figure plots event study estimates and 90 percent confidence intervals of Equation (4.1) in two separate regressions for male and female. Dependent variables are (panel a) the total number of spells per 1,000 female(male) workers, and (panel b) mean duration of sickness absences for male/female in the public administration sector. The reference period is pre-7. Standard errors are clustered at the province level.

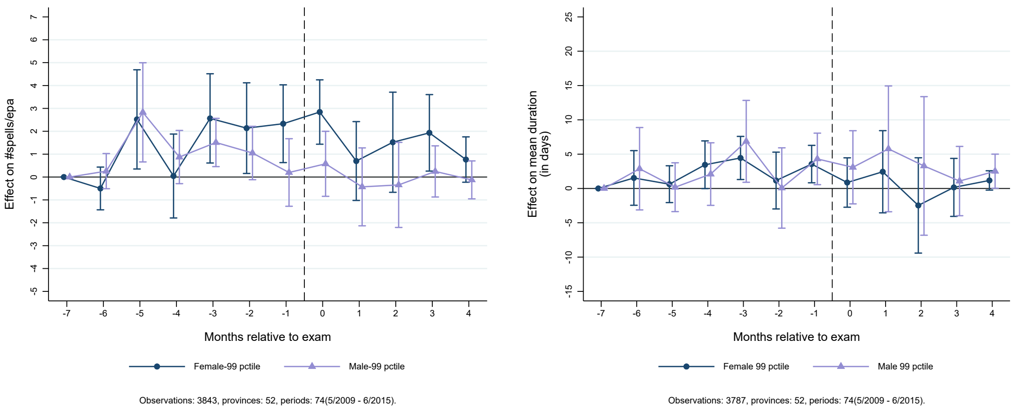
We find similar results for both men and women (for the two outcomes) in terms of direction and statistical significance, which is compatible with the idea that, in the public administration sector, both groups seem to react in a similar way to public examinations. For instance, female workers in provinces in which there is a large number of public positions show, on average, a higher number of sick-related absences of up to (around) 6 episodes (per 1,000 workers and month) in months preceding the public exams in comparison with women located in other provinces. In terms of the average mean duration, the results are consistent with baseline estimates in which we do not find much significant variation in the months preceding

[12]This implies that for this outcome we divide the numerator over the number of female(male) workers per economic sector.

public exams. Interestingly, after the examination date, we find some statistically significant variation for male workers, which might be compatible with the "sindrome del opositor" that we discussed in the previous section.

We next move to the education sector, and in Figure 4.6 plot the estimated results in separate regressions for women and men. For female workers, in panel (a), we see that in provinces with a large number of positions under examination women take, on average, a higher number of sick related absences in all the 6 months before the exams (except for the pre-4 month) relative to other provinces. We find the largest coefficient (around 4 spells per 1,000 female workers per province per month) three months before the examinations, which is compatible with the announcement of the positions in this sector. For the male group, we also find an increase in health related absences though of smaller magnitude and significance. Overall, it seems that in the education sector, female workers would be more affected.

Figure 4.6: Effect of Public Examinations on Sick Leave in the Education Sector: Women vs Men



(a) Outcome: #Spells per 1,000 workers

(b) Outcome: Mean duration

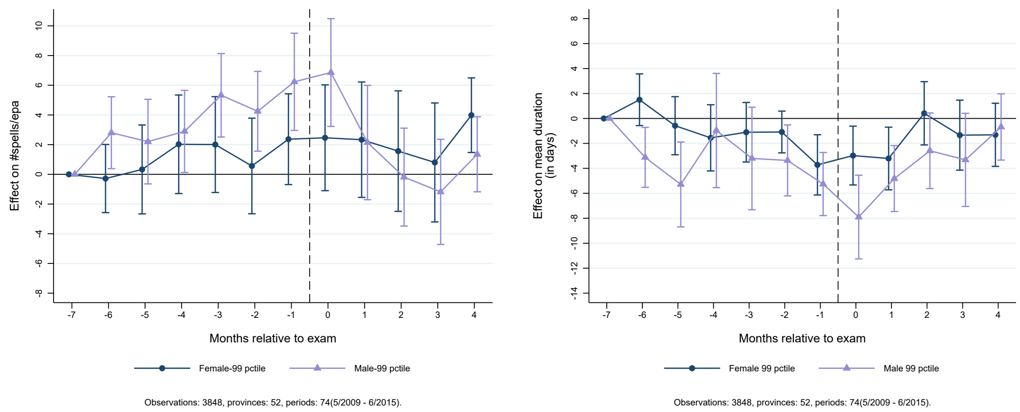
Notes: Figure plots event study estimates and 90 percent confidence intervals of Equation (4.1) in two separate regressions for male and female. Dependent variables are (panel a) the total number of spells per 1,000 female(male) workers, and (panel b) mean duration of sickness absences for male/female in the education sector. The reference period is pre-7. Standard errors are clustered at the province level.

Moving to the intensive margin, overall, it does not seem to be any change in the average mean duration of sickness spells, though we find some significant increase one month before the exams for women and men. In terms of magnitude, direction and statistical significance results are very similar for women and men.

With regards to how women and men's sick leave behavior responds to public examinations in the health & social services sector, in Figure 4.7, we show the estimated coefficients along with their 90 percent confidence intervals. For female workers, we do not find any significant variation in the number of sickness spells (panel a). For male workers the picture is different. For instance, we find an increase in the number of sickness episodes several months before the exams and the largest coefficient corresponds to that of the same month of the exam. After the examination date the coefficient drops and becomes insignificant.

Panel (b) shows the variation for the average mean duration of sickness spells. As in the baseline results, we observe a reduction in the average mean duration. This behaviour is similar for female and male workers thought the coefficients are larger for men. Overall, in this economic sector, we find evidence supporting that male workers might be more affected in terms of sickness absences.

Figure 4.7: Effect of Public Examinations on Sick Leave in the Health Sector: Women vs Men



(a) Outcome: #Spells per 1,000 workers (b) Outcome: Mean duration

Notes: Figure plots event study estimates and 90 percent confidence intervals of Equation (4.1) in two separate regressions for male and female. Dependent variables are (panel a) the total number of spells per 1,000 female(male) workers, and (panel b) mean duration of sickness absences for male/female in the education sector. The reference period is pre-7. Standard errors are clustered at the province level.

From the gender heterogeneity analysis, in general, we get mixed evidence on whether women (men) are more responsive than men(women) in terms of sick-related work absences. We find results supporting a gender-related heterogeneity across economic sectors and within sectors. For instance, in the public administration sector both groups are affected. In the education sector, we see a clearer and persistent increase in sickness absences for women months before the public examination. However, when we look at the variation in the average number of days there is not significant relationship. In the health sector, male workers seem to be more affected in terms of an increase in sickness spells in face to public examinations. Literature on how external factors affect sick leave behavior is very scarce and therefore limited evidence is available for gender heterogeneity. However, Eriksen et al. (2016), in a research where they assess workplace bullying and sickness absence, find that women have higher, persistent increases in long-term sickness absence and worse long-term health, pointing to different coping strategies between men and women. Finally, in the Appendix 4A, in Tables 4A.5, and 4A.6 we present the estimates obtained in this section.

4.6 Type of Diseases in Sick Leave Episodes

In this section, we explore the composition of sick leave spells with respect to the most common illnesses used in medical certificates. As mentioned in previous sections, the data contains sick leave episodes certified by a physician, including detailed classification of the type of disease according to the ICD-10 official classification, therefore, we are able to exploit this information to explore which type of diseases are mostly used to get the medical certificates. For this, we classify all diseases into twenty main categories based on the macro groups of the ICD-10 classification. In Table 4A.2 of the Appendix 4A, we show descriptive statistics per illness category.

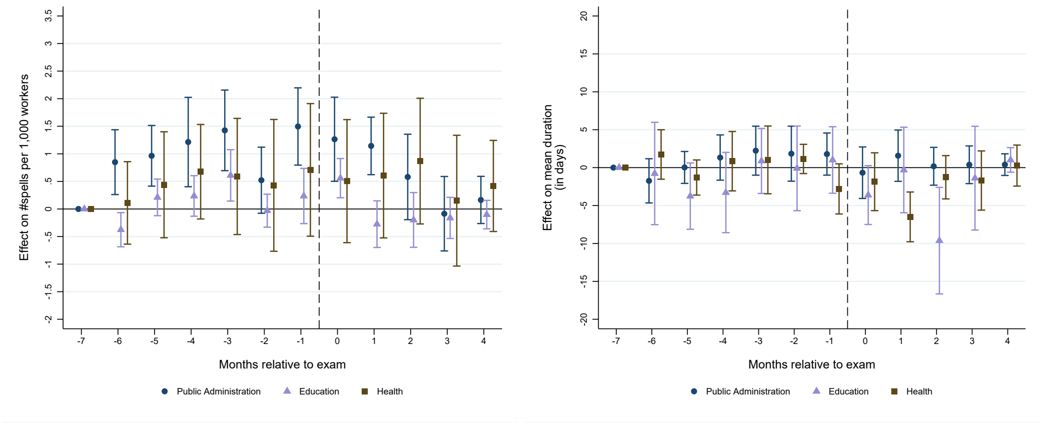
To begin with the analysis, in Figure 4B.3 of the Appendix 4B, we provide a set of plots of the distribution of average mean duration of sickness spells for selected illness categories and economic sectors. In addition in Table 4A.2 of the Appendix 4A, we show the average mean duration of the different diseases across the three economic sectors. In here, we see that conditions such as neoplasms, mental and blood disorders have the highest average mean duration (up to 102 days for neoplasms), and illnesses related to infectious and respiratory conditions report the lowest number of (mean) days (around 10 days).

In this exercise we expect to find a significant increase in the absence rate coming from diseases that are mostly related with stress and anxiety. Recall from the Introduction of this chapter that workers who are faced with time constraints and the need to study might reduce the allocation of time for sleeping, reduce the allocation of time for leisure, and increase stress related to civil service recruitment testing, which may increase the rate of illness. Under the same idea, we should not find any significant difference in sickness absences coming from diseases that are unrelated to stress or anxiety such as neoplasms (both benign and malignant tumors). We will analyse this later in the robustness check section.

To formally assess whether stress-related health conditions change, we estimate Equation (4.1) using sickness spells coming from mental disorders (example: anxiety, depression, panic attacks), osteoarticular (back pain, joint pain) and digestive conditions (non-infectious gastrointestinal problems). We include the same fixed effects, treatment definition, and standard errors as in our baseline equation. Figure

4.8 shows the results for this exercise plotting the estimated coefficients along with their 90 percent confidence intervals for the two outcomes of interest. Each series represents a separate regression for each of the three economic sectors. In the Appendix 4A, in Table 4A.7, we report the estimated coefficients and their standard errors.

Figure 4.8: Effect of Public Examinations on Sick Leave by Illness Category: "Stress-related Conditions"



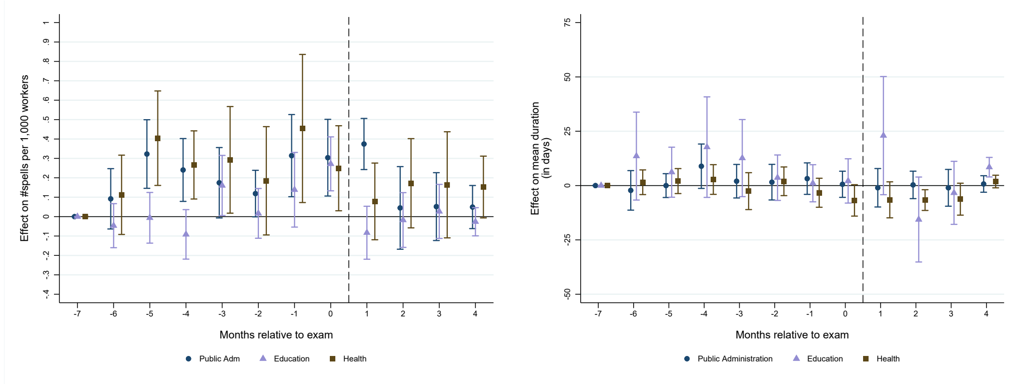
(a) Outcome: #Spells per 1,000 workers (b) Outcome: Mean duration

Notes: Figure plots event study estimates with 90 % confidence intervals of separate regressions for each economic sector. Dependent variables are (panel a) the total number of spells per 1,000 workers, and (panel b) mean duration of sickness absences. Reference period is pre-7. Standard errors are clustered at the province level.

Panel (a) shows a significant increase in the absence rate for workers of the public administration sector. This suggests that stress-related conditions increase before public examinations. The largest effect is seen one month prior to the exams, and is around 1.5 additional sickness absences per (1,000) workers per province and month. Moving to the education sector, most of the coefficients are of the expected sign though not statistically significant. However, we find that the largest (and significant) coefficients are three months and the month of the exam (as in the overall results for this sector). In this sector we also observe, that after the recruitment tests, the coefficient of interest decreases relative to the reference period (7 months before the exams). Finally, for the health & social services sector, at first sight, we do not find any significant variation in terms of the number of sickness spells for this type of medical conditions.

Panel (b) depicts the results for the mean duration outcome. Here, overall, we do not find differences in provinces with large calls relative to other provinces, which suggests that the average length of sickness episodes does not change in this group of medical conditions. We present a detailed table with descriptive statistics per economic sector in Appendix 4A, Table 4A.2.

Figure 4.9: Effect of Public Examinations on Sick Leave by Illness Category: "Mental Health Conditions"



(a) Outcome: #Spells per 1,000 workers (b) Outcome: Mean duration

Notes: Figure plots event study estimates with 90 % confidence intervals of separate regressions for each economic sector, for the sickness absences due to mental health conditions. Dependent variables are (panel a) the total number of spells per 1,000 workers, and (panel b) mean duration of sickness absences. Reference period is pre-7. Standard errors are clustered at the province level.

Next, we focus on sickness episodes coming exclusively from mental health conditions that mainly comprise anxiety and depression. For this, we re-do the above analysis using mental health diseases as outcome. Figure 4.9 depicts the estimated coefficients along with their 90 percent confidence intervals for the outcomes under analysis.

Overall from Figure 4.9, we find an increase in the number of sickness episodes in each of the three economic sectors. If we examine employees of the health sector, we find a sharp increase months before the public examinations. With respect to the education sector, we also find variation pointing to an increase in this type of health-related absences. In the public administration sector, we also see significant variation several months prior to the exams as well as in the month of the examinations. Moving to the average mean duration of these sickness spells, we do not find any significant variation. Results from this exercise, make us wonder whether applicants who combine work and exam preparation are experiencing negative unexpected consequences of the recruitment process in terms of increased levels of stress. We will come back to this point later on this chapter.

4.7 Robustness Checks

In this section, we include some robustness checks to add validity to our findings. First, we add support to the hypothesis that we established in the empirical strategy section, in which we argue that sick leave behavior might not respond to small number of workplaces announced. Second, we change the treatment definition and consider calls in which the number of positions is greater or equal to the 95 percentile of the positions distribution. Third, as a placebo test, we show that diseases that are less related to (short-term) stress/anxiety do not seem to be affected when there is a public examination approaching.

4.7.1 Continuous Event Variable

In our baseline results, we dichotomize the event variable based on whether the number of positions under examination is greater or equal to the 99 percentile of the positions distribution, under the assumption that sick leave behavior might not respond to a small number of workplaces announced as preparation requires big efforts in terms of study, and, more importantly, the perceived probability of "wining the place" is higher when more positions are available. To support this statement, in this subsection, we exploit all the available variation and estimate a generalized event study design as suggested by Schmidheiny and Siegloch (2022), where the event is constructed as a continuous variable measuring the number of positions per 1,000 workers for each economic sector. For the rest of the estimated equation, we use the same specification as in Equation (4.1).[13]

Figure 4.10 shows the results for this exercise. Panel (a) depicts the estimates for the public administration sector. In here, we see that we still find some significant effects in terms of increased number of sick leave episodes before the examination date. With respect to the education and health sectors, in panel (b), we see that we do not find any significant increase in sickness absence for the health sector. For the education sector, we only find a small increase in the outcome of interest three months before the examination date.

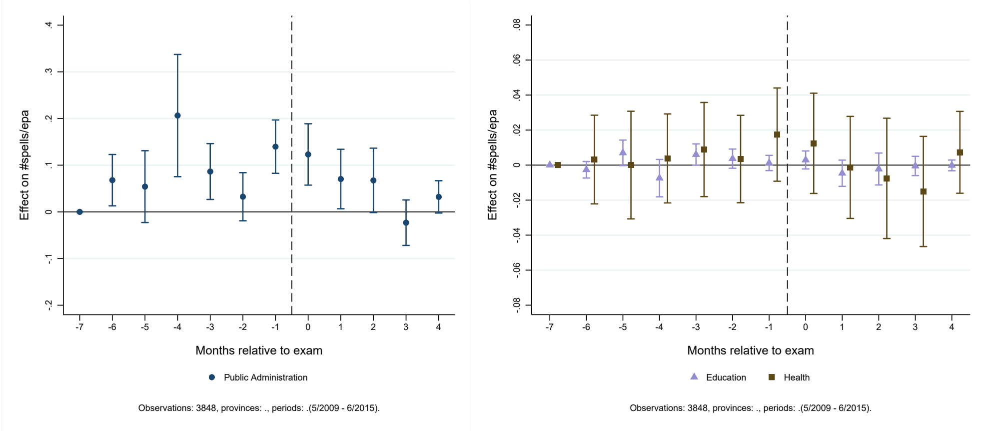
The results of this subsection show that even when we consider calls with a small number of positions, we still find some significant differences in the absence rate. However, we believe that by using only the largest calls (with a huge number of positions under competition), we are capturing in a more appropriate way the effects of public examination on sick leave behavior for the reasons already exposed.

4.7.2 Changing the Treatment Definition

In this robustness check we change the treatment definition used in the baseline analysis. Now, we consider that an event takes place when the number of positions under examination is greater or equal to the 95 percentile (instead of 99) of the po-

[13]We present the estimated results in Table 4A.8 of Appendix 4A.

Figure 4.10: Effect of Public Examinations on Sick Leave Spells: Continuous Event Definition



(a) Public administration sector (b) Education and Health sectors

Notes: Figure plots estimated coefficients of a generalized event study design where the event variable is continuous, along with their 90 percent confidence intervals. Panel (a) shows results for the public administration sector. Panel (b) presents results for the education and health sectors, separately. Dependent variable is the total number of spells per 1,000 workers in each economic sector. The reference period is pre-7 (seven months before the examination). Standard errors are clustered at the province level.

sitions distribution. Formally, we re-estimate Equation (4.1) for the two outcomes: (i) the number of spells per 1,000 workers, and (ii) the average mean duration.

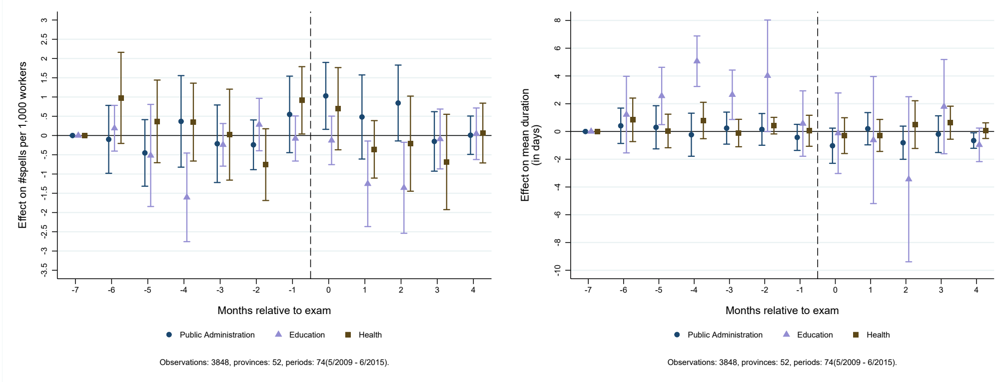
With this exercise, we expect to still find some statistically significant variation in sick leave behavior. Figure 4.11 shows that this is the case. For instance, for the absence rate (panel a), we get the expected signs in most of the coefficients and some significant variation (increase) in the previous and in the very month of the examinations. In addition, for the three economic sectors the estimated coefficients are smaller in magnitude compare to those of previous periods. Moving to the average mean duration of sickness spells, the education sector shows a significant variation, which can reach up to 5 more days on the length of average sickness absences per province per month. All the estimates from this subsection are shown in Table 4A.9 of the Appendix 4A.[14]

4.7.3 Placebo Illness Test

In this section, we perform a placebo test focusing on sickness spells with a diagnosis of "neoplasms". The intuition behind is that we should no expect to find any significant variation in terms of sickness absences, as it is very unlikely that participants both working and preparing an examination develop a tumor as consequence

[14]The rest of the results using the treatment definition of this subsection (gender and type of diseases analysis) are consistent and available upon request.

Figure 4.11: Robustness Checks - Different Treatment Definition



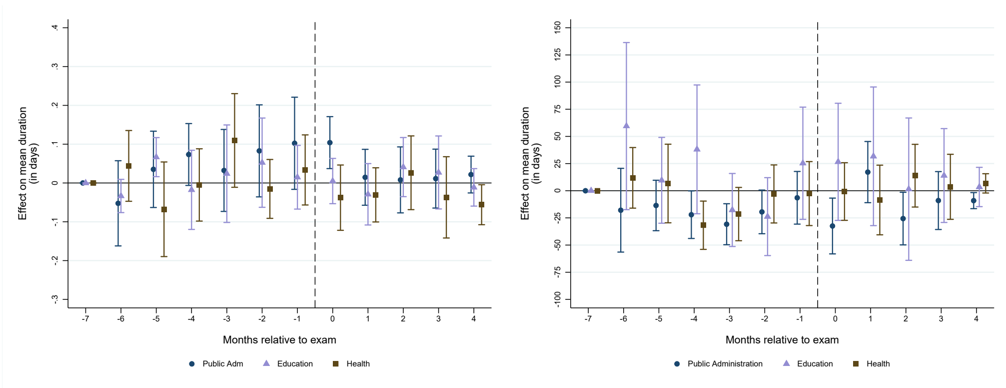
(a) Outcome: #Spells per 1,000 workers (b) Outcome: Mean duration

Notes: Figure plots event study designs along with 90 percent confidence intervals for sickness absences where the treatment definition is based on calls with a number of position greater or equal to the 95 percentile of the positions distribution. Dependent variables are (i) the total number of spells per 1,000 workers in each economic sector (separate regressions), and (ii) mean duration of sickness absence. The reference period is 7 months before the examinations. Standard errors are clustered at the province level.

of the tough recruitment process.

With this reasoning, we implement the same regression as in Equation (4.1) using as sub-sample all sickness episodes with a diagnostic of "neoplasms". In Figure 4.12, we plot how neoplasms-related sickness spells change months before (and in the same month of) public examinations. We perform separate regressions for the three economic sectors under study (public administration, education and health & social services) using the baseline treatment definition.

Figure 4.12: Effect of Public Examinations on Sick Leave Behaviour: Neoplasms



(a) Outcome: #Spells per 1,000 workers (b) Outcome: Mean duration

Notes: Figure plots event study designs along with 90 percent confidence intervals for sickness absences coming from neoplasms. Dependent variables are (i) the total number of spells per 1,000 workers (neoplasms-related) in each economic sector (separate regressions), and (ii) mean duration of these sickness absences. The reference period is 7 months before the examinations. Standard errors are clustered at the province level.

Regarding the extensive margin, in panel (a) of Figure 4.12, we do not find any statistically significant variation in the absence rate per 1,000 workers months before the exams in none of the three economic sectors. Likewise, in panel (b), we show the results for the average mean duration of these sickness spells. Once again, in general, we do not find any significant variation in this disease. We only find some significant changes, though of the contrary sign, in the public administration sector. Altogether, the results of this section add validity to our findings, which supports a significant increase in health-related absences several months before the examination date. The results seem to be mostly driven by stress related absences, which might be experienced by applicants who combine work and exam preparation. However, an additional possible explanation for our results might be that individuals may be using sickness absence as an strategic tool to gain more days to study exaggerating illnesses to increase sick leave. In this sense, employees could be more prone “to use” diseases that are more difficult to be objectively diagnosed, which is clearly not the case for the neoplasms. This also goes in line with some literature that finds that some type of medical conditions are more sensitive to incentives (for instance, changes in the payment scheme, insurance claims, job stability, etc) that may affect sick leave behavior (Godard et al., 2022; Bratberg and Monstad, 2015). Finally, we present all estimates from Figure 4.12 in Table 4A.10 of Appendix 4A.

4.8 Workers’ Health

Our main hypothesis is that the traditional recruitment system to access civil servant positions might be generating negative externalities in terms of increased health-related work absences. In this section, we provide some evidence that supports that large calls might be related to an increase in the use of health care services among public sector employees, which, in turn, might be compatible with deteriorating health. Formally, we test whether there is an association between large calls and some health (care) outcomes among public sector employees. We explore the existence of changes in self-assessed health and in health care use: general practitioner, emergency and specialist visits.

4.8.1 Public Administration Workers

We start comparing health-related outcomes of public sector employees with individuals working in other economic sectors. For this, we use data from the survey “Barómetro Sanitario” conducted by the *Centro de Investigaciones Sociológicas* of Spain for the years 2009 to 2015. This is a nationally representative survey that samples individuals aged 18 and older living in Spain. It provides information on health care use, waiting times, quality of services, self-assessed health and several socioeconomic characteristics at the individual level. It has several rounds per year.

We assess a set of outcomes such as general practitioner, specialists and emergency visits, in-hospital stay, self-assess health, and the number of general practitioner and specialist visits. For this, we use the corresponding questions from the survey asking whether individuals have used these services in the last twelve months. Using the responses to these questions, we construct the outcomes of interest, which are dummy variables taking the value of one for a positive response and zero otherwise. For the number of visits, these variables are conditional on having used the referred health services. As we did in the baseline estimates, we restrict the sample to individuals between 18 and 65 years old. The survey has important information that allows us to distinguish between public sector employees and other workers, which fits with our analysis.

To formally explore our hypothesis, we use a linear regression of the form:

$$Y_{ipt} = \alpha + \beta_1 Public_i + \beta_2 Exam_{pt} + \beta_3 Public_i * Exam_{pt} + \tau X_{ipt} + \gamma_p + \delta_t + \epsilon_{ipt} \tag{4.2}$$

where Y_{ipt} represents the health care use and health outcomes for individual i in province p and time t. $Public_i$ is a variable that takes the value of one for individuals working in the public sector and zero otherwise. $Exam_{pt}$ is a public examination indicator and takes the value of one for large calls (greater or equal to the 99 percentile of the positions distribution) in province p and time t. The coefficient of interest is β_3, the interaction term between public sector employees and the public exam indicator. We include a set of control variables τX_{ipt} at the individual level: sex, age group categories and a socioeconomic status category which is designed and provided by the survey. We also include province fixed effects γ_p to account for any time invariant shock across provinces and time δ_t fixed effects. Standard errors are clustered at the province level (52 clusters). We use probability weights to take into consideration the survey design.

Table 4.3 shows the estimates of Equation (4.2). We find a significant increase in the probability of having used the general practitioner and specialist for public sector workers living in a province with a large number of public examinations. We also find some variation in the number of specialist visits (significant at the 10% level). For the other outcomes, even we do not find a significant variation, the estimated coefficients are of the expected signs.

A possible interpretation for the increase in general practitioner visits, might be

Table 4.3: Health Care Use Results for Public Sector Workers

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	GP Visits	**Emergency**	**Specialist**	**Hospital**	**SAH**	**# GP visits**	**# Specialist**
Public	-0.0038	-0.0064	0.0530***	-0.0091*	0.0393***	-0.1003	0.1087
	(0.0123)	(0.0103)	(0.0074)	(0.0050)	(0.0107)	(0.0835)	(0.1038)
Exam	-0.0321	0.0480*	0.0152	0.0000	-0.0063	0.5294	-0.0935
	(0.0249)	(0.0257)	(0.0257)	(0.0128)	(0.0263)	(0.5391)	(0.2367)
Public*Exam	**0.1816*** **	**0.0570**	**0.2201** **	**0.0312**	**-0.0977**	**0.8323**	**1.8187***
	(0.0504)	(0.0811)	(0.1028)	(0.0433)	(0.1062)	(0.7259)	(0.9901)
Female dummy	0.1176***	0.0568***	0.1243***	0.0277***	-0.0438***	0.7864***	0.3357***
	(0.0094)	(0.0057)	(0.0057)	(0.0028)	(0.0084)	(0.0780)	(0.0944)
Age categories	x	x	x	x	x	x	x
SES categories	x	x	x	x	x	x	x
Province FE	x	x	x	x	x	x	x
Time FE	x	x	x	x	x	x	x
Observations	34,684	34,699	34,684	34,692	29,664	23,454	14,599

Notes: OLS estimates of Equation (4.2). Outcomes “GP visits”, “Emergency”, “Specialists”, “Hospital” are dummy variables assessing health care use at different levels of care. Outcome “SAH” is a dummy variable measuring self-assessed health (Good/Bad) for which there is no data for year 2009. “# GP visits” and “# Specialist visits” are the number of visits to the general practitioner and specialist, respectively. Standard errors are clustered at the province level (52 clusters). P-weights used in estimation to account for the survey design. ***p<0.01, **p<0.05, *p<0.10.

that individuals need to visit the physician to get the certificate to justify the sick leave, which can be interpreted as an immediate consequence of the public examinations. Regarding the probability of having visited the specialist, it also increases as well as the number of visits to this level of health care, which means that individuals already visiting the specialists, now increase the number of visits. This is compatible with deteriorating health for public sector employees compared with private sector workers in the presence of a large number of public examinations.

One concern about the previous results is that we should not find that health of permanent public sector workers is affected by opening of new positions. In fact, a permanent worker should not be affected, but a temporary one may fear for the current position, or getting uncertain about the future. To rule out that this is not the case, we now restrict the sample to individuals working exclusively to the public sector and estimate a model similar to Equation (4.2) where, instead of having a public employee indicator independent variable, we have a dummy taking the value of one for temporal workers and zero for permanent employees (both in the public sector). In this exercise, we expect to find negative health-related outcomes for individuals under temporal contracts in face of opening positions. We perform the analysis for the same outcomes and use the same specification as in Equation (4.2). Table 4.4 shows the results of this exercise.

Overall, we find the expected signs for all outcomes. In addition, we get a significant increase in the probability of visiting the emergency room for temporal

Table 4.4: Health Care Use Results for Public Sector Workers: Temporal vs Permanent Contracts

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	GP Visits	**Emergency**	**Specialist**	**Hospital**	**SAH**	**# GP visits**	**# Specialist**
Temporal	0.0389*	-0.0072	-0.0280	0.0080	0.0269	-0.1054	-0.1527
	(0.0219)	(0.0257)	(0.0209)	(0.0153)	(0.0211)	(0.2508)	(0.3118)
Exam	0.0595	0.0342	0.2109	-0.0186	-0.0620	0.8003	0.7302
	(0.0876)	(0.0920)	(0.1285)	(0.0518)	(0.0709)	(0.9630)	(1.0954)
Temporal*Exam	**0.0844**	**0.5851*** **	**0.1449**	**0.1806**	**-0.7638*** **	**0.3429**	**0.4736**
	(0.1030)	(0.1863)	(0.2239)	(0.1912)	(0.0735)	(1.3002)	(1.3397)
Female dummy	x	x	x	x	x	x	x
Age categories	x	x	x	x	x	x	x
SES categories	x	x	x	x	x	x	x
Province FE	x	x	x	x	x	x	x
Time FE	x	x	x	x	x	x	x
Observations	3,229	3,235	3,234	3,234	2,766	2,203	1,619

Notes: OLS estimates of a sample of public sector employees. Outcomes "GP visits", "Emergency", "Specialists", "Hospital" are dummy variables assessing health care use at different levels of care. Outcome "SAH" is a dummy variable measuring self-assessed health (Good/Bad) for which there is no data for year 2009. "# GP visits" and "# Specialist visits" are the number of visits to the general practitioner and specialist, respectively. Standard errors are clustered at the province level (52 clusters). P-weights used in estimation to account for survey design. ***p<0.01, **p<0.05, *p<0.10.

public sector workers relative to permanent workers. Likewise, we find that temporal workers in provinces with a large number of public examination report worst (self-assessed) health compared to their counterparts.

4.8.2 Gender Differences in Public Administration Workers' Health

In this section, we analyse public administration workers' health separately for women and men. For this, we use the same survey data from the "*Centro de Investigaciones Sociológicas of Spain*" and perform a regression of the form of Equation (4.2) for two different sub-samples: men and women.

Table 4.5 shows the results. Panel (A) suggests that the increase in the probability of using the specialist services might be coming from female public sector employees. We also see, that female workers of the public administration sector experience an increase of around 4 additional visits to the specialist per quarter per province in face of public examinations. An interesting result for women is the negative significant result in the probability of reporting bad health compared to their female counterparts working in the private sector when there are public examinations. In addition, we also observe some variation (though at the 10% significance level) for both general practitioner visits and the number of visits.

In panel (B), we present the results for the male sample. We find that this group

Table 4.5: Health Care Use Results: Women vs Men

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	GP Visits	**Emergency**	**Specialist**	**Hospital**	**SAH**	**# GP visits**	**# Specialist**
A. Female sample							
Public	0.0056	-0.0120	0.0395***	-0.0035	0.0334**	0.0858	0.2257
	(0.0153)	(0.0156)	(0.0129)	(0.0096)	(0.0144)	(0.1709)	(0.1633)
Exam	-0.0301	0.0886*	-0.0270	-0.0020	-0.0031	0.1737	0.0366
	(0.0367)	(0.0457)	(0.0342)	(0.0234)	(0.0306)	(0.3543)	(0.3136)
Public*Exam	**0.1508***	**0.0701**	**0.3254****	**0.0846**	**-0.4006****	**1.7692***	**4.6044*****
	(0.0802)	(0.2286)	(0.1374)	(0.1509)	(0.1902)	(0.9118)	(1.5475)
Observations	15,980	15,990	15,986	15,992	13,655	11,835	7,895
B. Male sample							
Public	-0.0131	-0.0042	0.0615***	-0.0127	0.0462***	-0.2875	-0.0255
	(0.0155)	(0.0114)	(0.0139)	(0.0095)	(0.0136)	(0.1931)	(0.1667)
Exam	-0.0338	0.0152	0.0516	0.0045	-0.0128	0.7317	-0.2745
	(0.0393)	(0.0393)	(0.0451)	(0.0173)	(0.0394)	(1.0505)	(0.2875)
Public*Exam	**0.2104*****	**0.0703**	**0.1738**	**0.0144**	**0.0913**	**0.4855**	**0.5762**
	(0.0700)	(0.1308)	(0.1176)	(0.0556)	(0.0874)	(1.0774)	(0.9404)
Observations	18,704	18,709	18,698	18,700	16,009	11,619	6,704
Control variables							
Age categories	x	x	x	x	x	x	x
SES categories	x	x	x	x	x	x	x
Fixed effects							
Province FE	x	x	x	x	x	x	x
Time FE	x	x	x	x	x	x	x

Notes: OLS estimates of Equation (4.2). Panel A corresponds to the female sample and Panel B to the male sample. Outcomes "GP visits", "Emergency", "Specialists", "Hospital" are dummy variables assessing health care use at different levels of care. Outcome "SAH" is a dummy variable measuring self-assessed health (Good/Bad) for which there is no data for year 2009. "# GP visits" and "# Specialist visits" are the number of visits to the general practitioner and specialist, respectively. Standard errors are clustered at the province level (52 clusters). P-weights used in estimation to account for survey design. Control variables and fixed effects are applied to both the female and male samples. ***p<0.01, **p<0.05, *p<0.10.

might be driving our results of increased visits to the general practitioner. For instance, we find an increase in the probability of using this service by 21.04 percentage points for male working in the public administration sector relative to private workers when there is a large number of public examinations.

Overall, we find evidence that supports some differences between the female and male samples. We show that women might be experiencing conditions leading them to make more use of the specialist care. Finally, we provide evidence supporting that there might be a deterioration of workers' health in face of public examinations which could be due to concerns about working stability or a general physical/mental overload due to the recruitment system.

4.9 Conclusions

Public examinations is the most common recruitment system to select civil servants in several countries. For instance, in Europe it is used in France, The UK, Spain, among others. Likewise, it is also used in some countries of Asia and South America. In spite of its wide use, there is scarce research conducted with respect to any potential unexpected negative consequences of the recruitment process. On this point, Bagüés (2005) in a document assessing the determinants of public exams success, highlights as a drawback the great impact on society in terms of income search costs due to the great effort needed to pass the exams. And additional example is a working document conducted by the World Bank done for south-pacific Asian countries, in which they found that the traditional public examination system induces corruption to get the positions. In this sense, assessing whether the system generates negative externalities in terms of health-related work absence results relevant for governments to decide future improvements in the selection mechanism.

In this chapter, we contribute to the literature on civil servants recruitment process assessing the impact of new openings of civil servant positions on sickness absences. For this, we use a unique administrative data from Spain on the universe of sickness absences due to common illness certified by a physician, and on the universe of public examinations, both for years 2009 to 2015. Empirically, we estimate an intention-to-treat effect exploiting the variation coming from the different number of positions under examination across time and provinces. We find a significant increase in health-related absences several months before the examination date. This effect is stronger for individuals working in the educational sector as well as for calls offering a large number of positions.

Next, we explore whether there are differences between women and men within each economic sector. We find that, overall, women seem to be taking more sick leave (in terms of number of episodes) several months before the exams. This is particularly visible in the education sector, which could be explained by the fact that the education sector is women-predominant in Spain. For the health sector the impact is mostly seen on men employees. In addition, we analyze which type of diseases are driving the results, and find that the effect is mostly coming from stress-related absences.

Our results provide evidence supporting the idea that the public examination process may be generating negative externalities in terms of increased sickness absences. We go a step further and find that stress-related conditions may explain these changes.

In addition, we explore health and health care use outcomes of public sector employees. We find evidence that supports a short term (self-assessed) health deterioration in temporal public sector workers compared to permanent civil servants in face of public exams. Moreover, we find that public sector workers have higher probabilities of visiting the general practitioner as well as the specialist in face of public exams, which is consistent with a deterioration in public sector workers' health. With this, we support the idea that the increase in sickness absence in the months previous to the examinations could be actually a true negative health effect due to an increase in stress among applicants who combine work and exam preparation. However, we have to recognize that there is still room for our results to be coming from a strategic behavior from participants.

Our results are important from a policy perspective as they highlight the existence of important negative consequences of the civil service recruitment process, which should be taken into account in future designs of improvement of the process. In this chapter we find evidence supporting that participants might be facing increased levels of stress and anxiety.

Appendix 4A: Additional Tables

Table 4A.1: Total Number of Positions Announced by Region

Region	**2009**	**2010**	**2011**	**2012**	**2013**	**2014**	**2015**
Andalusia	6,873	4,137	4,335	8	592	446	3,088
Aragon	2,838	590	1,074	129	546	700	1,128
Cantabria	745	352	14	53	90	12	87
Castille-La Mancha	5,547	627	0	0	210	0	330
Castille and Leon	1,679	1,179	901	0	176	0	895
Catalonia	5,197	8,511	1,849	90	0	88	370
Ceuta & Melilla	72	18	0	5	1	10	0
Community of Madrid	5,556	1,564	497	3,765	445	387	1,545
Navarre	1,066	1,069	308	172	9	0	13
Valencian Community	4,011	1,297	1,976	0	303	3	0
Extremadura	910	973	1,752	52	676	0	182
Galicia	3,691	998	439	58	481	325	802
Balearic Islands	2,966	1,069	245	0	18	0	97
Canary Islands	805	1,430	850	421	395	2,761	208
La Rioja	325	209	48	0	80	0	50
Basque Country	821	1,010	3,673	828	0	0	1,032
Principality of Asturias	796	136	197	448	24	52	188
Region of Murcia	1,931	581	0	0	214	0	0
Total	45,799	25,750	18,158	6,029	4,260	4,784	10,015

Elaboration: The Authors. Source: Administrative data from the Ministry of Territorial Policy and Public Function in Spain, for years 2009-2015. Includes all positions offered at the regional level that require a passing exam as entry condition, for which we have the information on the exam date.

Table 4A.2: Sick Leave by Disease Category and Economic Sector: Descriptive Statistics

	Public Administration			**Education sector**			**Health sector**		
Main disease category	**# of Spells**	**% Spells**	**Mean duration**	**# of Spells**	**% Spells**	**Mean duration**	**# of Spells**	**% Spells**	**Mean duration**
Osteo-articular system	467,512	20.11	52.13	181,225	16.77	45.09	661,068	18.15	53.14
Respiratory system	403,019	17.34	11.20	226,503	20.97	9.48	607,888	16.69	10.33
Infectious diseases	193,386	8.32	8.64	101,612	9.41	7.70	363,008	9.97	7.34
Injuries	188,603	8.11	50.01	72,899	6.75	45.67	238,497	6.55	50.15
Digestive system	143,358	6.17	27.48	59,068	5.47	22.84	240,595	6.61	20.57
Mental disorders	163,447	7.03	81.35	71,848	6.65	72.13	251,004	6.89	74.98
Senses and nervous system	106,542	4.58	37.81	45,447	4.21	30.98	194,106	5.33	32.09
Genitourinary system	59,277	2.55	35.86	29,555	2.74	32.50	105,734	2.90	34.12
Pregnancy complications	55,073	2.37	54.32	44,366	4.11	49.02	103,446	2.84	55.70
Circulatory system	49,054	2.11	72.00	19,483	1.80	58.06	72,934	2.00	58.64
Neoplasms	42,698	1.84	103.26	21,452	1.99	104.42	64,988	1.78	101.61
Supplementary factors(code V)	40,822	1.76	46.27	21,817	2.02	45.17	63,117	1.73	45.57
Skin diseases	28,883	1.24	30.82	10,229	0.95	25.88	41,194	1.13	29.64
Procedures (code P)	29,995	1.29	53.99	10,879	1.01	45.69	47,030	1.29	50.77
Endocrine diseases	16,105	0.69	49.02	6,317	0.58	48.45	25,347	0.70	50.12
External causes of injuries (code E)	8,621	0.37	47.98	3,513	0.33	40.40	12,619	0.35	42.10
Blood diseases	4,343	0.19	71.17	2,162	0.20	61.21	7,538	0.21	64.50
Congenital anomalies	3,548	0.15	63.81	1,611	0.15	55.86	6,068	0.17	60.40
Perinatal period diseases	1,770	0.08	67.16	984	0.09	55.01	2,542	0.07	67.13
Diseases not well defined	214,199	9.21	29.97	106,764	9.88	25.03	372,023	10.21	24.79
No diagnosis	23,128	0.99	73.46	8,153	0.75	55.42	32,957	0.90	59.36
Missing data	81,142	3.49	35.99	34,493	3.19	42.29	128,386	3.53	31,23

Elaboration: The Authors. Source: Administrative data from the Spanish Social Security Institution for years 2009-2015, and from the Ministry of Health, Consumption and Wellbeing (Ministerio de Sanidad, Consumo y Bienestar Social) on disease classification, available at `https://eciemaps.mscbs.gob.es/ecieMaps/browser/index_9_mc.html`.
Notes: Mean duration is measured in days.

Table 4A.3: Effect of Public Examinations on Sick Leave Behavior: (#) of Sick Leave Spells by Economic Sector

Economic sector	**Public Adm.**	**Education**	**Health**
	# Spells (1,000 workers)	# Spells (1,000 workers)	# Spells (1,000 workers)
Exam date_Pre,t-6	2.0287*	0.0454	0.4688
	(1.0524)	(0.4489)	(1.2071)
Exam date_Pre,t-5	3.1744*	2.7804**	0.9755
	(1.6306)	(1.2974)	(1.6643)
Exam date_Pre,t-4	5.4190***	0.4782	2.6664
	(1.8409)	(0.9596)	(1.8557)
Exam date_Pre,t-3	3.9072***	2.5574***	2.8145
	(1.3255)	(0.9442)	(1.7307)
Exam date_Pre,t-2	1.3447	2.2377**	1.3780
	(1.1408)	(0.9336)	(1.7709)
Exam date_Pre,t-1	4.1394***	1.8507***	3.2167*
	(1.2443)	(0.6095)	(1.6828)
Exam date	3.8657**	2.3344***	3.2785*
	(1.7227)	(0.5740)	(1.9524)
Exam date_Post,t+1	2.2636*	0.6406	2.5531
	(1.1850)	(0.8189)	(2.0188)
Exam date_Post,t+2	1.5171	1.4000	1.7150
	(1.4535)	(1.0509)	(2.0807)
Exam date_Post,t+3	-0.8662	1.6883***	0.8173
	(1.1822)	(0.5486)	(2.0192)
Exam date_Post,t+4	0.8125	0.7968*	3.6432**
	(0.8352)	(0.4306)	(1.4377)
Month FE	x	x	x
Year FE	x	x	x
Province FE	x	x	x
No. of observations	3,848	3,848	3,848
No. of provinces	52	52	52

Notes: Robust standard errors in parentheses clustered at the province level. Event study designed with binned endpoints. Reference category is pre-7. 74 time periods included in the estimation sample (05/2009-06/2015). All regressions are weighted with the number of employees at the province level. ***p<0.01, **p<0.05, *p<0.10.

Table 4A.4: Effect of Public Examinations on Sick Leave Behavior: Mean Duration of Sickness Spells by Economic Sector

Economic sector	**Public Adm.**	**Education**	**Health**
	Mean duration (in days)	Mean duration (in days)	Mean duration (in days)
Exam date_Pre,t-6	0.4067	1.5677	0.2526
	(1.3521)	(1.5571)	(1.1629)
Exam date_Pre,t-5	0.7594	-0.8606	-2.0898
	(0.9226)	(1.6465)	(1.2508)
Exam date_Pre,t-4	1.2051	2.8310*	-1.8786
	(1.3386)	(1.4947)	(1.6423)
Exam date_Pre,t-3	1.4943	3.9841*	-1.9372
	(1.1995)	(2.0615)	(1.4168)
Exam date_Pre,t-2	1.9366	-1.0002	-1.7583*
	(1.5047)	(1.9343)	(1.0107)
Exam date_Pre,t-1	0.1013	2.6275*	-4.4368***
	(1.5170)	(1.4088)	(1.3893)
Exam date	-0.4335	-0.7396	-3.9169***
	(1.4136)	(1.7859)	(1.3502)
Exam date_Post,t+1	1.4708	2.3953	-3.8881**
	(0.9530)	(2.6702)	(1.5195)
Exam date_Post,t+2	-0.5289	-4.6777	-0.5120
	(0.7960)	(3.7854)	(1.6092)
Exam date_Post,t+3	0.6502	-1.1965	-2.2705
	(0.9288)	(2.4280)	(1.7872)
Exam date_Post,t+4	-0.4514	0.9860	-1.6247
	(0.6255)	(0.7963)	(1.5658)
Month FE	x	x	x
Year FE	x	x	x
Province FE	x	x	x
No. of observations	3,848	3,848	3,848
No. of provinces	52	52	52

Notes: Robust standard errors in parentheses clustered at the province level. Event study designed with binned endpoints. Reference category is pre-7. 74 time periods included in the estimation sample (05/2009-06/2015). All regressions are weighted with the number of employees at the province level. ***p<0.01, **p<0.05, *p<0.10.

Table 4A.5: Effect of Public Examinations on Sick Leave Behavior: Gender Analysis by Economic Sector

Economic sector	**Public Adm.**		**Education**		**Health**	
Outcome:	# Spells	# Spells	# Spells	# Spells	# Spells	# Spells
(per 1,000 workers)	FEMALE	MALE	FEMALE	MALE	FEMALE	MALE
Exam date_Pre,t-6	3.4073***	1.7183	-0.4987	0.2581	-0.2832	2.8027*
	(1.1169)	(1.0474)	(0.5680)	(0.4671)	(1.3947)	(1.4710)
Exam date_Pre,t-5	3.3106	2.6320*	2.5201*	2.8280**	0.3292	2.2032
	(2.4353)	(1.3220)	(1.3184)	(1.3180)	(1.8201)	(1.7317)
Exam date_Pre,t-4	6.8828***	4.8835***	0.0448	0.8729	2.0245	2.8945*
	(2.3868)	(1.6244)	(1.1153)	(0.7073)	(2.0167)	(1.6804)
Exam date_Pre,t-3	3.9167*	3.7588***	2.5639**	1.5100**	2.0045	5.3249***
	(2.1326)	(1.2219)	(1.1859)	(0.6415)	(1.9619)	(1.7074)
Exam date_Pre,t-2	1.9296	1.2310	2.1375*	1.0502	0.5642	4.2467**
	(1.9676)	(1.0686)	(1.2047)	(0.7093)	(1.9570)	(1.6348)
Exam date_Pre,t-1	6.4178***	2.9424***	2.3306**	0.1989	2.3655	6.2317***
	(2.1112)	(0.9669)	(1.0336)	(0.8977)	(1.8584)	(1.9891)
Exam date	4.7923	3.5103***	2.8419***	0.5785	2.4627	6.8574***
	(3.1147)	(1.1509)	(0.8554)	(0.8640)	(2.1666)	(2.2063)
Exam date_Post,t+1	1.8102	2.2984**	0.6994	-0.4292	2.3338	2.1441
	(2.1800)	(0.8786)	(1.0475)	(1.0352)	(2.3619)	(2.3408)
Exam date_Post,t+2	-1.2748	3.2072***	1.5209	-0.3443	1.5651	-0.1852
	(2.4186)	(1.1754)	(1.3304)	(1.1305)	(2.4666)	(2.0043)
Exam date_Post,t+3	-2.5819	0.2602	1.9314*	0.2451	0.8006	-1.1825
	(2.0610)	(0.9063)	(1.0164)	(0.6795)	(2.4355)	(2.1529)
Exam date_Post,t+4	1.3415	0.3061	0.7645	-0.1247	3.9827**	1.3522
	(1.4085)	(0.6201)	(0.6025)	(0.5040)	(1.5276)	(1.5360)
Month FE	x	x	x	x	x	x
Year FE	x	x	x	x	x	x
Province FE	x	x	x	x	x	x
No. of observations	3,848	3,848	3,843	3,843	3,848	3,848
No. of provinces	52	52	52	52	52	52

Notes: Robust standard errors in parentheses clustered at the province level. Event study designed with binned endpoints. Reference category is pre-7. Outcome variable is the total number of spells per 1,000 female/male workers per economic sector. 74 time periods included in the estimation sample (05/2009-06/2015). All regressions are weighted with the number of employees at the province level. ***p<0.01, **p<0.05, *p<0.10.

Table 4A.6: Effect of Public Examinations on Sick Leave Behavior: Gender Analysis by Economic Sector

Economic sector	**Public Adm.**		**Education**		**Health**	
Outcome:	Mean duration (in days)		Mean duration (in days)		Mean duration (in days)	
	FEMALE	MALE	FEMALE	MALE	FEMALE	MALE
Exam date_Pre,t-6	-1.0611	1.4487	1.5330	2.8765	1.4992	-3.1200**
	(1.3538)	(1.3640)	(2.4197)	(3.6487)	(1.2596)	(1.4606)
Exam date_Pre,t-5	1.6103	-0.0251	0.6239	0.1843	-0.5815	-5.2874**
	(1.1367)	(1.3429)	(1.6316)	(2.1633)	(1.4151)	(2.0676)
Exam date_Pre,t-4	-0.7815	1.2535	3.4533	2.0996	-1.5543	-0.9662
	(1.1118)	(1.4944)	(2.1148)	(2.7710)	(1.6125)	(2.7823)
Exam date_Pre,t-3	0.7221	1.0227	4.4456**	6.8673*	-1.1055	-3.2053
	(1.1126)	(1.1557)	(1.9107)	(3.6249)	(1.4493)	(2.4988)
Exam date_Pre,t-2	0.4514	2.5301	1.1555	0.0700	-1.0860	-3.3646*
	(1.3856)	(1.9855)	(2.5161)	(3.5589)	(1.0167)	(1.7296)
Exam date_Pre,t-1	-1.3106	0.1708	3.5561**	4.3074*	-3.7196**	-5.2556***
	(1.7375)	(1.5160)	(1.6580)	(2.2826)	(1.4681)	(1.5366)
Exam date	0.2098	-2.0279	0.8681	3.0856	-2.9745**	-7.9023***
	(1.5788)	(1.6433)	(2.1880)	(3.2406)	(1.4322)	(2.0385)
Exam date_Post,t+1	1.5911	0.0521	2.4379	5.7737	-3.2091**	-4.8185***
	(1.3030)	(1.4128)	(3.6399)	(5.5728)	(1.5261)	(1.6064)
Exam date_Post,t+2	-0.5748	-0.3415	-2.4705	3.2835	0.4144	-2.5898
	(1.2901)	(1.5297)	(4.2220)	(6.1390)	(1.5416)	(1.8417)
Exam date_Post,t+3	0.8938	0.5735	0.1575	1.0795	-1.3378	-3.3190
	(0.9899)	(1.2302)	(2.5678)	(3.0697)	(1.7057)	(2.2681)
Exam date_Post,t+4	-0.8741*	-0.7090	1.1710	2.5176	-1.3116	-0.6815
	(0.4976)	(0.6826)	(0.8572)	(1.5100)	(1.5381)	(1.6152)
Month FE	x	x	x	x	x	x
Year FE	x	x	x	x	x	x
Province FE	x	x	x	x	x	x
No. of observations	3,848	3,848	3,841	3,787	3,848	3,848
No. of provinces	52	52	52	52	52	52

Notes: Robust standard errors in parentheses clustered at the province level. Event study designed with binned endpoints. Reference category is pre-7. Outcome variable is the mean duration of spells for female/male workers per economic sector. 74 time periods included in the estimation sample (05/2009-06/2015). All regressions are weighted with the number of employees at the province level. ***p<0.01, **p<0.05, *p<0.10.

Table 4A.7: Effect of Public Examinations on Sick Leave Behavior - Stress-related Conditions

Economic sector	**Public Adm.**		**Education**		**Health**	
Outcome	# Spells per 1,000 workers	Mean Duration (in days)	# Spells per 1,000 workers	Mean Duration (in days)	# Spells per 1,000 workers	Mean Duration (in days)
Exam date_Pre,t-6	0.8493**	-1.7419	-0.3759*	-0.7738	0.1103	1.7366
	(0.3571)	(1.7715)	(0.1880)	(4.1028)	(0.4546)	(1.9836)
Exam date_Pre,t-5	0.9644***	0.0256	0.2104	-3.7560	0.4391	-1.3040
	(0.3344)	(1.2800)	(0.2017)	(2.6584)	(0.5842)	(1.4094)
Exam date_Pre,t-4	1.2140**	1.3295	0.2365	-3.2849	0.6757	0.8490
	(0.4925)	(1.8132)	(0.2235)	(3.2181)	(0.5208)	(2.3812)
Exam date_Pre,t-3	1.4258***	2.2362	0.6100**	0.8803	0.5898	1.0208
	(0.4443)	(1.9637)	(0.2833)	(2.6066)	(0.6401)	(2.7145)
Exam date_Pre,t-2	0.5218	1.8387	-0.0306	-0.0932	0.4283	1.1455
	(0.3643)	(2.2041)	(0.1819)	(3.3946)	(0.7266)	(1.1708)
Exam date_Pre,t-1	1.4962***	1.7881	0.2357	1.0086	0.7091	-2.8074
	(0.4259)	(1.6895)	(0.3043)	(2.6664)	(0.7309)	(2.0085)
Exam date	1.2646***	-0.6685	0.5591**	-3.6197	0.5051	-1.8576
	(0.4637)	(2.0661)	(0.2161)	(2.3597)	(0.6784)	(2.3182)
Exam date_Post,t+1	1.1439***	1.5759	-0.2750	-0.3086	0.6054	-6.4947***
	(0.3170)	(2.0573)	(0.2570)	(3.4270)	(0.6871)	(1.9939)
Exam date_Post,t+2	0.5808	0.1770	-0.1986	-9.6292**	0.8683	-1.2677
	(0.4708)	(1.5154)	(0.3018)	(4.2715)	(0.6928)	(1.7359)
Exam date_Post,t+3	-0.0856	0.3771	-0.1620	-1.3900	0.1508	-1.7020
	(0.4104)	(1.5104)	(0.2266)	(4.1591)	(0.7208)	(2.3728)
Exam date_Post,t+4	0.1640	0.3974	-0.1019	1.0177	0.4179	0.2703
	(0.2600)	(0.8681)	(0.1569)	(0.9839)	(0.5021)	(1.6437)
Month FE	x	x	x	x	x	x
Year FE	x	x	x	x	x	x
Province FE	x	x	x	x	x	
No. of observations	3,848	3,848	3,848	3,822	3,848	3,848
No. of provinces	52	52	52	52	52	52

Notes: Robust standard errors in parentheses clustered at the province level. Outcome variables are (i) the total number of sickness spells per 1,000 workers per economic sector, and (ii) the mean duration of sickness spells per economic sector, both due to osteoarticular, digestive and mental conditions. Event study designed with binned endpoints. Reference category is pre-7. 74 time periods included in the estimation sample (05/2009-06/2015). All regressions are weighted with the number of employees at the province level. ***p<0.01, **p<0.05, *p<0.10.

Table 4A.8: Robustness Check: Continuous Treatment Definition

Economic sector	**Public Adm.**	**Education**	**Health**
	# Spells (1,000 workers)	# Spells (1,000 workers)	# Spells (1,000 workers)
Exam date_Pre,t-6	0.0679**	-0.0027	0.0032
	(0.0334)	(0.0029)	(0.0154)
Exam date_Pre,t-5	0.0541	0.0070	0.0000
	(0.0468)	(0.0044)	(0.0187)
Exam date_Pre,t-4	0.2063**	-0.0074	0.0038
	(0.0796)	(0.0065)	(0.0154)
Exam date_Pre,t-3	0.0864**	0.0060	0.0089
	(0.0364)	(0.0037)	(0.0163)
Exam date_Pre,t-2	0.0325	0.0037	0.0035
	(0.0313)	(0.0033)	(0.0152)
Exam date_Pre,t-1	0.1397***	0.0012	0.0174
	(0.0348)	(0.0026)	(0.0162)
Exam date	0.1230***	0.0029	0.0124
	(0.0400)	(0.0031)	(0.0174)
Exam date_Post,t+1	0.0704*	-0.0047	-0.0013
	(0.0387)	(0.0046)	(0.0177)
Exam date_Post,t+2	0.0675	-0.0022	-0.0076
	(0.0419)	(0.0055)	(0.0209)
Exam date_Post,t+3	-0.0232	-0.0005	-0.0151
	(0.0296)	(0.0033)	(0.0191)
Exam date_Post,t+4	0.0321	-0.0002	0.0073
	(0.0211)	(0.0019)	(0.0142)
Month FE	x	x	x
Year FE	x	x	x
Province FE	x	x	x
No. of observations	3,848	3,848	3,848
No. of provinces	52	52	52

Notes: Robust standard errors in parentheses clustered at the province level. Event study designed with binned endpoints. Reference category is pre-7. 74 time periods included in the estimation sample (05/2009-06/2015). All regressions are weighted with the number of employees at the province at level. ***p<0.01, **p<0.05, *p<0.10, respectively.

Table 4A.9: Robustness Check: Different Treatment Definition

Economic sector	**Public Adm.**		**Education**		**Health**	
Outcome	# Spells per 1,000 workers	Mean Duration (in days)	# Spells per 1,000 workers	Mean Duration (in days)	# Spells per 1,000 workers	Mean Duration (in days)
Exam date_Pre,t-6	-0.0996	0.4114	0.1889	1.2162	0.9774	0.8396
	(0.4499)	(0.6497)	(0.3030)	(1.4076)	(0.6035)	(0.8002)
Exam date_Pre,t-5	-0.4521	0.3018	-0.5180	2.5590**	0.3669	0.0374
	(0.4411)	(0.7929)	(0.6766)	(1.0534)	(0.5477)	(0.6168)
Exam date_Pre,t-4	0.3671	-0.2338	-1.6057***	5.0638***	0.3484	0.7915
	(0.6067)	(0.7913)	(0.5873)	(0.9285)	(0.5160)	(0.6692)
Exam date_Pre,t-3	-0.2135	0.2418	-0.2418	2.6447***	0.0229	-0.1118
	(0.5138)	(0.5884)	(0.2825)	(0.9083)	(0.6026)	(0.5044)
Exam date_Pre,t-2	-0.2411	0.1516	0.2846	4.0145*	-0.7565	0.4200
	(0.3294)	(0.5832)	(0.3472)	(2.0496)	(0.4753)	(0.3048)
Exam date_Pre,t-1	0.5483	-0.4252	-0.0773	0.5714	0.9142**	0.0502
	(0.5069)	(0.4792)	(0.2989)	(1.2023)	(0.4462)	(0.5696)
Exam date	1.0294**	-1.0264	-0.1265	-0.1255	0.6969	-0.2990
	(0.4433)	(0.6467)	(0.3213)	(1.4809)	(0.5448)	(0.6581)
Exam date_Post,t+1	0.4815	0.2005	-1.2541**	-0.6168	-0.3594	-0.2840
	(0.5573)	(0.5933)	(0.5668)	(2.3354)	(0.3808)	(0.5882)
Exam date_Post,t+2	0.8440*	-0.8096	-1.3576**	-3.4424	-0.2113	0.4951
	(0.5026)	(0.6095)	(0.6027)	(3.0343)	(0.6301)	(0.8766)
Exam date_Post,t+3	-0.1533	-0.1930	-0.0898	1.7918	-0.6869	0.6327
	(0.3943)	(0.6742)	(0.3965)	(1.7328)	(0.6318)	(0.6067)
Exam date_Post,t+4	0.0083	-0.6534**	0.0457	-0.9631	0.0638	0.0593
	(0.2541)	(0.2849)	(0.3415)	(0.6167)	(0.3960)	(0.2874)
Month FE	x	x	x	x	x	x
Year FE	x	x	x	x	x	x
Province FE	x	x	x	x	x	x
No. of observations	3,848	3,848	3,848	3,848	3,848	3,848
No. of provinces	52	52	52	52	52	52

Notes: Robust standard errors in parentheses clustered at the province level. Event study designed with binned endpoints. Reference category is pre-7. Outcome variables are (i) the total number of sickness spells per 1,000 workers per economic sector, and (ii) the mean duration of sickness spells per economic sector. Treatment definition: calls with a number of positions greater or equal to the 95 percentile of the positions distribution. 74 time periods included in the estimation sample (05/2009-06/2015). ***p<0.01, **p<0.05, *p<0.10.

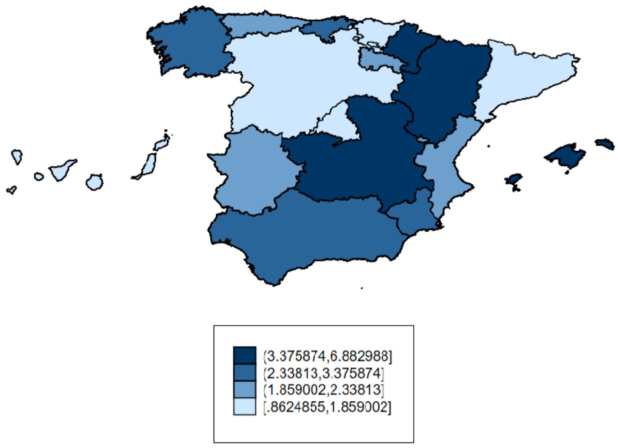
Table 4A.10: Robustness Check: Placebo Illness Test Neoplasms

Economic sector	**Public Adm.**		**Education**		**Health**	
Outcome	# Spells per 1,000 workers	Mean Duration (in days)	# Spells per 1,000 workers	Mean Duration (in days)	# Spells per 1,000 workers	Mean Duration (in days)
Exam date_Pre,t-6	-0.0524	-17.9211	-0.0335	59.4926	0.0440	11.8702
	(0.0560)	(19.6625)	(0.0219)	(39.2362)	(0.0465)	(14.2297)
Exam date_Pre,t-5	0.0351	-13.5663	0.0666**	9.6000	-0.0678	6.6629
	(0.0502)	(11.8447)	(0.0257)	(20.1463)	(0.0622)	(18.4556)
Exam date_Pre,t-4	0.0734*	-22.0389*	-0.0177	38.0890	-0.0050	-31.7400***
	(0.0406)	(11.1691)	(0.0520)	(30.2432)	(0.0475)	(11.3359)
Exam date_Pre,t-3	0.0323	-30.8182***	0.0239	-17.7145	0.1096*	-21.4700*
	(0.0539)	(9.6261)	(0.0641)	(17.1382)	(0.0616)	(12.5174)
Exam date_Pre,t-2	0.0828	-19.4661*	0.0524	-23.8100	-0.0153	-2.9492
	(0.0605)	(10.2555)	(0.0587)	(18.2923)	(0.0387)	(13.6247)
Exam date_Pre,t-1	0.1024*	-6.4822	0.0148	25.3171	0.0336	-2.6112
	(0.0606)	(12.3618)	(0.0419)	(26.3470)	(0.0461)	(15.0153)
Exam date	0.1040***	-32.4471**	0.0050	26.6083	-0.0378	-0.7234
	(0.0342)	(13.0990)	(0.0298)	(27.4989)	(0.0429))	(13.5321)
Exam date_Post,t+1	0.0147	17.1688	-0.0291	31.7017	-0.0306	-8.5103
	(0.0368)	(14.3663)	(0.0403)	(32.5526)	(0.0356)	(16.3551)
Exam date_Post,t+2	0.0080	-25.6101**	0.0409	1.4578	0.0263	13.8614
	(0.0434)	(12.3508)	(0.0389)	(33.4211)	(0.0485)	(14.7243)
Exam date_Post,t+3	0.0113	-9.0246	0.0272	13.9977	-0.0370	3.6217
	(0.0387)	(13.5907)	(0.0480)	(22.0369)	(0.0534)	(15.2729)
Exam date_Post,t+4	0.0217	-9.0967**	-0.0113	3.4547	-0.0560**	6.7949
	(0.0242)	(3.7749)	(0.0246)	(9.2332)	(0.0263)	(4.5286)
Month FE	x	x	x	x	x	x
Year FE	x	x	x	x	x	x
Province FE	x	x	x	x	x	x
No. of observations	3,848	3,636	3,848	2,872	3,848	3,722
No. of provinces	52	52	52	52	52	52

Notes: Robust standard errors in parentheses clustered at the province level. Event study designed with binned endpoints. Reference category is pre-7. Outcome variables are (i) the total number of sickness spells per 1,000 workers per economic sector related to neoplasms, and (ii) the mean duration of these sickness spells. 74 time periods included in the estimation sample 74 (05/2009-06/2015). ***p<0.01, **p<0.05, *p<0.10.

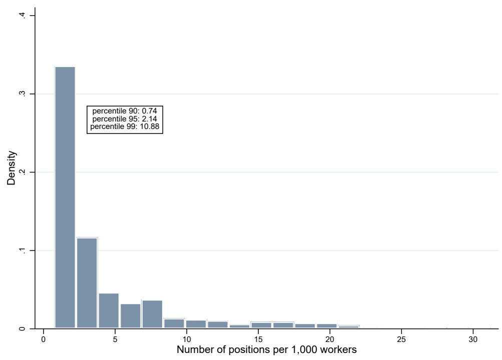
Appendix 4B: Additional Figures

Figure 4B.1: Cross-sectional Variation in the Number of Positions per 1,000 Workers at the Regional Level in year 2009.

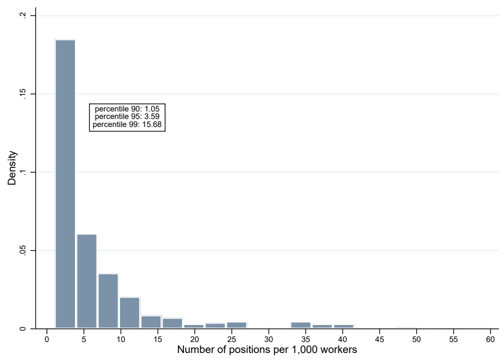


Notes: Own elaboration using administrative data from the Ministry of Territorial Policy and Public Function in Spain, for the year 2009. It also includes data from the Spanish Labor Force survey on the total number of workers for year 2009. It includes all positions that require a passing exam as entry condition in which the place of work is Spain. Excludes all temporal contracts.

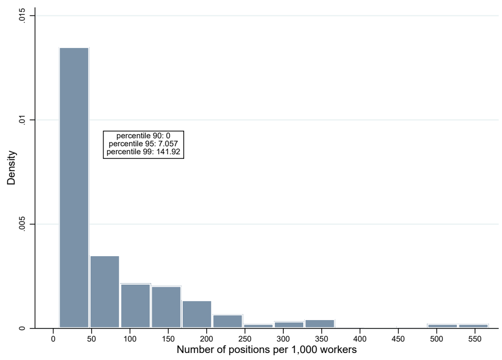
Figure 4B.2: Distribution of Large Events per Economic Sector



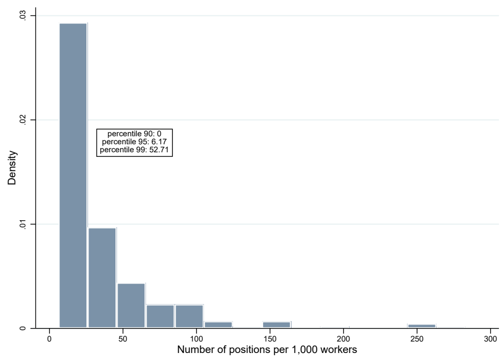
(a) All sectors per 1,000 workers



(b) Public administration per 1,000 workers



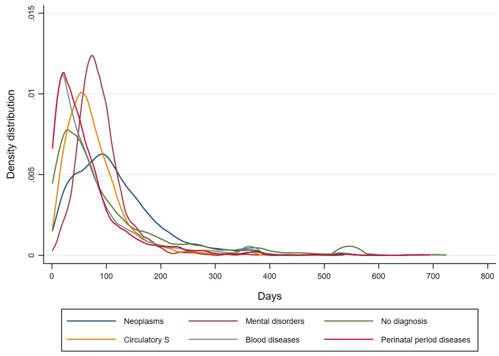
(c) Education per 1,000 workers



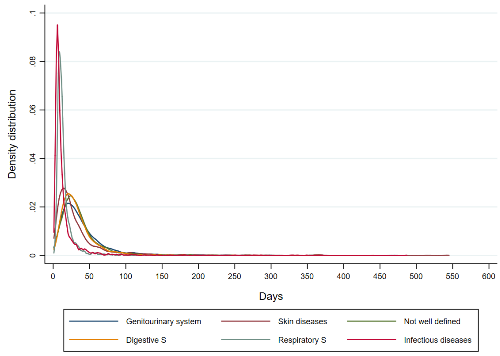
(d) Health per 1,000 workers

Notes: Figure shows the distribution of large events, which are defined as calls with a number of positions equal or greater than the 95 or 99 percentile of the positions distribution in each economic sector. uthors' calculations based on data from the Ministry of Territorial Policy and Public Function in Spain and the Spanish Labor Force Survey, for years 2009-2015.

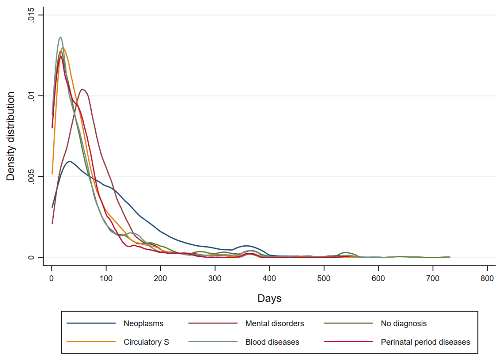
Figure 4B.3: Distribution of Average Mean Duration of Sickness Absences per Illness Category



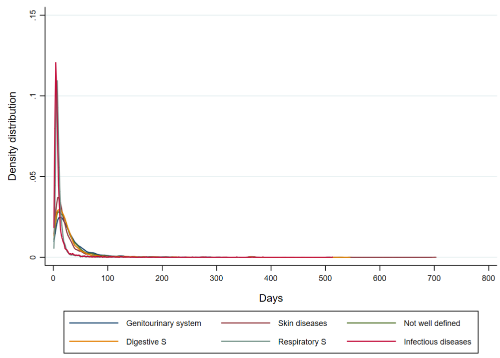
(a) Public Administration



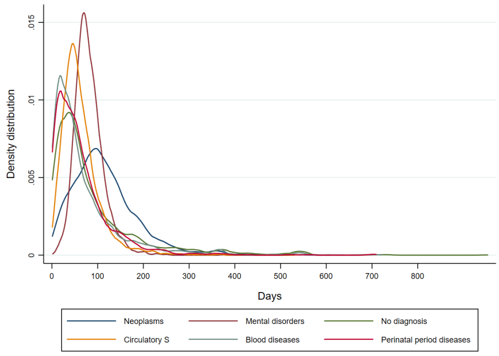
(b) Public Administration



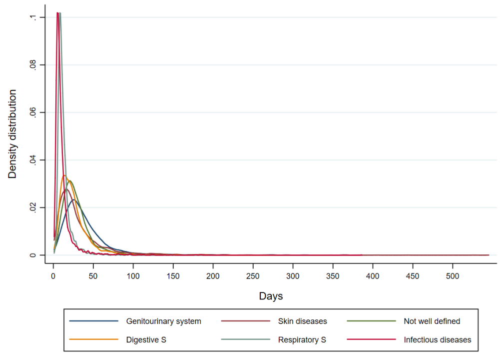
(c) Education



(d) Education



(e) Health



(f) Health

Notes: Figure shows the distribution of average mean duration of sickness absences per illness category within every economic sector. Authors' calculations based on data from the Social Security Institution.

5 Covid-19 Lockdown in Ecuador: Are there Gender Differences in Unemployment?

5.1 Introduction

The Covid-19 crisis is one of the most challenging events that the world has faced in recent decades. According to the World Health Organization (WHO), as of July 2022, Covid-19 is purported to have caused more than 547 millions infections and more than 6 million deaths worldwide. Under the pretext of reducing contagion, a large number of countries undertook non-pharmaceutical interventions (NPIs) to restrict personal mobility and social contact among individuals. These restrictions produced detrimental effects on the economy, which saw the labor market severely affected. For instance, for the Latin American and Caribbean region, according to the International Labor Organization (ILO), the unemployment rate rose from 8.1% (in 2019) to 10.6% (in 2020), which translates to 5.4 million additional job seekers and 30.1 million in total by the end of 2020.

Furthermore, this (health) economic crisis has negatively affected employment for women, less educated, young, and older people, deepening already existing inequalities in labor markets (Escalante and Maisonnave, 2022; Serrano et al., 2019; Hoehn-Velasco et al., 2021; Aldan et al., 2021). The Americas region experienced the greatest reduction in women's employment as a result of the crisis. One plausible explanation for this might be that female employees are more concentrated in economic sectors strongly affected by the crisis (Maurizio, 2021). Moreover, according to data from 17 countries in the Latin American region, the labor participation rate for women fell (on average) by 8.1% in 2020, compared with 5% among men, while the employment rate among men contracted by 7.2% and among women by 10.2% in 2020 (Arreaza et al., 2021).

In this chapter, we assess the short-term consequences that the lockdown policy may had on women's labor market participation in Ecuador. We analyze whether there is an increase in new unemployment spells in the formal labor market for women workers. The restriction policy in this country, as in many others, was targeted to a group of economic activities, which allows us to exploit the existing variation in the affected relative to the non-affected economic sectors. We apply a difference-in-differences-in-differences approach to estimate the relation between the lockdown policy and the number of new unemployment spells, comparing differences between restricted vs non-restricted economic activities (first difference), and female vs male employees (second difference) before and after the closure of economic activities. For this, we use a unique, novel, and non-explored administrative data set, provided by the Ecuadorian Social Security Institution, on the universe of the new number of unemployment registered in the social security, across all provinces of the country. We complement this analysis using survey data to explore other labor market outcomes.

As a first insight of the results, we find that the sector-specific lockdown policy is negatively associated with formal employment for women working in the restricted economic sectors. Thus, we find an increase of approximately, 15% in new unemployment spells for the group of interest. This result is robust to a set of checks such as an event study design, the exclusion of geographic areas with a higher share of employed individuals, and a placebo in the timing of the lockdown policy. Moreover, we provide some evidence on a set of other labor market outcomes, which show that women working in the restricted sectors are more likely to be unemployed (5.76 percentage points). In addition, we provide evidence that suggests that younger, older and less educated women seem to have been more likely to lose their jobs.

Research that assesses Covid-19 lockdown policies and labor market outcomes is very rich. However, most of the literature focuses on developed economies. For instance, for the United States of America (U.S.), Beland et al. (2020) find that the Covid-19 increases the unemployment rate and decreases hours of work and labor force participation. Likewise, Baek et al. (2021) find that each week of stay-at-home orders increases unemployment by 1.9%. With respect to the European context, Bauer and Weber (2021) and Juranek et al. (2021) also find negative effects in terms of unemployment.[1]

With regards to research that focuses on developing economies, evidence is limited. To the knowledge of the authors, there are very few papers that explore the impact of the Covid-19 lockdown on labor markets. For instance, Morales et al. (2022) explore the effects of Covid-19 on the Colombian labor market, and find that sector-specific restriction policies had negative effects on employment. Campos-Vazquez et al. (2021), using job ads from a leading job search website in Mexico, find that

[1]For an extensive literature review of this topic, see Brodeur et al. (2021).

there is a decline in the number of job advertisements, but that there is no structural change in labor demand. Moreover, Olivera et al. (2021) find a reduction in employment of around 16.5 percentage points for regions more affected by the lockdown in Venezuela.

As seen, the Covid-19 lockdown had large negative impacts on labor market outcomes in several countries. However, this effect might be different across demographic characteristics like gender (Farré et al., 2022; Adams-Prassl et al., 2020), age (Abraham et al., 2022; Gustafsson, 2020; Lee et al., 2021; Bui et al., 2020), education (Aldan et al., 2021; Holder et al., 2021), and others.

In Latin American countries, few studies analyze the differential effect that the lockdown policy might had on female compared to male workers. For instance, Escalante and Maisonnave (2022) show that women suffered more than men from the negative impacts of the crisis in Bolivia. Specifically, unskilled women are the most affected. Hoehn-Velasco et al. (2021) use administrative data that covers Mexico's formal labor market to investigate the employment effects of the Great Lockdown and find that the most affected economic agents during the crisis include the youngest workers (15–29- years-old), the oldest workers (over 60 years old), and low-income earners.

This chapter contributes to the existing literature on the Covid-19 restrictions and labor outcomes in several ways. First, we provide evidence on whether female employees have been affected differently by the lockdown, which is, to the knowledge of the authors, very scarcely studied for the Latin American region. We also explore differences across demographic groups like age and education level. Second, we analyze two margins: (i) unemployment spells in the formal sector using social security records, and (ii) the probability of being employed using survey data. Third, we exploit the exogeneity of the lockdown policy, which in Ecuador was implemented before the peak of the number of contagions, which allows us to draw reliable conclusions (Morales et al., 2022). Finally, for the main analysis, we use a high-quality official administrative data set that covers all formal sector workers at the monthly level, avoiding problems that may arise with the use of survey data.

Ecuador provides an interesting setting for studying NPIs that, on the one hand, are effective in terms of health outcomes but, on the other hand, might be detrimental to the economy. Ecuador is a country with important levels of inequality compared to Brazil, Chile, Colombia, Costa Rica, Guatemala, Honduras, and Panama (Acevedo et al., 2022; Delaporte et al., 2021), and the loss of formal employment may increase these already existing inequalities as individuals who lose their jobs might transition from the formal to the informal labor market. In addition, losing a formal position means losing access to social security benefits such as health care (for the employee and their dependents) and personal loans (directly provided by the social security and which are available only to employees contributing to social security), among others. Moreover, involuntary job loss might be associated with social problems such as increases in criminal behavior, poor health care, and limited access to education (Escalante and Maisonnave, 2022; Khanna et al., 2021; Britto, 2022; Rose, 2021). The results of this chapter are of important policy relevance, because they may be a tool to help policy makers to identify the most affected groups and, in this way, design recovery mechanisms focused on them.

The rest of the chapter proceeds as follows. Section 5.2 describes the Covid-19 scenario in Ecuador. In Section 5.3, we present the data and some descriptive analysis. Section 5.4 explains the empirical strategy. Section 5.5 describes the main results. In Section 5.6, we provide a set of robustness checks of the baseline results. In Section 5.7, we explore other labor market outcomes and, finally, in Section 5.8, we state the conclusions of the chapter.

5.2 Covid-19 in Ecuador

By the end of February 2020, the first case of Covid-19 was officially reported in Ecuador. In the Latin American region, the case of Ecuador is particularly relevant, as Ecuador was, according to official data, one of the most affected countries in terms of number of cases and fatalities per million inhabitants during the first months of the crisis. According to Cuéllar et al. (2021), Ecuador had an excess death rate of around 64%.[2] Figure 5B.1 in the Appendix 5B shows (raw) excess deaths per month during the year 2020 in Ecuador. From here, we see that the peak was around April and May, two months after the onset of mandated home confinement.

According to ILO (2020), this crisis affected Latin American labor markets in a non-negligible way. The average unemployment rate for the region reached 10% in 2020 (Comisión Económica para América Latina y el Caribe, 2020). In Ecuador, it was around 13% during the second quarter of 2020 (Instituto Nacional de Estadísticas y Censos - INEC, 2020). One of the main problems behind job loss in Ecuador is the large number of micro, small, and medium-sized firms, which represent around 95% of all formal companies and generates around 50% of formal employment (Superintendencia de Compañías Valores y Seguros, 2018). Micro and small companies are more vulnerable to bankruptcy than large firms (Camino-Mogro et al., 2020; Beverinotti and Deza, 2020; Carrillo-Maldonado et al., 2020), which makes the Ecuadorian labor market even more sensitive to negative shocks that affect formal employment.

Like many other countries in the region (and the world), Ecuador implemented a strict lockdown policy, which was legally established by means of the Presidential Decree Number 1017 of 2020. In this document, the authorities placed restrictions on personal mobility and closed establishments, beginning on March 16, 2020.

The measures taken by the government included a quarantine of the entire population and restricted mobility that was allowed only in justified cases such as medical emergencies or going to workplaces considered essential. The group of essential economic activities included public and private health services, national security, the financial and insurance sectors, food provision, agriculture and mining activities, electricity and water supply, information and communication, and the transport sector, which was allowed to operate only in relation to health and public security services. On the other hand, the activities affected by the lockdown included wholesale and retail trade (in the section not corresponding to food supply), construction, other types of transportation, accommodation and food services, some types of information and communication, professional services, administrative and support

[2]For an extensive review of statistics on COVID-19, see Roser and Hasell (2020).

services, and both public and private education (all levels), among others. We will be more specific about affected and non-affected economic sectors in the following section.

5.3 Data and Descriptive Statistics

5.3.1 Main Sources

Regarding flows from employment to unemployment, we use unique administrative data provided by the Social Security Administration of Ecuador (IESS, in Spanish) from January 2019 to December 2020. This data comprises the universe of new unemployment spells for individuals working in the formal sector, that is, those that contribute to the social security system. The unit of observation is the unemployment spell in an economic sector in a given province, for which there is information on the age, gender, and province of residence of the individual that lost the job. We also observe the type of contribution scheme of the employee, which can be private, public, autonomous or domestic workers. In addition, there is information on the economic activity sector of the firm or institution at the four-digit level according to the International Standard Industrial Classification (ISIC). Finally, we have the date on which the formal job loss occurred. With this, we are able to obtain the outcome of interest, which is the monthly (log) total number of new ended contracts.

From the universe of new unemployment spells, we restrict the sample to unemployment episodes that occurred only in private firms (private sector of the economy) as this was the group that was (officially) targeted with the lockdown policy. In addition, we focus on individuals of legal working age (from 15 to 65 years old) and contributing to the general social security system.[3] This definition leaves us with a total of 2,105,368 new ended contracts in the period of analysis.

With respect to the Covid-19 data, we use official administrative data released by the Ecuadorian institution in charge of registering births and deaths “Registro Civil”. From here, we construct a variable that contains excess deaths per province and month between January 2019 and December 2020. We use excess deaths, instead of Covid-19 cases as there might be concerns regarding the accuracy of confirmed cases, especially during the so-called first wave of the pandemic in the country. In Figure 5B.1 of the Appendix 5B, we show monthly (absolute) excess deaths for the year 2020. We can see that the peak is between April and May 2020 (which is compatible with the peak of Covid-19 cases), and, as mentioned in the introduc-

[3]This implies that we do not consider domestic workers and individuals in the agricultural and mining sectors, which pertain to a special social security scheme.

tion, the lockdown restrictions were imposed several weeks before the peak of the crisis.

5.3.2 Treatment and Control Groups

As pointed out in the introduction, our main approach compares differences across affected versus non-affected economic activities (first difference) and women versus men (second difference), before and after the implementation of the lockdown (third difference).

Regarding the first difference, we rely on what was established in the Decree 1017 (signed by the president of the country) with respect to the restricted economic activities. From here, using the ISIC classification at the four-digit level, we derive both groups. The treated group consists of activities that were ordered (by law) to cease their functioning, such as construction, accommodation and food service activities, education, wholesale and retail trade, and manufacturing, among others, which totals 1,394,708 unemployment spells. In the control group, the economic activities that were not forced to close, we have groups mainly in wholesale and retail sectors (in the section that corresponds to food provision), human health and social work activities, financial and insurance activities, and manufacturing (in activities that corresponds to vital goods), which adds up to 710,660 new ended contracts. In total, we have 21 groups of economic activities that we categorize into affected or non-affected sectors depending on whether the economic activity at the 4-digit ISIC level was restricted by law.[4]

An extensive list with the treated economic activities is shown in Table 5A.1 in the Appendix 5A. In addition, in Table 5.1, we present the total number of unemployment spells across gender and affected and non-affected economic sectors. For the second difference, we construct the female indicator variable using the reported sex of the employee. We generate a dummy variable which takes the value of one for women and zero for men.

[4] We use this aggregation because using the economic sectors directly at the four-digit level may generate groups with too few observations, which in the estimates may raise concerns about few observations within clusters.

Table 5.1: Unemployment Spells in Ecuador by Economic Sector and Gender.

Economic sector activity (ISIC)	All	Female	Male
Affected sectors	1,394,708	409,751	984,957
Non-affected sectors	710,660	251,151	459,509
Total	2,105,368	660,902	1,444,466

Elaboration: the Authors using administrative records from the Social Security Administration in Ecuador (IESS) from January 2019 to December 2020.

5.3.3 Descriptive Evidence

We start by exploring the monthly evolution of new unemployment spells in 2019 and 2020 across treated and control groups. Panel (a) of Figure 5.1 shows raw trends for women working in restricted sectors versus women in non-restricted sectors. We see that the first group seems to experience a steeper increase in the number of unemployment spells right after the implementation of the lockdown. When we compare unemployment spells for women working in affected sectors with their male counterparts (men working in restricted sectors), both groups seem to have experienced increases in the total number of new ended contracts as depicted in panel (b). In any case, this is only descriptive evidence that we will contrast in a later regression framework.

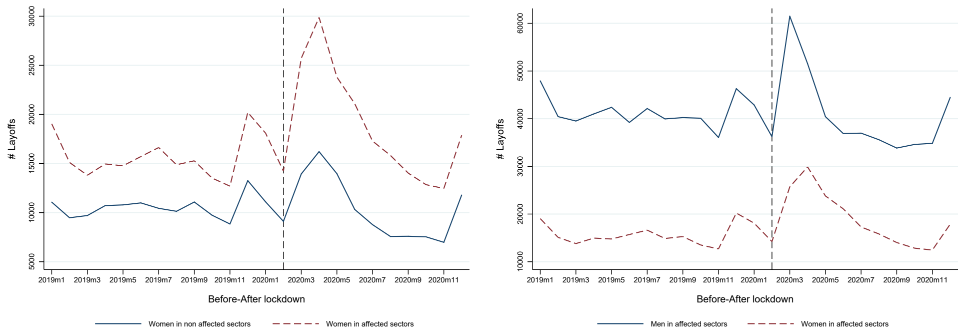
We also present in Figure 5.1 raw trends for monthly unemployment spells separately by gender, that is, female vs male. As one can observe from panel (c), both groups seem to have followed very similar trends, though men, in absolute numbers, show more unemployment spells. Once more, this is a first bit of descriptive evidence that we need to explore further. Moreover, panel (d) depicts the evolution of unemployment spells comparing restricted sectors vs non-restricted sectors, where the affected group shows a steeper positive slope. Finally, in Figure 5B.2 in the Appendix 5B, we present the monthly evolution of unemployment spells for the entire economy.

5.4 Empirical Strategy

5.4.1 Motivating Difference-in-Differences Estimates

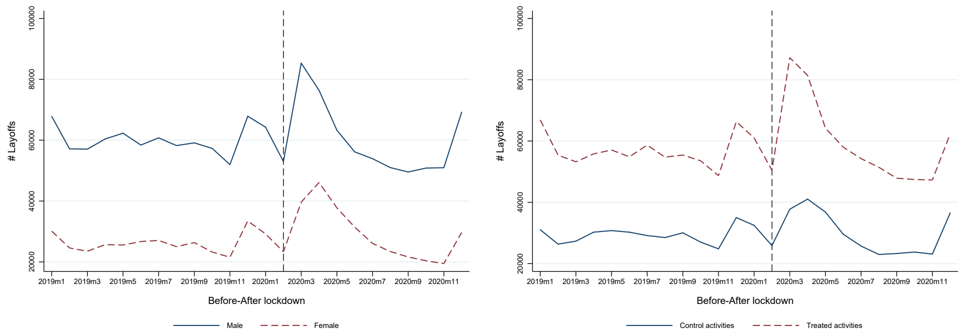
The baseline specification is a difference-in-difference-in-differences (DDD) which compares differences across economic activities (first difference) and gender (sec-

Figure 5.1: Evolution of Unemployment Spells per Group



(a) Female affected sectors vs female non-affected sectors

(b) Female affected sectors vs male affected sectors



(c) Female vs Male

(d) Affected sectors vs non-affected sectors

Notes: Figure shows the evolution of unemployment spells. Panel (a) depicts trends between women working in restricted sectors vs women in non-restricted sectors. Panel (b) shows the evolution split by women and men both working in affected economic sectors. Panel (c) compares raw trends for women vs men, and panel (d) presents affected economic sectors vs non-affected sectors. Own elaboration using administrative records from social security data.

ond difference) before versus after the lockdown in Ecuador (third difference).

Why might the COVID-19 lockdown in Ecuador be particularly harmful for the employment (measured through unemployment spells) of women working in restricted economic activities? First, as discussed in the introduction, there are a number of studies that find that this crisis negatively affected employment more for women than for men. According to Alon et al. (2020), one reason for this may be that women's employment is concentrated in less cyclical sectors such as health care and education. In addition, the COVID-19 crisis had a bigger impact on activities requiring more in-person contact such as services, in which women represent an important share of the labor supply. Regarding the situation in Ecuador, according to the National Employment Survey (ENEMDU in Spanish) for the year 2019, 51.9% of individuals employed in the wholesale and retail trade sectors, 66.49% in accommodation and food service activities, 65% in the education sector, and 62.30% in other personal service activities are women. These sectors were hit the hardest by the closure of economic activity. Second, these activities are also less likely to be performed on-line or via telecommuting, which creates differences between male and female labor outcomes. Third, in Ecuador (as in many other countries), educational activities for children and adolescents moved online, which might have increased the demand for childcare at home, a task that is more likely to be carried out by women (Alon et al., 2020).[5]

From Figure 5.1, we learned that women working in affected economic sectors seemed to experience a sharper increase in the number of unemployment spells after the lockdown compared with women working in non-affected sectors. We assess this change by performing a difference-in-differences regression, which compares female workers in restricted sectors with women in non-restricted sectors. To do this, we collapse female unemployment spells data at the province, economic sector, and time (month-year) level, and estimate an equation of the form:

$$\begin{aligned} Y_{spmt} = \alpha + \beta_1 TreatSector_s + \beta_2 Post_{mt} + \beta_3 TreatSector_s * Post_{mt} \\ + \beta_4 Control_{pmt} + \delta_s + \rho_p + \gamma_m + \tau_t + \epsilon_{spmt} \end{aligned} \tag{5.1}$$

where $TreatSector_s$ is a binary indicator that takes the value of one for restricted economic sectors and zero otherwise. $Post_{mt}$ denotes the pre-post lockdown order. Y_{spmt} is the natural logarithm of the number of new unemployment spells in the female sample. We also include controls $Controls_{pmt}$ to account for the regional variation in the spread of Covid-19, the share of excess deaths, in province *p*, month *m*, year *t*. We include a set of fixed effects: (i) month γ_m and year τ_t fixed effects to account for potential common time shocks across units, (ii) economic sector δ_s, and province fixed effects ρ_p accounting for time-invariant heterogeneity in each labor market. There are 21 economic sectors and 24 provinces. We cluster standard errors at the economic sector level since this is the level at which the effect takes place (Wooldridge, 2003; Cameron et al., 2008). As there are 21 clusters, we use wild bootstrapped clustered standard errors, applying 999 replications.[6]

In Equation (5.1), β_3 captures the post-lockdown effect of being a woman working in affected economic sectors relative to women working in non-affected sectors. Column (1) of Table 5.2 shows the results. As expected, the coefficient of interest is positive and significant, which suggests that the number of unemployment spells for women working in restricted sectors increased by approximately 41% compared with women working in other sectors after the implementation of the lockdown.

[5] In Ecuador, education moved online on March, 13th, 2020, a situation which lasted for more than a year.

[6] For computation, we use the boottest Stata command (Roodman et al., 2019).

Table 5.2: Difference-in-Differences Estimates for (ln) # of Unemployment Spells

	(1)	(2)	(3)	(4)
	Entire Female Sample	**Entire Affected Sectors Sample**	**Entire Male Sample**	**Full Sample**
DD (TreatSector*Post)	**0.4169****	**-**	**0.1750**	**0.1305**
	[0.0370]		[0.4625]	[0.6226]
DD (Female*Post)	**-**	**0.0728***	**-**	**-**
		[0.0621]		
Post-lockdown dummy	-0.2975	0.1170*	-0.0983	-0.1318
	[0.2923]	[0.0971]	[0.7247]	[0.7327]
Treat sector dummy	-0.4244	-	-1.3945	-0.9309
	[0.2633]	-	[0.1892]	[0.2402]
Female dummy	-	-0.8011**	-	-
		[0.0070]	-	-
Share excess deaths	-0.0000	-0.0000*	-0.0000	-0.0000
	[0.2783]	[0.0791]	[0.4805]	[0. 4745]
Province FE	x	x	x	x
Economic activity FE	x	x	x	x
Month FE	x	x	x	x
Year FE	x	x	x	x
Observations	12,013	12,259	12,013	12,013
Number of clusters	21	12	21	21

Notes: Column (1) shows estimates for DiD that compare women working in restricted sectors (treated group) relative to women working in non-restricted sectors. Column (2) reports DiD estimates that compare women working in restricted sectors (treated group) with men working in restricted sectors before-after the mandated quarantine. Column (3) reports DiD estimates that compare men working in restricted sectors with men working in non-restricted sectors. Column (4) reports DiD estimates that compare all workers working in restricted sectors with workers in non-restricted sectors. P-values refer to standard errors clustered at the economic activity level in square brackets, calculated using the wild cluster bootstrap with 999 replications. **p<0.05, *p<0.10.

Furthermore, it is also relevant to analyze differences in unemployment spells between women and men working in restricted economic sectors. For this, we focus on new ended contracts that took place in the affected economic sectors and collapse the data set at the province, gender and time (month-year) levels. Formally, we estimate a DD model that compares women with men:

$$\begin{aligned} Y_{sgpmt} = \alpha + \beta_1 Female_g + \beta_2 Post_{mt} + \beta_3 Female_g * Post_{mt} \\ + \beta_4 Control_{pmt} + \delta_s + \rho_p + \gamma_m + \tau_t + \epsilon_{sgpmt} \end{aligned} \tag{5.2}$$

In this specification, the outcome of interest, Y_{sgpmt}, the variable $Post_{mt}$, controls $Control_{pmt}$, and the fixed effects are the same as in Equation (5.1). $Female_g$ is a binary variable that takes the value of one for female workers and zero for male workers. Standard errors are clustered at the economic sector level (12 clusters) and computed using wild bootstrapped replications. In this equation, β_3 captures the post-lockdown effect of being a woman relative to men (when both are) working in the affected economic sectors on the (ln) number of new unemployment spells. From Column (2) of Table 5.2, we see that the number of new unemployment spells after the lockdown was mandated increased by about 7.28%.

We also perform the same specification as in Equation (5.1) for two additional sub-samples: unemployment spells (i) for males and (ii) in the whole economy. Columns (3) and (4) of Table 5.2 show the estimated coefficients. When we compare unemployment spells for male workers working in the affected sectors with those for male workers in non-affected sectors, we do not find a significant result. Moreover, when comparing unemployment spells in the restricted economic sectors relative to the non-affected sectors for the whole economy, we do not find a significant coefficient, either.

5.4.2 Baseline Equation

The differences found in the previous subsection motivate a DDD estimator which includes differences over time, gender, and economic sectors. Recall that we are interested in assessing whether the lockdown policy affected women differently than men in terms of the number of inflows from employment to unemployment in the formal sector (private sector). To formally assess this, we collapse the unemployment spells data at the province, economic sector, gender, and time (month-year) levels, and estimate a DDD equation of the form:

$$\begin{aligned} Y_{spgmt} = \alpha + \beta_1 Female_g + \beta_2 Post_{mt} + \beta_3 Treat_s * Post_{mt} \\ + \beta_4 Female_g * Post_{mt} + \beta_5 Female_g * Treat_s \\ + \beta_6 Female_g * Treat_s * Post_{mt} + \beta_7 Control_{pmt} \\ + \delta_s + \rho_p + \gamma_m + \tau_t + \epsilon_{spgmt} \end{aligned} \tag{5.3}$$

where Y_{spgmt} is the natural logarithm of the number of new unemployment spells in economic sector *s*, province *p*, gender group *g*, month *m*, year *t*. The $Post_{mt}$, $Treat_s$, $Female_g$, and $Controls_{pmt}$ variables are defined in the same way as in Equations (5.1) and (5.2). The coefficient of interest is β_6, which captures the effect of the lockdown on new ended contracts for women working in restricted economic sectors. We also include month γ_m, year τ_t, economic sector δ_s, and province fixed effects ρ_p.[7] There are 21 economic sectors and 24 provinces. We compute wild bootstrapped clustered (at the economic sector level) standard errors, applying 999 replications.

With respect to the outcome variable, which is the natural logarithm of the number of new unemployment spells, it would be ideal to use instead the ratio of work separations over the number of employees per province, time and economic sector. However, because of limitations on data availability, the working population disaggregated at the province and economic sector levels is not available for use by the authors.[8] Using a ratio instead of the number of work separations, may help to have a relative measure that takes into account the size of the workforce. Despite this drawback of the data available, one should keep in mind that our estimates are still a first step in quantifying gender differences in unemployment spells. Using a relative measure such as the proportion of unemployment spells over the number of employed individuals will provide a more accurate picture of the relation between the lockdown and unemployment spells.

The difference-in-differences estimator intends to provide an unbiased estimate of the treatment effect in a situation in which, in the absence of the treatment, the outcome in the two groups would have followed the same trend. This is more critical for the pre-treatment period; that is, one needs to add evidence that supports the parallel trend assumption before the implementation of the policy. On this point, Figure 5.1 provides preliminary descriptive evidence that adds validity to this assumption. In panel (c), we show the evolution of unemployment spells (in the private sector) for female and male workers. Panel (d) depicts the evolution of the outcome variable for restricted and non-restricted economic activities. From here, we can see that, in general, pairwise groups follow similar trends before the lockdown policy. We also see a sharp increase in the number of new ended contracts right after the restriction for the affected economic activities. Though the same increase is seen in the non-affected sectors, the increase seems to be sharper for restricted sectors.[9]

The identification strategy of this chapter is supported by several facts. The first is the exogenous characteristic of Covid-19. Second, as mentioned before, the lockdown was established before the peak of contagion in Ecuador. Third, we analyse

[7]Table 5A.2 in the Appendix 5A presents the estimates after including gender-specific linear time trends and economic sector-specific linear time trends, separately. The results are very similar.

[8]The official national statistics institution of the country, "Instituto Nacional de Estadísticas y Censos" (INEC), does not publish disaggregated data on the workforce. This data is available only at the national level. In addition, to this date, we do not have access to the number of employees that contribute to social security for the periods of our analysis.

[9]Though descriptive graphs are a valid tool, we have to keep in mind that they do not consider covariates or fixed effects as does a regression framework. In the robustness check section, we provide additional evidence that supports parallel trends.

private sector workers as there might be important differences between public and private employees. Fourth, the DDD specification allows us to control for two kinds of potentially confounding trends: changes that affect the number of unemployment spells of women unrelated to the lockdown, and changes that affect unemployment spells of all women and men working in the affected economic sectors. Finally, we provide, in the following sections, several robustness checks to support our results. For instance, we perform an event study to check for potential pre-trends. We also use placebo lockdown dates and restrict the sample by excluding the most populated provinces in the country.

5.5 Results on Unemployment Spells

Table 5.3 shows the estimates of Equation (5.3). We report results for the outcome of interest, first without any disease-spread control variable, and then including excess deaths. Both regressions include province fixed effects, month and year fixed effects, and economic activity fixed effects. Standard errors are clustered at the economic activity level.

From the results of Column (1), we see that the restriction policy is associated with an increase of approximately 15.06% (estimated coefficient 0.1506) in the number of new unemployment spells for women working in the affected economic sectors relative to other workers. This difference is statistically significant at the 5% level. In Column (2), when we include the share of excess deaths per province and month, the estimated coefficient is very similar in magnitude and significance and points to a detrimental result for female employees. In this column, we also observe that the disease-related control variable is negative though very close to zero and slightly significant.[10]

The results above suggest that female workers might have been differently hit by the lockdown. In fact, several studies indicate that the demographic composition of (un)employment varies between low and high education level, race, age group, and gender. For instance, Hoynes et al. (2012) find that general recessions are worse for men. However, literature on Covid-19 and labor market outcomes for the United States of America finds that women seem to have been more affected in terms of unemployment and hours worked (Forsythe et al., 2020; Beland et al., 2020; Cajner et al., 2020; Mongey et al., 2021). In this line, there are several studies conducted (under the Covid-19 framework) for the U.S. that show larger impacts on women

[10] As the pandemic is also very dynamic, as a robustness, we replace year and month fixed effects with month-year fixed effects. The results from this exercise are very similar and available upon request.

Table 5.3: Baseline Estimates Unemployment Spells

	(1)	**(2)**
	Log. unemployment spells	**Log. unemployment spells**
DDD Female*Treat*Post	**0.1506****	**0.1459****
	[0.0340]	[0.0400]
Interaction Female*Post	-0.0762	-0.0775
	[0.2703]	[0.2673]
Interaction Female*Treat	-0.3445	-0.3445
	[0.3724]	[0.3493]
Interaction Treat*Post	0.0052	0.0031
	[0.9600]	[0.9750]
Share excess deaths		-0.0000*
		[0.0901]
Province FE	x	x
Economic activity FE	x	x
Month FE	x	x
Year FE	x	x
Observations	20,391	20,391
Number of clusters	21	21

Notes: OLS estimates of Equation (5.3). P-values referring to standard errors clustered at the economic activity level in square brackets, calculated using wild cluster bootstrap with 999 replications. **p<0.05, *p<0.10.

(Montenovo et al., 2022; Forsythe et al., 2020; Beland et al., 2020; Mongey et al., 2021), but there is also some evidence that suggests that during the first stages of the lockdown, male employees were more affected (Béland et al., 2020). With respect to research on Europe, Fana et al. (2020) find mixed results in terms of gender differences across countries like Spain, Poland, Germany and the U.K. On the contrary, Farré et al. (2022), using household survey data from Spain, find that women were slightly more likely to lose their job than men. Regarding results for the Latin American scenario, to the knowledge of the authors, there are no documents that explore gender differences in labor market outcomes (under the Covid-19 context). The closest work is that of Morales et al. (2022), who estimate the impact of Covid-19 restrictions for Colombia, where they find an estimate equivalent to 13.9% more jobs lost in the excluded sectors relative to the non-excluded sectors. As seen, there is mixed evidence on whether women or men were more affected, so there is a need to explore the source of potential differences as Montenovo et al. (2022) establish in their research.

As we pointed out in previous sections, women tend to be more concentrated in sectors that require more in-person contact, but the composition between men and women might also be different within restricted and non-restricted sectors. Understanding these differences may be useful to interpret the results we find. To examine whether women and men are similar within affected and non-affected sectors, we provide in Table 5A.3 of the Appendix 5A some descriptive statistics for selected characteristics comparing female and male workers within treated and untreated economic sectors, separately. For this, we use survey data from the "National Survey of Employment" (ENEMDU) of Ecuador and restrict the sample using the same considerations as in Equation (5.3).

Table 5A.3 shows the results for the share of low educated workers, defined as those who have less than 10 years of education; the monthly payment, which corresponds to average monthly employment income; and the share of blue-collar workers, which corresponds workers performing manual labor.

The results show that in the treated sectors, the share of low educated female workers is lower than the share of male workers, and the difference is statistically significant. With respect to the monthly payment, on average women earn more compared to men, being the difference statistically significant. Likewise, women represent a lower share of blue-collar workers compared to men, the difference between men and women is also statistically significant. This may be because less educated women in Ecuador tend to do unpaid/paid domestic work, activity that is not considered in this analysis. Therefore, women who work in the treated sectors are probably more educated, which is compatible with a higher salary, and less probability of having a blue-collar position. The World Bank (2019) in a report for gender gaps in Ecuador find that almost 25 percent of employed women are unpaid workers (in or outside the household) or perform domestic chores, as compared to 6.6 percent of men performing unpaid work. They also find that women dominate in sectors of domestic service, hotel and restaurants, education, and services, while the representation of women in services like construction, manufacturing, transportation, and infrastructure is rather low. They also highlight that as a reflection of increased education, there are more women than men performing "Professional" occupations.

We find something similar in the untreated sectors, in which we see that the share of low educated workers is lower for women, women earn on average a higher salary, and represent a lower share of blue-collar positions. The differences between men and women are all statistically significant, except for the average monthly payment.

In general, for the characteristics and sample analyzed, we might think that women working in the affected sectors are less likely to lose their jobs, because they are more educated, with higher average salaries and less blue-collar positions compared to men. This descriptive evidence might support the statement that the lockdown policy implemented in Ecuador is highly associated with job losses for women working in the restricted sectors.

Overall, the results of this section suggest that the lock down policy is negatively associated with an important loss in the extensive margin of formal jobs for women working in the restricted economic sectors. This finding is relevant because these individuals might transition from the formal to the informal labor market. Informality has several disadvantages for workers. For instance, it is related to lack of health insurance, contributory pensions, limited career advancement, and to higher levels of inequality (Messina and Silva, 2018; Busso et al., 2021). Finally, it would also be interesting to assess differences in the number of hours worked and wages. However, this information from administrative records is not available for external use.

5.6 Robustness Checks of Baseline Results

In this section, we include a battery of robustness checks to the baseline results of Equation (5.3). First, we perform an event-study design to explore potential pre-trends. Second, we explore whether our results hold when we exclude the most populated provinces of the country. Finally, we perform a placebo test in which we change the timing of the lockdown policy.

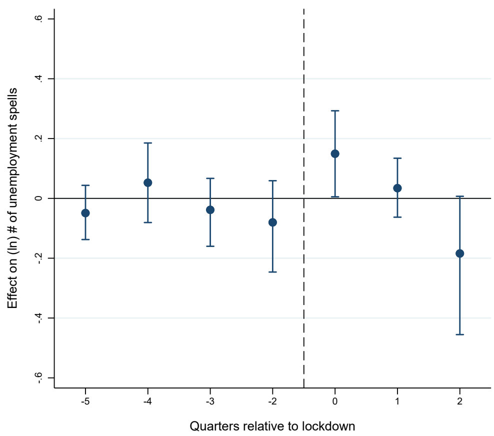
5.6.1 Event Study Estimates

To evaluate possible differential pre-trends between treatment and control groups, we perform an event study version of Equation (5.3). To do this, we generate five pre- and three post-lockdown periods (as dummy variables) on a quarterly basis. As work separations may be seasonal, and we want to exploit all the time horizon we have available (from January 2019 to December 2020), we construct the pre/post-policy periods using "quarterly base" categories. We have five pre-lockdown and three post-lockdown time periods. We set the period before the policy implementation ($t = -1$) as the omitted category (normalized to 0) (Freyaldenhoven et al., 2019; Fuest et al., 2018; Schmidheiny and Siegloch, 2022). In addition to exploring potential pre-trends, event study designs are useful for examining dynamic effects after the onset of the policy. Formally, we estimate an equation of the form:

$$Y_{spgmt} = \alpha + \beta_1 Female_g + \sum_{j=-5}^{2} \beta_2^j Post_{mt=j} + \sum_{j=-5}^{2} \beta_3^j Treat_s * Post_{mt=j} + \sum_{j=-5}^{2} \beta_4^j Female_g * Post_{mt=j} + \beta_5 Female_g * Treat_s + \sum_{j=-5}^{2} \beta_6^j Female_g * Treat_s * Post_{mt=j} + \delta_s + \rho_p + \epsilon_{spgmt} \tag{5.4}$$

In this specification, we include the same fixed effects and compute standard errors as in Equation (5.3). Figure 5.2 depicts the set of pre- and post-lockdown coefficients of the triple interaction (DDD) for the outcome of interest along with their 95 percent confidence intervals.

Figure 5.2: Event Study Graphs of Baseline Estimates



Notes: Own elaboration using administrative Ecuadorian social security data on the universe of new unemployment spells for formal private sector employees. Figure 5.2 plots the set of coefficients that correspond to the triple interaction terms from the event study of Equation (5.4). Reference category is one period (quarter) prior to the lockdown policy and is omitted from the graph. Confidence intervals at the 95 percent level.

Overall, we do not find evidence of differential pre-trends, which adds validity to the parallel trends assumption.[11] Regarding the post-lockdown period, we see that the increase in unemployment spells is particularly strong during the first three months (dummy label as "0" on x axis) following the mobility restrictions. After this, the coefficient decreases and becomes negative. From June (post 1 dummy) until December 2020, we do not find any significant coefficient. This result suggests that the mobility restrictions seem to be particularly problematic in terms of new

[11]The F test of jointly significance of the set of pre-lockdown dummies is 2.0134, from which we cannot reject the null hypothesis, $H_0 = \beta_6^{-5} = \beta_6^{-4} = \beta_6^{-3} = \beta_6^{-2} = 0$

unemployment spells during the so-called first wave of Covid-19 in Ecuador. The relaxation of the imposed quarantine, which started at the beginning of June 2020, is compatible with the zero effect seen in the last periods of our analysis.

5.6.2 Excluding Provinces

In Ecuador, according to the National Employment Survey (ENEMDU in Spanish) for September 2018, the city of Guayaquil, located in the Guayas province, accounted for approximately 24% of urban workers, and the capital, Quito (located in the province of Pichincha), registered 17.38% of urban workers (INEC, 2018). In this sense, one may wonder whether our results are driven by these two large provinces (in terms of number of workers). To answer this question, we re-estimate Equation (5.3) excluding unemployment spells that occurred in both the Guayas and Pichincha provinces.

To support the empirical strategy, one should still find positive and significant increases in the (log) number of unemployment spells. From Table 5.4, we see that after excluding unemployment spells of workers living in the Guayas province, the negative result remains and corresponds to an increase of approximately 14.21% in the number of new unemployment spells for women working in the restricted economic sectors. We find something similar when we exclude the Pichincha province (where the capital of the country is located). Here, we find a similar number in terms of magnitude and direction of the estimated coefficient. Therefore, even if we do not consider the biggest provinces, we still find a detrimental relation between jobs lost for female employees in the restricted economic activities and the lockdown policy.

5.6.3 Placebo

To indirectly assess the common time trend assumption, we conduct a placebo test in which we change the timing of the lockdown policy and perform a series of regressions of the form of Equation (5.3).[12] Here, we suppose that the event takes place one, two, three, and up to thirteen months before the actual lockdown date (March 2020), and we exclude all observations from March 2020 onwards. To add support to the parallel trends assumption, we should not find any significant coefficient in this placebo test.

Figure 5.3 plots the DDD coefficients along with their 95 percent confidence intervals where each point estimate corresponds to a separate regression. From

[12] All regressions are estimated using the same fixed effects as in our baseline setting. Standard errors are also computed in the same way.

Table 5.4: Robustness Check - Excluding Guayas and Pichincha Provinces

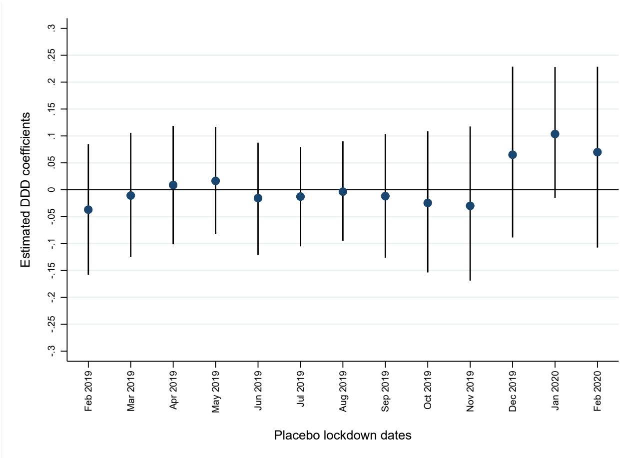
	(1)	(2)
	Excluding Guayas (log) unempl spells	**Excluding Pichincha (log) unempl spells**
DDD Female*Treat*Post	**0.1421****	**0.1400***
	[0.0420]	[0.0651]
Interaction Female*Post	-0.0735	-0.0727
	[0.2723]	[0.3013]
Interaction Female*Treat	-0.3528	-0.3560
	[0.3634]	[0.3834]
Interaction Treat*Post	0.0118	0.0092
	[0.8859]	[0.8999]
Province FE	x	x
Economic activity FE	x	x
Month FE	x	x
Year FE	x	x
Observations	19,338	19,337
Number of clusters	21	21

Notes: OLS estimates of Equation (5.3) excluding unemployment spells from the Guayas (column 1) and Pichincha (column 2) provinces. P-values refer to standard errors clustered at the economic activity level in squared brackets, calculated using the wild cluster bootstrap with 999 replications. **p<0.05, *p<0.10.

here, we see that there is no statistically significant coefficient at any of the fake event dates. In addition, all the coefficients are smaller (and some are very close to zero) than the actual estimated effect, which adds validity to our identification strategy.

Altogether, we find evidence that supports the robustness of our findings. For instance, our results, overall, do not show evidence of pre-trends (seen in the event study), keep their significance, and have a similar magnitude when excluding the most populated provinces (in terms of the working population). In addition, our results are also robust to a placebo test, in which we change the timing of the lock-down policy.

Figure 5.3: Robustness Checks - Placebo Event Dates



Notes: Own elaboration using administrative Ecuadorian social security data on the universe of new unemployment spells for formal private sector employees. Each point estimate (and its 95 percent confidence intervals) represents a separate regression in which the event date is set according to the x axis.

5.7 Other Labor Market Results

5.7.1 Probability of Being Employed and Cause of Job Loss

So far, we find evidence that supports an increase in unemployment spells for women working in the restricted sectors relative to other groups of workers. Now, we go a step further and explore whether there are any changes in other labor market outcomes such as the probability of being employed, full time jobs, and cause of job loss.

We use survey data taken from the "National Survey of Employment" of Ecuador (ENEMDU, for its name in Spanish), released by the National Institute of Statistics (INEC) for the years 2019 and 2020. The data is a pooled cross-section at the individual and quarter level. Each cross section consists of individual level data. The ENEMDU constitutes the official source for calculating labor market indicators in the country and is nationally representative, including both urban and rural areas (INEC, 2022). Regarding labor market information, the survey offers data on employment, job searches, hours worked, and cause of job loss, among others.

For this analysis, we use three labor market outcomes: (i) the probability of being employed, (ii) the probability of having a full-time job, and (iii) the cause of job loss (conditional on having moved from employment to unemployment). For the first outcome, we construct a dummy variable that takes the value of one when the individual reports that he/she worked the previous week. In this question, all types of jobs such as full- or part-time are included, and zero corresponds to unemployed respondents. For the second outcome, we construct a binary variable when the individual declares to have worked 40 hours or more (which corresponds to full-time jobs) and zero otherwise. In this way, we measure the probability of having worked 40 or more hours the previous week. With the last outcome, we intend to assess whether women working in the restricted economic sectors voluntarily left their jobs[13] or not. We construct a dummy variable that takes the value of one when leaving work corresponds to a voluntary act and zero when it corresponds to being fired.

As in the baseline estimates of this chapter, we focus on individuals who are between the ages of 15 and 65 and work for the private sector. We exclude the agricultural and mining sectors, domestic workers, and self-employed individuals.

To formally explore these labor market outcomes, we rely on the same econometric approach as in Equation (5.3). We use a DDD approach which compares restricted and non-restricted economic sectors (first difference), women and men (second difference), before vs after the lockdown policy (third difference). We estimate an equation of the form:

$$\begin{aligned} Y_{igst} = {} & \alpha + \beta_1 Female_g + \beta_2 Post_t + \beta_3 Treat_s * Post_t + \beta_4 Female_g * Post_t \\ & + \beta_5 Female_g * Treat_s + \beta_6 Female_g * Treat_s * Post_t \\ & + \beta_7 Deaths_t + \delta' X_i + \rho_s + \tau_t + \epsilon_{igst} \end{aligned} \tag{5.5}$$

where Y_{igst} represents one of the three labor market outcomes. $Female_g$ takes the value of one for women and zero for men. $Post_t$ is a dummy variable for the pre-post lockdown, and $Treat_s$ takes the value of one if the economic activity in which the individual works was restricted. In addition, we include a set of control variables such as the share of excess deaths (at the country level) $Deaths_t$. At the individual level, we control for education attainment, marital status, age group, and urban/ rural location, $\delta' X_i$.[14] Control and treatment groups are defined and constructed as in Equation (5.3). We cluster standard errors at the economic activity level, and compute them using the wild cluster bootstrap with 999 replications. We use probability weights to account for the survey design.

Table 5.5 presents the results for the exercise described above. Column (1) shows a decrease of 5.76 percentage points in the probability of being employed for women in restricted economic activities relative to other groups of employees after the lock-

[13]Voluntary termination occurs when an employee makes the decision to leave a job or end a contract early.

[14]The ENEMDU for the period of analysis does not provide desegregation at the province level. Therefore, we can only control for the urban or rural place of residence of the respondents.

down policy. This represents a decrease of 6.35% with respect to the pre-policy mean. With this result, we support the statement that the lockdown policy might have disproportionately affected women compared with men. In addition, we find evidence that points to a deterioration in employment and not only an increase in unemployment spells for our group of interest.

Table 5.5: Other Labor Market Results

	(1)	(2)	(3)
	Employed	**Full time job**	**Voluntary**
DDD Female*Treat*Post	**-0.0576***	**-0.0084**	**-0.1231***
	[0.0721]	[0.8328]	[0.0831]
Interaction Female*Post	0.0093	0.0537	-0.0113
	[0.7287]	[0.1962]	[0.7718]
Interaction Female*Treat	0.0152	-0.0287	0.0518
	[0.2282]	[0.5475]	[0.3954]
Interaction Treat*Post	-0.0030	-0.0544**	0.0775
	[0.8669]	[0.0440]	[0.2302]
Share excess deaths	-0.1058***	-0.2086***	0.0162
	[0.0000]	[0.0000]	[0.5556]
Mean dependent variable	0.9077	0.5600	0.4224
% Effect	**-6.35%**	**-1.5%**	**-29.14%**
Education level	x	x	x
Marital status	x	x	x
Age category	x	x	x
Urban/rural	x	x	x
Economic activity FE	x	x	x
Time FE	x	x	x
Observations	43,251	43,251	3,784
Number of clusters	22	22	21

Notes: Estimates from Equation (5.5). All outcome variables are dummy indicators. The time horizon includes the years 2019 and 2020 at the quarter level. The year 2020 has no information during the first trimester. P-values refer to standard errors clustered at the economic activity level in squared brackets, calculated using the wild cluster bootstrap with 999 replications. ***p<0.01, **p<0.05, *p<0.10.

Column (2) presents the results for the probability of working 40 hours or more the previous week. Here, even we do not find a significant result, it points to the expected side, that is, a decrease in the share of women (in the restricted economic sectors) working full-time. Finally, Column (3) shows that the probability of a voluntary termination of the work contract is lower for female employees (working in the restricted economic sectors) compared with other groups of workers (however, as we will show in the following paragraph, we do not make conclusions based on this outcome because of the existence of a pre-trend). The difference represents a decrease of around 29.14% in the share of voluntary job loss with respect to the pre-lockdown mean. In spite of the slightly significant result of outcome 1, we still consider it is interesting to further explore it across socioeconomic characteristics, an exercise that we will perform later in this chapter.

To add support to the validity of the results of this section, we perform an event study version of Equation (5.5) for the three outcomes. In Figure 5B.3 of the Appendix 5B, we plot the set of DDD coefficients along with their 90 percent confidence intervals. We include four pre- and three post-lockdown periods at the quarter level for each outcome. The post lockdown periods are the second, third and fourth quarters of year 2020.[15] In these regressions, we use the same control variables, fixed effects, and method for computation of standard errors as used to obtain the estimates in Table 5.5. Overall, we do not find evidence of pre-trends in any of the outcomes, except for voluntary job loss, outcome for which we cannot make more conclusions. For the probability of being employed, in panel (a), none of the pre-policy coefficients are significant and are of a very small size (close to zero). For the post period, we find one coefficient significant at the 10% level, which is compatible with the corresponding result of column (1) of Table 5.5. A similar pattern is seen for the probability of working full-time. All pre-event estimates are very close to zero and not significant. With respect to post-policy estimates, we do not find significant results, but signs for the last periods are as expected.

Overall, we find evidence that the COVID-19 lockdown is associated with an increase in job loss for women working in the affected sectors compared to other workers. Furthermore, we find a significant increase in the number of unemployment spells, a result that is supported by a general decrease in the probability of being employed. According to the International Labor Organization ILO (2021), women suffered a disproportionate loss of employment and income because of their prominent presence in the most affected sectors, such as accommodation and catering services. In addition, the ILO states that The Americas region experienced the greatest loss of female employment as a result of the crisis (-9.4 %).

5.7.2 Heterogeneity Analysis: Age Groups and Education Level

In the previous subsection, we found evidence that suggests a decrease in the probability of being employed for women working in the restricted economic sectors.

[15]The periodicity of the ENEMDU survey is quarterly.

As a consequence, one may wonder whether there are differences across socioeconomic characteristics. To answer this question, we now focus only on the first outcome (probability of being employed) because the "cause of job loss" showed a pre-trend. We begin by splitting the sample into age group categories and construct four groups: (i) 15-24 y/o, (ii) 25-34 y/o, (iii) 35-44 y/o, and (iv) 45-65 y/o, where each group corresponds to 20%, 33%, 23%, and 22% of the total number of respondents, respectively. Formally, we estimate Equation (5.5) for the four different (age) sub-samples. Table 5.6 shows the results of this exercise.

Table 5.6: Heterogeneity Analysis - Age Groups

	Outcome: Probability of being employed			
	15- 24 y/o	**25-34 y/o**	**35-44 y/o**	**45-65 y/o**
DDD Female*Treat*Post	**-0.1187****	**-0.0232**	**-0.0051**	**-0.1569****
	[0.0470]	[0.5095]	[0.9109]	[0.0220]
Interaction Female*Post	0.0394	-0.0058	-0.0330	0.0775
	[0.2873]	[0.8188]	[0.1902]	[0.1031]
Interaction Female*Treat	0.0440	0.0139	-0.0102	0.0274
	[0.1742]	[0.4755]	[0.5566]	[0.1241]
Interaction Treat*Post	0.0134	-0.0043	-0.0240	0.0106
	[0.4264]	[0.8398]	[0.2432]	[0.7948]
Share excess deaths	-0.1408***	-0.1183***	-0.0851***	-0.0822***
	[0.0000]	[0.0000]	[0.0000]	[0.0000]
Education level	x	x	x	x
Marital status	x	x	x	x
Urban/rural	x	x	x	x
Economic activity FE	x	x	x	x
Time FE	x	x	x	x
Observations	8,742	14,281	10,191	10,037
Number of clusters	22	22	22	21

Notes: The outcome variable is the probability of being employed. The time horizon includes the years 2019 and 2020 at the quarter level. The year 2020 has no information during the first trimester. P-values refer to standard errors clustered at the economic activity level in squared brackets, calculated using the wild cluster bootstrap with 999 replications. ***p<0.01, **p<0.05.

For all four age groups, we obtain the expected signs, that is, a decrease in the probability of being employed the previous week. Interestingly, we find that the biggest coefficient corresponds to the oldest group (45-65 years old) of women working in the restricted economic sectors. We find a decrease of 15.69 percentage points in the probability of being employed. Moreover, we also find a statistically significant decrease for the youngest group, of around 11.87 percentage points.

Under the Covid-19 economic shock context, older groups of workers seem to be more negatively affected (Bui et al., 2020). This group might have more difficulties adapting to new technologies (Pit et al., 2021), which might place them in a more vulnerable position and therefore increase their probability of job loss in a situation such as the Covid-19. A plausible explanation for these findings may be that, in Ecuador, older individuals are concentrated in the economic sectors that were restricted because of the lockdown. For instance, according to data from ENEMDU year 2019, 35.6% of employees working in these sectors were between 45 and 65 years old. Likewise, older women represent around 51.38% of the workforce in the wholesale sector, 71.84% in accommodation and food service activities, and 63.78% in the education sector. These sectors were the most severely affected because of the nature of their activities.

Younger groups of workers are also affected by the lockdown. For example, Abraham et al. (2022), using data from India, find that women were seven times more likely to lose work during the nationwide lockdown, and in a sample stratified by gender, found that young workers, whether men or women, were more likely to face job loss. Gustafsson (2020) found, for the United Kingdom, that both younger and older workers experienced the brunt of the hit to jobs during the Covid-19 crisis. In the same line, Bui et al. (2020), using “Current Population Survey” data from the U.S., found that the recession generated by the Covid-19 crisis (during its “first wave”) disproportionately affected older workers. In addition, these authors also found that women reached higher unemployment rates than men across all age groups. Finally, in research conducted for Mexico, Hoehn-Velasco et al. (2021) found that the youngest workers, the oldest workers (over 60 years old), and low-income earners were the most impacted in terms of formal employment.

The result for older workers is of particular interest because they may face more difficulties to reenter the labor market. For instance, they may encounter more age discrimination in hiring. It may also take longer for older workers to find a job (Neumark and Button, 2014), which is particularly problematic as this could contribute to a decline in earnings, less savings, less access to health care use, expanding working years before retirement, and reductions in quality of life (Bui et al., 2020).

We continue to explore whether there are differences across workers’ education levels. To do this, we split the sample into three groups: (i) basic education, which corresponds to up to 10 years of studies, (ii) secondary education, which corresponds to 13 years of studies and the end of non-tertiary education, and (iii) tertiary education, which corresponds to a professional degree. The composition of the sample is as follows: 26.20% of individuals have a basic education, 46.21% have a high school diploma, and 26.84% have a professional degree. As we did for the age group estimates, we use the same econometric specification as in Equation (5.5) for the three sub-samples where the outcome is the probability of being employed.

Table 5.7 shows the heterogeneity across education level groups. For all groups, we get the expected signs, that is, an average decrease in the probability of being employed. However, less educated workers seem to be driving the results. In particular, we find that the group of individuals with basic education experienced a decrease of 12.78 percentage points in the probability of being employed. These results provide evidence that supports that less educated women working in the restricted economic sectors might be disproportionately affected by the lockdown. One explanation for this might be that less educated women may be more concentrated in economic sectors that require more in-person interactions and therefore would be more affected by mobility restrictions.

Table 5.7: Heterogeneity Analysis - Education Groups

	Outcome: Probability of being employed		
	Basic	**Secondary**	**Tertiary**
DDD Female*Treat*Post	**-0.1278****	**-0.0456**	**-0.0218**
	[0.0180]	[0.2392]	[0.4074]
Interaction Female*Post	0.0156	0.0009	0.0347
	[0.6426]	[0.9690]	[0.1131]
Interaction Female*Treat	0.0250	0.0214	0.0137
	[0.3213]	[0.2492]	[0.2853]
Interaction Treat*Post	0.0018	-0.0102	-0.0208
	[0.9730]	[0.4645]	[0.5415]
Share excess deaths	-0.0868***	-0.1113***	-0.1117***
	[0.0000]	[0.0000]	[0.0000]
Age category	x	x	x
Marital status	x	x	x
Urban/rural	x	x	x
Economic activity FE	x	x	x
Time FE	x	x	x
Observations	10,851	19,251	12,883
Number of clusters	20	22	22

Notes: The outcome variable is the probability of being employed. The time horizon includes the years 2019 and 2020 at the quarter level. The year 2020 has no information during the first trimester. P-values refer to standard errors clustered at the economic activity level in squared brackets, calculated using the wild cluster bootstrap with 999 replications. ***p<0.01, **p<0.05.

Our results are in line with others in the literature. For instance, Escalante and Maisonnave (2022), using data from Bolivia, find that women suffered more than men from the negative impacts of the Covid-19 crisis in terms of employment. In particular, they find that unskilled women were the most affected, as they experienced the greatest losses of formal employment. Likewise, Adams-Prassl et al. (2020) found that women and less educated workers were more affected by the crisis in the German context. Holder et al. (2021) found evidence that supports greater job losses for less educated workers in the U.S. Likewise, Lee et al. (2021) show that the initial negative impact on employment was larger for women, minorities, and the less educated (also for the U.S.). Finally, Aldan et al. (2021), using quarterly Turkish household labor force surveys, found that the pandemic decreased employment and labor force participation of almost all groups, with women more affected than men, along with the least educated.

5.8 Conclusions

The Covid-19 crisis has seriously affected the world in various dimensions. Governments isolated their citizens by imposing a strict quarantine referred to as a "lockdown", which produced an economic crisis whose effects on labor markets were particularly damaging. Ecuador also applied a lockdown and a number of activities considered "non-essential" were forced to stop their operations.

Economic crises may affect demographic groups differently. In this line, several studies conducted by international organizations show that women are disproportionately affected in terms of job loss, declining earnings and hours worked. In addition, women may also face delays reentering the labor force, which reinforces already existing gender gaps. For these reasons, assessing whether the lockdown affected women more than men in terms of job loss becomes relevant.

In this chapter, we evaluate the short-term relation of the lockdown imposed in Ecuador in March 2020 with labor market outcomes for women working in restricted economic sectors. Using a a triple difference-in-differences approach, we find that the lockdown policy is associated with an increase in the number of unemployment spells for the population under study. We also show that the probability of being employed decreased. These results are consistent with others in the literature that find women more negatively affected in terms of labor market outcomes as a consequence of the Covid-19 (economic) shock (Bluedorn et al., 2022).

We also add evidence that supports differences across age groups and educational level. We find that younger, older, and less educated women (in the restricted economic sectors) seem to have been more affected than their male counterparts. These results are consistent with Escalante and Maisonnave (2022), Hoehn-Velasco et al. (2021), and Serrano et al. (2019), who find that specific demographic groups were more affected, at least in the short term, after the Covid-19 lockdown.

The results are robust to a battery of robustness checks, which adds validity to the estimates. However, there are some limitations that stem mainly from data unavailability. First, we observe inflows into unemployment but not outflows. Because we are interested only in short-term relations and, in the immediate post-lockdown period, the hiring of new employees was very unlikely, we believe our results are still of relevance. Second, it would be ideal to work, as an outcome, with the share of unemployment spells over the working population. Unfortunately, to date, we have not gotten access to this information, so our results should be interpreted taking this into account. Third, we acknowledge that there might be other factors that influence both supply and labor demand such as increases in transportation costs, reductions in production capacity, and decreases in job searches, among others. Fourth, this evidence should be interpreted with caution, as it is correlational. Unfortunately, the labor market is different for men and women in Ecuador and this causes men and women to have different characteristics, so their comparison leads us to associative conclusions.

Despite these drawbacks, we believe that the lockdown policy was a channel associated with the increase in unemployment spells for female workers in the formal sector of the Ecuadorian economy. We believe our results are of relevance in assessing the short-term consequences of the Covid-19 lockdown on the Ecuadorian labor market.

Finally, providing evidence with regards to (formal) job loss in a context such as the Ecuadorian becomes not only interesting but relevant. Ecuador is a country with high shares of informality, which might have increased because of the transition from formal employment to unemployment. Moreover, at the individual level, having no formal job in Ecuador is linked to losing access to social security benefits such as health care, personal loans, etc. Likewise, older groups may face more barriers to rejoining the labor market, which may negatively affect earnings, savings, and health care use and increases the chances of falling into poverty. In this sense, having a clearer panorama of the labor market consequences of the lockdown policy is an important tool for policy makers to design programs or strategies focused on the most affected groups to objectively target public policies.

Appendix 5A: Additional Tables

Table 5A.1: Treated Economic Activities

ISIC	ISIC description	Treated
C - Manufacturing	C1101, C1102, C1200, C1312, C1313, C1393, C1394, C1399, C1410, C1420, C1430, C1511, C1512, C1520, C1622, C1629, C1812, C1820, C2391, C2392, C2393, C2394, C2395, C2396, C2399, C2511, C2592, C2652, C2680, C2750, C2814, C2815, C2817, C2818, C2826, C2910, C2920, C2930, C3011, C3012, C3091, C3100, C3211, C3212, C3220, C3230, C3240, C3290.	Yes
F - Construction	F4100, F4210, F4220, F4290, F4311, F4312, F4322, F4329, F4330, F4390.	Yes
G - Wholesale and retail trade	G4510, G4520, G4530, G4540, G4610, G4641, G4663, G4690, G4742, G4751, G4752, G4753, G4759, G4761, G4762, G4763, G4764, G4771, G4773, G4774, G4782, G4789.	Yes
H - Transportation and storage	H4911, H4921, H4922, H5011, H5021, H5110, H5221, H5222, H5223.	Yes
I - Accommodation and food service	I5510, I5520, I5590, I5610, I5621, I5629, I5630.	Yes
J - Information & commun.	J5913, J5914.	Yes
L - Real estate activities	L6810, L6820.	Yes
M - Professional, scientific and technical activities	M7010, M7020, M7110, M7120, M7210, M7220, M7320, M7410, M7420, M7490.	Yes
N - Administrative and support service activities	All divisions except N7229, N7730, N8220, N8291, N8292, and N8299.	Yes
P - Education	All divisions.	Yes
R - Arts, entertainment and recreation	All divisions.	Yes
S - Other service activities	All divisions except S9603.	Yes

Notes: Own elaboration using data from the National Institute of Statistics of Ecuador (INEC, 2020). English version downloadable at $https://unstats.un.org/unsd/publication/seriesm/seriesm_4rev4e.pdf$

Table 5A.2: Robustness - Gender and Economic Sector Time Trends

	(1)	(2)
	Log. unemployment spells	**Log. unemployment spells**
DDD Female*Treat*Post	**0.1513****	**0.1489****
	[0.0480]	[0.0490]
Interaction Female*Post	-0.0535	-0.0756
	[0.4274]	[0.2963]
Interaction Female*Treat	-0.3445	-0.3424
	[0.3754]	[0.4024]
Interaction Treat*Post	0.0032	0.0064
	[0.9710]	[0.9249]
Share excess deaths	-0.0000*	-0.0000
	[0.0821]	[0.1371]
Province FE	yes	yes
Economic activity FE	yes	yes
Month FE	yes	yes
Year FE	yes	yes
Gender-specific time trend	yes	no
Economic sector time trend	no	yes
Observations	20,391	20,391
Number of clusters	21	21

Notes: OLS estimates of Equation (5.3). P-values refer to standard errors clustered at the economic activity level in square brackets, calculated using the wild cluster bootstrap with 999 replications.
**p<0.05, *p<0.10.

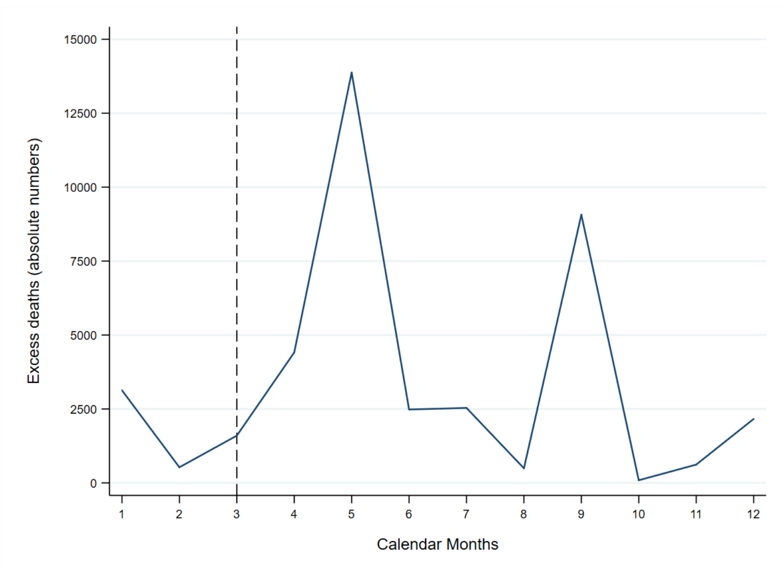
Table 5A.3: Mean Gender Differences across Treated and Untreated Sectors

	Treated sectors			**Untreated sectors**		
Variable	**Males (0)**	**Female (1)**	**Difference (2)=(0)-(1)**	**Males (0)**	**Female (1)**	**Difference (2)=(0)-(1)**
Low educated	0.3429	0.1993	0.1435***	0.2219	0.1261	0.0958***
	(0.0034)	(0.0040)	(0.0056)	(0.0044)	(0.0043)	(0.0064)
	[18,985]	[9,667]	[28,652]	[8,755]	[5,844]	[14,599]
Monthly payment	8,743.748	12,779.35	-4,035.604***	14,448.06	15,052.61	-604.5552
	(685.4696)	(1,192.818)	(1,286.946)	(1,285.901)	(1,645.782)	(2,078.492)
	[17,426]	[8,534]	[25,960]	[8,241]	[5,256]	[13,497]
Blue collar jobs	0.6893	0.4869	0.2024***	0.5487	0.3829	0.1657***
	(0.0033)	(0.0050)	(0.0059)	(0.0053)	(0.0063)	(0.0083)
	[18,985]	[9,667]	[28,652]	[8,755]	[5,844]	[14,599]

Notes: Variable low educated corresponds to the share of individuals that have less than 10 years of education. Monthly payment corresponds to average monthly employment income. Blue collar jobs corresponds to the share of individuals performing manual labor. Standard errors are in parenthesis and the number of observations is in square brackets. ***p<0.01

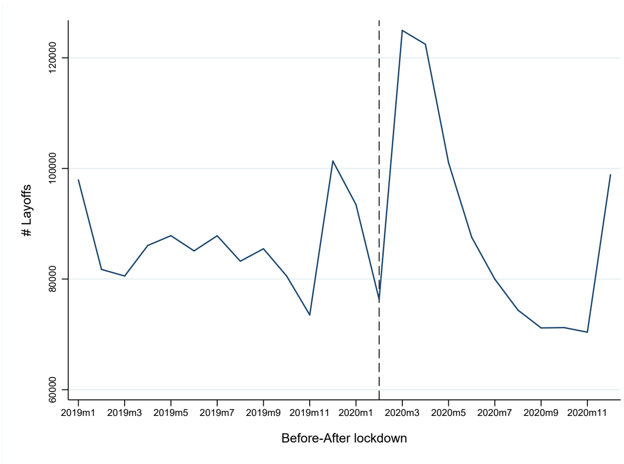
Appendix 5B: Additional Figures

Figure 5B.1: Excess Deaths per Months - year 2020



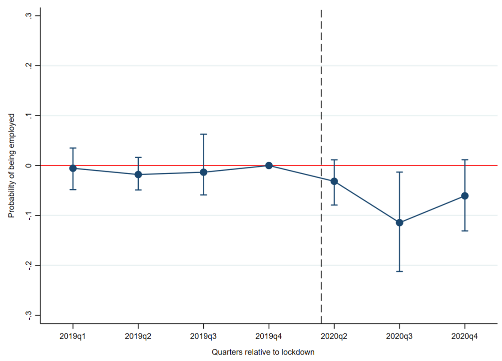
Notes: Own elaboration using administrative data from national birth and deaths registration authority "Registro Civil del Ecuador". The time period includes monthly excess deaths from the year 2020 across the whole country.

Figure 5B.2: Evolution of Unemployment Spells for the Entire Economy

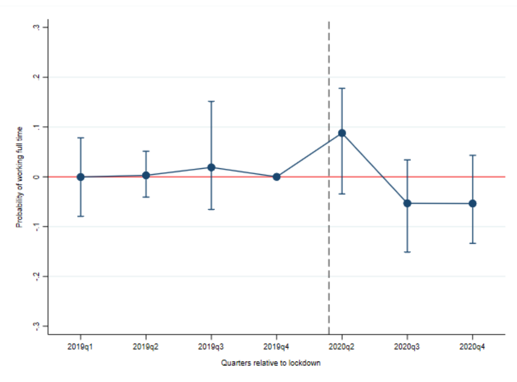


Notes: Own elaboration using administrative data from social security records.

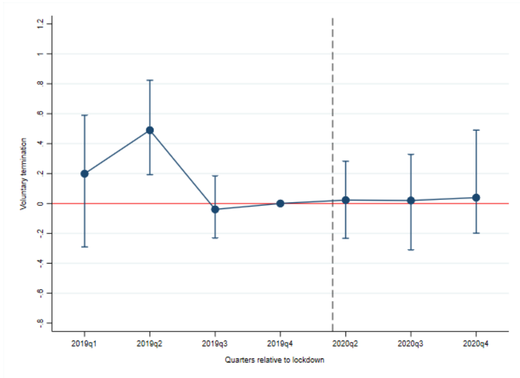
Figure 5B.3: Event Study Graphs of "Other Labor Market Results"



(a) Employed



(b) Full time



(c) Voluntary termination

Notes: Own elaboration using survey data from ENEMDU for the years 2019-2020. Figure 5B.3 shows the set of coefficients that correspond to the triple interaction term from the event study versions of estimates of Table 5.5. Time periods are defined on quarterly basis. Points estimates are shown along with their 90 percent confidence intervals.

6 Concluding Remarks

6.1 Summary of Findings

This dissertation contributes to two research lines of the health economics literature, newborn's health and workers health. Chapters 2 and 3 focus on newborn's health and analyze the effects of exogenous shocks on women and its in-utero transmission with resulting detrimental health for newborns. Chapter 4 builds on the effect of the civil servant selection process on workers health, which we approach using sickness absence episodes in Spain. Chapter 5 explores the relation between the lockdown policy caused by the Covid-19 pandemic and female employment, which we proxy using inflows from employment to unemployment.

In the development of this dissertation, we generated data and tools that are available for the research community. In Chapter 3, I create a composite index, which measures the intensity of exposure to illicit drugs in the year 2012 of each of the provinces of Ecuador. The steps to generate the index are detailed in Appendix 3A.[1] In Chapter 4, we build a unique database, which we generated from official administrative records, on the universe of public employment calls in Spain during years 2009 to 2015. This dataset comprises all calls for public examinations at the national, regional and local level.

Chapter 2, *"Terrorist Attacks, Islamophobia and newborns' health"*, focuses on the impact that jihadist terrorist attacks may have on newborn's health, as a consequence of increased emotional distress in Muslim mothers subject to Islamophobia. This chapter provides robust evidence that Islamophobia-related stress is possibly one of the channels affecting health at birth. This chapter contributes to the literature on discrimination and health, and on maternal stress and birth outcomes.

The main findings, estimated using a difference-in-differences-in-differences approach, document a significant increase in the share of low birth weigh newborns to Muslim mothers living in the cities that were affected by the attacks, and an increase in the share of deliveries with complications, with the effect being different throughout the three months of gestation. In addition, we also document an increase in the rejection against Muslim, known as Islamophobia, which is associated with

[1] The data used to generate the index is available upon request to the competing authorities.

increased emotional distress for Muslim mothers (Bader and Berg, 2013). The intuition is that the terrorist attacks perpetrated by a jihadist group in year 2017 in Catalonia-Spain, is a source of increased Islamophobia, which is particularly problematic for Muslim pregnant women, because this may increase stress, which has been found to be negative for the development of the fetus (Black et al., 2016; Persson and Rossin-Slater, 2018). Finally, the results are robust to several robustness checks such as an event study version of the main equation and other placebo exercises.

Chapter 3, *"Drug Tenancy, Thresholds, Consumption and Newborn's Health"*, finds a negative relation between women's drug use/abuse and health at birth. This chapter explores how a public policy in Ecuador designed not to criminalize drug users, but rather distinguish them from dealers, and in this way, recognize their rights in terms of need for health care, may have unintended negative consequences in drug consumption among women of childbearing age, which in turn, may be transmitted to the developing fetus, resulting in bad health at birth.

Formally, I use a difference-in-difference approach, which exploits the variation in the initial drug exposure across provinces, before versus after the enacting of the policy. To measure drug exposure, I develop a composite index at the province level, which is a continuous variable. Therefore, the treatment is also continuous. The results reveal that newborns to women living in provinces that were initially more intense in drug exposure are more likely to have low and very low birth weight. They also have, overall, higher hospitalization rates, especially those related to maternal drug use. The last result is compatible with that of low birth weight, because worse health at birth might induce higher probabilities of neonatal hospitalizations (Creanga et al., 2012; Hwang et al., 2017). In terms of hospital mortality rates, I do not find any significant difference, which is compatible with the absence of a significant result for preterm births (Perin et al., 2022). With respect to women's health, this chapter documents an increase in the hospitalization rate to conditions related to drug use/abuse. This chapter presents several robustness checks. For example, I do not find evidence that suggests any changes in fertility responses and prenatal care use. I also perform a falsification test changing the date of the policy implementation to rule out any anticipatory effect.

This chapter contributes to a growing literature evaluating the effect of drug decriminalization policies on health-related outcomes, however, this has been mainly focused on legalization of drugs, which is not the aim of the policy I analyze in this chapter. Moreover, most of the work is focused on marijuana liberalization and on the opioid crisis in the U.S (Williams and Bretteville-Jensen, 2014; Laqueur et al., 2020; Kim, 2021). This chapter also contributes to the literature on the determinants of child and adulthood outcomes.

In Chapter 4, *"Job Competition in Civil Servant Public Examinations and Sick Leave Behavior"*, we investigate how the civil servant recruitment system used in Spain may affect workers' health. We do so by measuring changes in work sickness absences months before the public examinations take place. The intuition is that the tough process, which demands high levels of study, might be generating increased levels of stress/anxiety among participants who combine work and study at the same time, which may be reflected in an increase of the absence rate.

Empirically, we implement an event study framework to examine whether there is an increase in the absence rate before the public examinations and its average length. The findings point to a significant increase in health-related absences months before the exams, which is stronger for workers of the education sector. In addition, the effect is mostly driven by stress-related conditions. We also provide evidence that supports a deterioration of temporal public sector workers' health. The results of this chapter are robust to several tests. This chapter contributes to the literature on workers' health, and on the negative consequences of the civil servant recruitment system, which is still under-explored.

The last empirical chapter of this dissertation, *"Covid-19 Lockdown in Ecuador: Are there Gender Differences in Unemployment?"* sheds light on the short-term negative consequences that the lockdown imposed to stop the spread of the contagious of the Covid-19 had on women's employment in Ecuador. Using a difference-in-differences-in-differences framework, we find strong evidence that suggests a positive association between the lockdown and unemployment spells for female workers of the restricted economic sectors. In addition, the lockdown is also associated with a lower probability of being employed for women. The groups that seem to be more affected are the least educated, youngest and oldest workers. This chapter adds to the rich literature on Covid-19 and labor market outcomes. Specifically, we contribute to the analysis of labor markets of developing economies with a gender focus, which has been scarcely studied. The findings are robust to several tests.

Three key lessons arises from this dissertation. First, mother's health and behaviors are important for newborn's health. Mother's health conditions can be affected not only by traumatic events such as terrorist attacks, but also as unintended consequences of public policies designed for different purposes. Second, workers health is very sensitive to common-practice procedures of civil servants selection, which have the potential to affect workers mental health. Third, unexpected health shocks, such as the Covid-19 pandemic, may importantly influence socioeconomic outcomes such as employment, which can be particularly harmful for female workers.

6.2 Policy Implications, Limitations, and Further Work

The Unicef (2017) establishes that the first 1000 days of life, the time spanning roughly between conception and one's second birthday, *"is a unique period of opportunity when the foundations of optimum health, growth, and neurodevelopment across the lifespan are established"*. The right health conditions during this period of time are important determinants for child survival, children's ability to grow, learn and rise out of poverty, which also contributes to society's prosperity as pointed out by Deaton (2013).

According to the previous definition, the period of pregnancy constitutes the beginning of future health and well-being, therefore early life shocks are important for many socioeconomic outcomes in adulthood such as education achievement, earnings and quality of life. Children that experienced bad health at birth, as measured by LBW, are more liked to experience developmental difficulties, increased risk of infant morbidity and mortality (Lee et al., 2017; Lambiris et al., 2022). Following Barker (1990) and its original work on the *"fetal origins hypothesis"*, positive associations between negative early life shocks and adult risk of chronic disease have been documented in humans. Likewise, fetal development may also influence human capital outcomes across the life cycle.

The findings of Chapters 2 and 3 add to the research on the effects of maternal stress and substance/use abuse on newborn's health. These chapters find detrimental effects in terms of outcomes such as low birth weight, very low birth weight, and neonatal hospitalizations. These results raise concerns about the inequalities in health that may widen for already disadvantaged groups, such as immigrants (Chapter 2), and new inequalities that may arise for the new generations (Chapter 3). Consequently, these results become an important tool for public policy, which may help governments to identify vulnerable groups that are in need of health care.

At this point, we must recognize that there are some limitations in our work, which mainly come from data availability restrictions. In Chapter 2, we are not able to assess other potential channels affecting health at birth such as changes in prenatal care use and labor market outcomes, which may be affected after the terrorist attacks. In Chapter 3, I am not able to evaluate whether exposed women are more prone to risky behaviors that may be harmful for the fetus. In addition, consumption of some type of illicit drugs is also associated with feeding problems of the mother, which can decrease the nutrients that the fetus needs for adequate development.

The results from these two chapters open the floor for future research in terms of analysing long-term negative effects of in-utero exposure to stress and drugs of abuse. For instance, its effects on childhood outcomes such as health, education achievement; and on adulthood outcomes like earning and health.

A different line of the health economics literature studies workers health. Hermann and Rockoff (2012) find that worker absences have large negative impacts in terms of productivity, mainly coming from hours lost and the replacement with a new employee. In Chapter 4, we find evidence supporting that the civil servant recruitment process may be generating negative unintended effects in terms of sickness absences, which is stronger for individuals working in the education sector. We also find that stress-related conditions seem to be the most common for sickness absences before public examinations. These findings also shed light on the potential negative consequences for worker's mental health, which in turn, may also be detrimental for labor productivity.

There are some limitations in the results of Chapter 4. Despite we provide evidence that supports an increase in stress among applicants who combine work and exam preparation, there is still room for our results to be coming from a strategic behavior from participants, who may take advantage of the social security system to access sick leave. Our results also open the way for future research to explore and implement improvements in the design of the civil service recruitment system.

Finally, with respect to the findings of Chapter 5, losing formal employment in the Ecuadorian context represents lack of access to health care and medicines. It also imposes barriers to cheaper credit access, and extend the years before retirement. Our findings support a strong association between the Covid-19 lockdown and short-term job loss, with a stronger consequence for oldest, youngest, and less educated women. This is particularly problematic, because older individuals may find more difficult to re-join the labor market, and puts them in higher risk of poverty.

In this last empirical chapter, we also faced some limitations. For instance, women and men are very likely to be different in characteristics that may be important for labor market outcomes. While, we provide several robustness checks, we must interpret the results as strong associations only. It would also be ideal to explore a longer time horizon, however, due to data limitations, we restricted the analysis to the end of year 2020. This last limitation becomes into an opportunity for further research on the long-term effects of the Covid-19 lockdown on labor market outcomes.

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