Three Essays on the Role of Federalism in the Swiss Healthcare System

Paola Salari

Faculty of Economics

Università della Svizzera Italiana
(Lugano, Switzerland)

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Doctoral committee:

Prof. Luca Crivelli, Università della Svizzera Italiana (Lugano), Internal member

Prof. Fabrizio Mazzonna, Università della Svizzera Italiana (Lugano), Supervisor

Prof. Eddy van Doorslaer, Erasmus University (Rotterdam), External member

To my parents, who taught me to walk

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Introduction

Switzerland is a country with a strongly decentralized political system, based on federalism and institutions of direct democracy, on a liberal economic culture, and on a well–developed tradition of mutualism and social security (generous social expenditure and welfare system). Switzerland is unique for its high level of decentralization and for the particular distribution of competences and roles between the central State (the Confederation) and the cantonal authorities. This particular setting allows cantons to have significant leeway in the organization of health care and the implementation of specific public policies.

My thesis aims to give a contribution to the health economics literature, assessing some of the internal implications of the Swiss federal setting in the health care sector (chapters 1 and 2) and drawing some more general public policy results, exploiting the Swiss context (chapter 3).

The first two chapters explore two different sides of the equity of the Swiss health care system. In the first chapter, I assess to what extent the political autonomy allocated to cantons leads to differences in the level of regressivity in the financing of the health care system and over time. The second chapter investigates the role of managed care contracts and higher deductible on the equity in health care utilization. Finally, in the third chapter I carry out a policy evaluation analysis to assess the causal effects of smoking bans on acute myocardial infarction. Different from the other two chapters, this work draws conclusions that abstract from the Swiss context. Switzerland is the perfect setting where to employ our empirical strategy, due to the different times of implementation of the policy if interest. However, the results have external validity.

Each work is based on a different database: the first and the third are based on administrative data, while the second one relies on survey data.

The first chapter, "The inequity of the Swiss health care system financing from a federal state perspective" (with Luca Crivelli) focuses on the equity issue in the health care system financing. It aims to detect differences at the cantonal level. In previous studies the Swiss health care financing results as particularly regressive. However, the currently available information does not allow to monitor income-related inequities in financing health care across cantons. This study aims at filling this information gap exploring the differences and the level of regressivity in the financing of the health care system across cantons and over time, considering the role of fiscal federalism. In the first part of the study, I describe the different strategies adopted by the cantons to finance health care, looking at how much financing relies on private resources and how much on public ones. In the second part, I use the Kakwani index as a summary measure of regressivity. I compute it for each canton and for each of the sources that have a role in financing the health care system, i.e. the taxes, both at federal (including VAT) and cantonal level, the net premiums (after subtracting possible subsidies), and other social contributions. I graph concentration curves and perform relative dominance tests, which utilize the full distribution of expenditures. The microdata comes from the "Swiss Household Income and Expenditure Survey" based on a sample of the Swiss population (about 3500 households per year), for the years 1998–2005. The empirical evidence confirms that the Swiss health care system financing has remained quite regressive since the major reform of 1996 and shows that the variations in equity across cantons are quite significant: the difference between the most and the least regressive canton is about the same as between two extremely different financing systems like the US and Sweden. There is no evidence, however, of a clear evolution of regressivity over time. The significant variation in equity across cantons can be explained by fiscal federalism and the related autonomy in the design of tax and social policies. In particular, the results highlight that earmarked subsidies, the policy adopted to smooth the regressivity of the premiums, appear to be not enough.

The second chapter, "Equity in the Swiss health care system: the role of managed care and higher deductible", investigates on a different side of equity: the equity in the access to health care services. The goal of this study is to analyze empirically the impact that alternatives forms of insurance, namely managed care and higher deductible, may have on the equity in health care utilization. I compute concentration indices for the distribution of visits to GP and to the specialist for the years 2007 and 2012. I use the two last waves of the "Swiss Health Survey", a specific survey addressing health issues on a sample of the Swiss population (about 20000 individuals per wave). I compute the need–standardized level of health care for the total number of visits, for the probability to see a doctor and for the conditional number of visits. The results suggest not only that specialist visits are pro–rich distributed, in line with previous literature, but also that the pro–rich inequity worsened over time. By contrast, no pro–rich inequity emerges in GP visits. I also perform a decomposition analysis (on the total probability of a specialist visit), to disentangle the contribution of each source of inequity. The findings suggest that the choice of a higher deductible acts as an entrance barrier for insured, as expected. Nevertheless, being this option more widespread among the better–off, this barrier prevents from going to the specialist more people belonging to the middle–rich income class. This study reveals that a high deductible seems to have two positive effects. First, it rebalances the concentration of visits towards a smaller pro–rich inequity and second, it discourages moral hazard behaviors. On the contrary, the analysis does not find any contribution to inequality due to being enrolled in a managed care plan, neither in 2007, or 2012.

In the third chapter, "Short terms effects of public smoking bans oh health" (with Fabrizio Mazzonna) I perform an empirical application aimed at assessing the causal effect of the public smoking bans on a health outcome, namely the acute myocardial infarction (AMI). Since a reduction in the AMI cases should be quantifiable immediately after the introduction of the smoking ban, focusing on AMI allows to measure the short-run effects of the policy on an objective metrics of health. I use a quasi-experiment, consisting in the fact that cantonal authorities implemented the same policies in different times, as allowed by the Swiss federalism. I exploit time and geographical variation in the implementation of the policy to provide evidence of a causal association between the policy and the health outcome of interest. Data come from the "Medical Statistics of Hospitals" a wide dataset that collects information on every patient for every Swiss hospital from 1999 to 2012, provided by the Swiss Federal Statistical Office. The results suggest that the effect of smoking ban globally reduced the number of AMI of about 7.5–10% percentage points. The effect seems to be stronger for men between 50 and 65 years old, showing a reduction in the number of AMI of about 15%, and for women over 66. Results are robust to alternative model specifications. The advantage of focusing on smoking bans is twofold. On one hand, it allows us to assess a causal relation of smoking (both active and second-hand smoke) on short-run health effects. On the other hand, assessing the efficacy of this policy measure on health outcomes offers an interesting contribution in the debate on the effectiveness of smoking bans on health.

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Chapter 1

The inequity of the Swiss health care system financing from a federal state perspective

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1.1 Introduction

The idea that health-care services should be paid for according to the ability to pay rather than according to the actual use of the health-care system finds its roots in the egalitarian concept of social justice and is generally adopted in Switzerland, as it is in most of the OECD countries. It is common in the economic literature to indicate this idea using the concept of *equity in financing* (Wagstaff *et al.*, 1989). This principle implies not only a form of solidarity between the sick and the healthy, which is implicit in any health insurance system, but also solidarity between rich and poor. In this context, two measures of equity exist. The first is "horizontal equity", which claims *equal treatments for equals*; i.e., that people with the same income have to contribute the same amount of money to the total expenditures. The second measure is "vertical equity", which states that people with different income must contribute appropriate amounts to the total expenditure.

All of the empirical studies on this topic measure the equity of a financing source in terms of progressivity; i.e., the extent to which higher-income people pay more as a proportion of their income than lower- income people.

When the new Health Insurance Act (HIA) came into force in 1996, many things changed in the Swiss health-care system. The main objectives of the act were to guarantee universal coverage, to establish a fair competition among health insurers and to equip the political system with better cost containment tools. Moreover, it also increased the importance of the financing equity aim: one of the objectives of the reform was to provide monetary assistance to low-income households in order to increase equity of financing. In the federal draft bill issued in the year 1991 (and approved some years later after many amendments as HIA) we can read that: *The main priority of the project is undoubtedly the strengthening of solidarity. The current law provides for individual premiums to be paid without taking into account the economic situation of the people insured (p.67).* In order to make this payment affordable for all citizens, the new law replaced the historically grown system of general state transfers to health insurance with a better endowed and more equitable system of allowances targeted to low-income households.

At present the health insurance in Switzerland is mandatory for every citizen, and it is based on community-rated premiums. The confederation jointly with cantons assigns a budget cap for earmarked subsidies targeting low-income households. Cantons have a great autonomy in deciding the distribution of subsidies; despite their role of mitigating vertical inequity, the heterogeneity of cantonal subsidy policies leads to a different treatment of similar households across cantons (horizontal inequity). As subsidies vary with cantons, people with the same income living in different cantons may pay a different amount of money for health care.

Some studies have demonstrated that the general level of health system financing is regressive in Switzerland, both before and after the reform (Bilger, 2008; Wagstaff *et al.*, 1999). This means that lower-income people pay more as a proportion of their income than higher-income people.

Nevertheless, these studies consider only the whole of Switzerland and no research has monitored the situation at the cantonal level. Due to the Swiss federal setting, each canton differs in the economic strategy it has adopted to finance the health care system; this leads to different levels of equity among them. The OECD Review of the Swiss Health System (OECD/WHO, 2011) considers the development of an information system able to monitor this inter-cantonal variation as one of the Swiss policy challenges for the future.

This paper aims to fill this information gap by analyzing the level of equity in the financing of the health-care system in the Swiss cantons. In this study we have used the methodology suggested by Wagstaff and Van Doorslaer (1992) to compute the regressivity level of each canton through the Kakwani index (Kakwani, 1977). The paper is organized as follows. The next section presents a review of the most relevant literature in this field. We have collected and summarized similar studies conducted for other countries, as well as for all of Switzerland. Section 1.3 provides a brief explanation of Swiss health system financing. Section 1.4 introduces the dataset, section 1.5 explains the methodology used and section 1.6 the results obtained. Section 1.7 discusses the results in more detail and outlines the limitations of the study. Finally, section 1.8 offers some conclusions.

1.2 Literature review

This work finds its place among the studies about the equity of health-care system financing.

Table 1.1 includes the main papers that have been published on this topic. Some of them have focused on a single country - namely Australia, Iran, Finland, Italy, the Netherlands, Malaysia, Palestine, Ireland, Tanzania, Ghana, Sweden, and Switzerland. Two papers present an international comparison, based on 13 Asian countries and 13 OECD countries, respectively.

Table 1.1: Main papers about equity in the health-care financing system

Country	**Author(s) and year**
Italy	Paci *et al.* (1993)
Australia	Lairson *et al.* (1995)
The Netherlands	Wagstaff and van Doorslaer (1997)
Finland	Klavus (1998, 2001)
Sweden	Gerdtham and Sundberg (1998)
Malaysia	Yu *et al.* (2008)
Switzerland	Bilger (2008)
Palestine	Abu-Zaineh *et al.* (2008)
Ireland	Smith (2010)
Tanzania	Mtei *et al.* (2012)
Iran	Alireza (2011)
Ghana	Akazili *et al.* (2011)
13 OECD countries	Wagstaff *et al.* (1999)
10 OECD countries	Wagstaff and Van Doorslaer (1992)
13 Asian countries	O'Donnell *et al.* (2008a)

Smith (2010), Klavus (1998), Akazili *et al.* (2011), and Abu-Zaineh *et al.* (2008) analyzed the inequity in financing through the Kakwani index as well as through a "disaggregated approach"; the latter provides summary measures over specific income groups, by using the dominance test to assess the level of progressivity for different deciles of income distribution.

Another different method consists of separating the redistributive effect into three parts: vertical, horizontal, and reranking effect. The works by Lairson *et al.* (1995) for Australia and Gerdtham and Sundberg (1998) for Sweden concentrate on this approach. Bilger (2008) adopted the same methodology in the case of Switzerland.

Many studies have analyzed the equity of health system financing at the national level, while very few studies have explored this issue at the subnational level. Abu-Zaineh *et al.* (2008) and Alireza (2011) extended the analysis by computing the Kakwani index for two different regions of a given country: the former included in the study the two regions of the Occupied Palestinian territory (i.e. the West Bank and the Gaza Strip) and the latter applied the Kakwani index to urban and rural areas of Iran. As far as Switzerland is concerned, the issue of the health-care system financing has been explored by Wagstaff and Van Doorslaer (1992); Wagstaff *et al.* (1999) and Bilger (2008). All three papers provide evidence in favor of highly regressive financing.

Wagstaff and Van Doorslaer (1992) presented an international comparison of health-care financing across 10 countries. The score of the Kakwani index for Switzerland was based on 1982 data and was significantly negative (-0.117). Moreover, the Swiss healthcare financing was the second most regressive, after the U.S. (Kakwani index of -0.145).

In 1999, the same authors, along with some others, updated their previous paper with more recent data. In line with their previous work, the analysis for Switzerland shows that the new Kakwani index, based on the data of 1992, is -0.1402. In ten years the inequity level of financing grew even worse.

Bilger (2008) provided another important contribution to the analysis of financing equity in Switzerland. For his analysis, he used the Swiss Household Income and Expenditure Survey (SHIES) of 1998, which contains data from 9295 Swiss households. He found evidence that health system financing in post-reform Switzerland remains very regressive. In particular, the paper concludes that the reform failed to reduce vertical inequity.

Starting from this strand of literature, the current paper aims to go a step further. The goal is to analyze the differences in the regressivity of the financing system at a sub-national level. Apart from the two previously noted exceptions (focusing on two regions of the same country), this is the first attempt to provide a systematic analysis of financing equity at the level of the single entities of a federal state (in this case, Switzerland at the cantonal level). Switzerland is an ideal country in which to add this new piece of evidence, as it can be set in a context of fiscal federalism, where the cantons have large freedom to decide on the financing sources for health care and in the design of the tax system (Kriesi *et al.*, 2005).

1.3 Swiss health system financing

1.3.1 General framework

The Swiss health care system is based on a mixed private-social health insurance, and it is financed through several sources (in brackets the 2009 shares on the total health care costs): public expenditure (28%), mandatory health insurance (29%), general social insurances (6%), complementary health insurance (9%), co-payments and out-of-pocket (27%), other private (1%) expenses.

Public expenditure

Federal, cantonal and local authorities contribute to the health care system mainly in two forms:

- Direct subsidy of specific services (inpatient hospital care, nursing homes, home care), mostly paid by cantons, and payment of public health interventions (like prevention and health promotion).

- Indirect financing through earmarked subsidies for lower-income people who cannot afford mandatory health insurance, funded jointly by the Confederation and the cantons, and other means-tested allowances for specific target groups.

All these activities are financed through tax revenues: they include both direct and indirect taxes for the Confederation, mostly direct taxes for cantons and municipalities. In particular direct taxes consist of income taxes for the Confederation and both income and property taxes for cantons and municipalities.

Mandatory health insurance (MHI)

The health insurance in Switzerland, mandatory since the reform of 1996, consists of a monthly community rated premium that does not depend on the level of income. However, different premiums are set for two age-classes: lower premiums for children (age <19) and higher for adults. Moreover, insurers are free to grant a discount for students between 19 and 25. Health insurance is offered through a variable number of private nonprofit sickness funds (ranging from 118 in 1998, to 85 in 2005 and to the current number of 66 in 2012). It covers a comprehensive benefit basket fixed at the federal level. Coverage starts once the yearly deductible (minimum of 300 CHF, about 330 US$, in the standard contract) has been reached and includes a 10 percent copayment (up to a maximum amount of 700 CHF per year). People can opt for higher deductibles (up to 2500 CHF) and get in exchange a premium discount. Premiums are set at the regional level by each health insurer. Some cantons are considered as a single premium community, whereas others (including urban and rural areas) are divided in two or three regions (the total number of premium communities declined between 1998 and 2005 from 78 to 43). As a consequence, premiums significantly differ across regions and cantons, reflecting the huge differences in health care expenditures. Significant variation exists also across health insurers (Ortiz, 2011), due in particular to a still ineffective risk adjustment mechanism (Beck *et al.*, 2010). Horizontal equity across geographic communities is not guaranteed at all, since individuals with the same income pay significantly more or less depending on what premium region they are resident in. Moreover, not even horizontal equity within the same premium region can be completely achieved, due to competition in insurance market, switching decisions by consumers, different deductibles and the option to sign managed care contracts.

As mentioned before, for people earning an income under a certain threshold, the Confederation and the cantons jointly fund earmarked subsidies that cover part or the total of the monthly premium (full coverage concerns in particular retired and disabled people receiving means-tested benefits).

General social insurances

This is Bismarckian insurance system designed to cover through in cash benefits other (non-health) risks like longevity, disability and accident. In some circumstances it provides also benefits through in kind health care services. It covers, in particular, spending for rehabilitation in case of disability and the health care costs in case of professional and non-professional accidents of employed persons.

Complementary health insurance

People can choose to pay a risk-rated premium (for individual contracts) or a community-rated premium (for collective contracts) to have a voluntary supplementary private health insurance. It covers some health-care services or inpatient hospital amenities that are not included in the compulsory benefit basket.

Out-of-pocket

The out-of-pocket expenses include the co-participation to costs (namely the deductible and the 10% co-payment) as well as all other health care expenses not covered in the general contract (e.g. nursing homes fees, dental care, OTC drugs, etc.).

Some of the financing strategy rules are decided at the federal level. However, Swiss federalism allows cantons to make their own decision regarding the exact design of the financing policy they want to adopt.

The cantonal scope of action can take several forms. Firstly, a different mixture of the financing sources chosen by each canton[1] determines a different level of regressivity: cantons that rely more on the MHI than on public expenditure are more likely to be regressive than those that choose to finance the system more through general taxation. Secondly, cantonal autonomy due to fiscal federalism allows significant leeway in deciding the subsidies policy for the worse-off (Balthasar *et al.*, 2008; Gilardi and Füglister, 2008; Kägi *et al.*, 2012). Cantons are allowed to make their own decisions regarding the eligibility criteria for receiving benefits, which creates heterogeneity in their distribution and which leads to differences in the level of regressivity among cantons. Thirdly, each canton has significant leeway in the definition of the tax rates for income and property taxes, whereas each municipality decides every year the percentage of the cantonal tax liability to be paid as local taxes.

According to cantons' choices, different equity (or inequity) levels are thus determined.

1.3.2 Evidence of heterogeneity at the cantonal level

Due to data availability problem, in the analysis that follows we are forced to rely only on the first three funding sources in the list, i.e. public financing, MHI and general social insurances. The part of financing that we are considering is called *socialized* health expenditure (SHE), since it reflects collective spending for the universally accessible basket of health care benefits. It accounts for approximately 60-65% of the total financing in Switzerland.

Hence, in our analysis, we will decompose SHE in four sources: *federal taxes* (including direct tax and VAT), *cantonal and municipal taxes* (henceforth we will refer to this second group of tax simply as cantonal tax), *mandatory health insurance* (computed as net premium, i.e. the difference between the privately paid premium and the earmarked subsidies), and *general social insurances*[2].

The first column in Table 1.2 considers the amount of the SHE for each canton and for Switzerland as a whole in absolute per capita values for 2005. The second column shows the incidence of SHE on average disposable income.

The last four columns of Table 1.2 represent the proportion of the expenditure financed by each source. We computed these weights as an average for the considered period (1998-2005).

The first evidence from the table relates to the absolute level of expenditure which differs

[1]E.g. until 2011 cantons had significant leeway in allocating the public financing to private hospitals. This has been changed by the new regulation of hospital financing that started in 2012.

[2]The available data does not provide any information about the direct federal health expenditure given to each canton or the amount of health expenditure from the general social insurance spent in each canton. To compute these two values, we took the total amount of expenditure and simply imputed the same average spending for each Swiss inhabitant, regardless of the canton in which she/he lives. In the expenditure of the confederation and the cantons, we also accounted for the amount of the subsidies distributed.

Table 1.2: Socialized health expenditure – Absolute value (2005) and share of each financing source (average 1998-2005)

	SHE per capita (CHF)	**SHE as a share of average disposable income**	**Sources of financing**			
			Federal taxes	**Cantonal and municipal taxes**	**General social insurances**	**Mandatory health insurance (net premiums)**
Zurich (ZH)	4198	6.66%	10.00%	29.00%	13.00%	48.00%
Bern (BE)	4297	8.52%	11.00%	30.00%	13.00%	46.00%
Lucerne (LU)	3461	7.19%	13.00%	25.00%	16.00%	46.00%
Uri (UR)	3379	6.85%	13.00%	26.00%	16.00%	45.00%
Schwyz (SZ)	3581	6.14%	13.00%	25.00%	15.00%	47.00%
Obwald (OW)	3155	7.40%	14.00%	27.00%	16.00%	43.00%
Nidwald (NW)	3385	6.21%	13.00%	27.00%	16.00%	44.00%
Glarus (GL)	3541	8.01%	12.00%	30.00%	14.00%	44.00%
Zug (ZG)	3885	5.15%	12.00%	30.00%	14.00%	44.00%
Fribourg (FR)	3876	8.42%	12.00%	28.00%	14.00%	46.00%
Solothurn (SO)	4002	7.78%	11.00%	26.00%	14.00%	49.00%
Basel City (BS)	5854	9.31%	8.00%	39.00%	9.00%	44.00%
Basel Land (BL)	4197	7.60%	11.00%	27.00%	13.00%	49.00%
Schaffausen (SH)	4070	9.06%	11.00%	31.00%	13.00%	45.00%
Appenzell O. Rhodes (AR)	3505	6.33%	13.00%	27.00%	16.00%	44.00%
Appenzell I. Rhodes (AI)	2988	6.81%	16.00%	20.00%	19.00%	45.00%
St. Gall (SG)	3465	6.76%	13.00%	24.00%	15.00%	48.00%
Grisons (GR)	3837	7.18%	12.00%	29.00%	14.00%	45.00%
Argovia (AG)	3575	6.60%	13.00%	20.00%	15.00%	52.00%
Thurgovia (TG)	3204	7.10%	14.00%	19.00%	16.00%	51.00%
Ticino (TI)	4852	9.66%	10.00%	29.00%	12.00%	49.00%
Vaud (VD)	4858	9.01%	9.00%	30.00%	11.00%	50.00%
Valais (VS)	3732	8.32%	12.00%	30.00%	14.00%	44.00%
Neuchâtel (NE)	5017	10.44%	9.00%	33.00%	11.00%	47.00%
Geneva (GE)	6578	10.10%	7.00%	43.00%	8.00%	42.00%
Jura (JU)	4381	10.43%	10.00%	30.00%	12.00%	48.00%
Switzerland (CH)	**4243**	**7.85%**	**11.00%**	**29.00%**	**13.00%**	**47.00%**

greatly across cantons. With respect to the Swiss average of 4243 CHF (registered in 2005, the latest year considered in the empirical analysis), canton Geneva and Basel City present the highest value of expenditures, with 6578 and 5854 CHF, respectively. Cantons Appenzell Inner Rhodes and Obwald register the smallest expenditure values (with 2988 CHF and 3155 CHF, respectively). As far as the incidence is concerned, it ranges from 5.15% of disposable income in the rich and low SHE canton of Zug to almost the double in Geneva, Jura and Neuchâtel. Whereas in Geneva it is the high level of SHE that determines the heavy burden for households, in Jura (where SHE is aligned with Swiss average) the high incidence is rather driven by a low average income.

However, the differences do not only concern the total amount of expenditure: another evident difference between cantons is the proportion of each of the single financing sources.

The table also shows that, in cantons where the total expenditure is higher, the share of the cantonal part tends to be also higher. This is because, where health care costs (and consequently premiums) are higher, many households cannot afford anymore the cost of mandatory health insurance, which means that cantons are forced to spend more in terms of subsidies for the worse-off.

The values in the table show that the most important part is financed by health insurers. This accounts, on average, for 47 percent of the SHE. The second source derives from the cantonal taxes that cover 29 percent of the SHE, followed by the social insurances and federal taxes that account for 13 percent and 11 percent, respectively.

Finally, it is worth noting the magnitude of the differences among cantons. There are two cantons for which the part financed by cantonal taxes is considerably larger (Geneva - 43 percent; Basel City - 39 percent), whereas in Appenzell Inner Rhodes the share financed by cantonal taxes corresponds to only 20 percent.

1.4 Dataset

The micro dataset is the Swiss Household Income and Expenditure Survey (SHIES), which allows computation of household income at a very detailed level, including all the taxes and social contributions paid, as well as the health expenditures (premiums and, only to some extent, out-of-pocket expenditures) and indemnities related to health (subsidies from the state and reimbursements from the insurance companies).

This survey is available from 1998 to 2005 (excluding 1999) and is based on a sample of the Swiss population (approximately 3500 households observed for each of the years between 2000 and 2005, and 9295 observations for 1998).

The SHIES does not guarantee that the households sample is representative at the level of all cantons for each year, but only at the level of seven macro-regions, each of which groups cantons together according to their geographical position[3]. As the focus of this work is to control for the differences due to the federal setting, we were interested in maintaining the cantonal dimension. For this reason, we decided to merge the dataset of each year to obtain a representative sample at the cantonal level. This has been possible because different people were interviewed each year for each canton. In this way we were able to control for the cantonal differences, but we lost the information about the variation over time for each canton.

Ultimately, this study concentrates on the between-variation for each canton and on the within-variation only for the whole Switzerland and the seven macro-regions. There are some differences in the questionnaires, especially for the year 1998 with respect to the other years. Only in the wave of interviews taken in 1998, were people asked to report all the medical expenses (greater than 150 CHF) they had during the whole year preceding the month of the survey and not only all the medical expenses they had during the month of the interview (as it is in the other waves of the survey). Moreover, only for the year 1998 responders had to register the reimbursements received by the insurance companies during the whole year; hence there was a correspondence between the expenditures and their reimbursements. The collection of this data makes it possible to compute the out-of-pocket expenditures for each households and, consequently, to have a robust proxy for the yearly out-of-pocket financing at least for the year 1998.

1.5 Methods: Kakwani index, concentration curves and dominance test

In order to measure the regressivity of each financing source, we used the procedure proposed by Wagstaff *et al.* (1999), (see also O'Donnell *et al.* (2008b) for the empirical implementation). We started by computing the Kakwani index for each financing source and then we computed a total Kakwani index aggregating these results with a weighted average, using as weights the proportion of each financing source with respect to the total financing. This procedure has been applied to 23 cantons[4], to the seven macro-regions, and to Switzerland as a whole.

This index is computed as the difference between the concentration index of each financing

[3] Macro-region 1 includes Vaud, Valais, and Geneva; macro-region 2 includes Bern, Fribourg, Solothurn, Neuchâtel, and Jura; macro-region 3 includes Basel City, Basel Land, and Argovia; macro-region 4 includes Zurich; macro-region 5 includes Glarus, Schaffhausen, Appenzell O. Rh., Appenzell I.. Rh., St. Gall, Grisons, and Thurgovia; macro-region 6 includes Lucerne, Uri, Schwyz, Nidwald, Obwald, and Zug; macro-region 7 includes Ticino.

[4] Canton Uri, canton Appenzell I. Rhodes, and canton Appenzell O. Rhodes have not been considered because the small number of observations in the dataset did not allow for any robust computation.

source and the *Gini* index calculated on the (equivalent disposable) household income before considering any health-related expenditures (we refer to this as pre-health income). If the two curves coincide, the difference between the CI and the *Gini* index is zero and the financing source can be considered as proportional to income.

The formula is as follows:

$$K_{ip} = CI_{ip} - Gini\_pre_i \qquad Range : [-2 : +1] \tag{1.1}$$

where i indicates the geographical unit of analysis and p the different financing sources available in the dataset.

CI_{ip} is the concentration index of each financing source p. CI is twice the area between the concentration curve of the source p and the 45-degree line. It indicates whether the variable of interest is more concentrated among the poor (the concentration curve lies above the equality line and the index has a negative value) or among the rich (the curve lies below the equality line and the index has a positive value).

$Gini\_pre_i$ is the *Gini* index for the *pre-health* income.

Since the Swiss fiscal system does not rely on earmarked taxes for health care, we simply imputed to the amount of taxes paid by each household the share of total health expenditure financed by that tax. We did this for federal tax, cantonal tax, and also for general social insurances.

$Gini\_pre_i$ is computed according to the following formula:

$$y\_pre_h = \sum_{k=1}^{N} y_{hk} - \sum_{s=1}^{m}(1-\mu_s)sc_{hs} - \sum_{v=1}^{p}(1-\tau_v)t_{hv} \tag{1.2}$$

where:

- $\sum_{k=1}^{N} y_{hk}$ is the gross income for the household h, which, according to the definition given in Bilger (2008), consists of the sum of all income earned from work and self-employment, interest, house rental, social insurance benefits, revenues from other insurances, and other indemnities.

- $\sum_{s=1}^{m}(1-\mu_s)sc_{hs}$ is the part of the general social insurance paid by the household h and not directed towards financing health care. In particular, μ_s is the share of health care services funded out of the budget of the social insurances. In 2005, μ_s was 3.3 percent for the pension and the disability insurance, and 24.5 percent for the accident insurance.

- $\sum_{v=1}^{p}(1-\tau_v)t_{hv}$ are the taxes (federal, cantonal and municipal) paid by the household h and not directed towards financing the healthcare system. As in Bilger (2008), federal indirect taxes were proxied with VAT, which accounts for two-thirds of total indirect taxes.

VAT has been computed from the data on consumptions that household declared in the survey.

As in the case of social contribution, τ_v is the part used to finance the health care system for each type of tax (v), so $1 - \tau_v$ is the residual part that is not used within the health sector.

To give an idea of the dimension of these coefficients, in 2005 τ_v was 5.4 percent for federal tax, 28.5 percent (on average) for cantonal tax and 4.3 percent (on average) for communal tax.

Disposable income before health-care financing has been corrected using an equivalence scale in order to make households with a different number of members comparable[5].

Moreover, as different years are pooled together, the values have been deflated through the consumer price index (CPI) for Switzerland.

Kakwani is a useful index for providing information about the shift from proportionality. However, it is an extremely summary measure, and its information can be sometimes misleading if the distributions underlying the Kakwani index are not considered alike. When curves cross, a value of Kakwani equal to zero could be the result of a concentration curve that is progressive for half of the population and regressive for the other half. To control for this, we look also at the concentration curves and the relative dominance test, that aims to define statistically whether one curve dominates another. The null hypothesis indicates that there is no significant difference between the two concentration curves considered. The choice criterion used here is the multiple comparison approach. In the comparison between the two curves, the null is rejected if there is at least one significant difference between them in one direction and no significant difference in the other direction. The ordinates are compared in 19 different quintiles, as suggested by O'Donnell *et al.* (2008b)

1.6 Results

Table 1.3 presents the results of the Kakwani index over time for the whole Switzerland and for the seven macro-regions.

The most notable result is that the total value of the Kakwani index is always negative, which means that the Swiss health-care system financing remains regressive, even after the major reform of 1996; this result is in line with the previous literature. The other notable point is that the results do not seem to vary widely in the years considered, neither for all of Switzerland nor for any of the macro-regions. The most regressive year appears to be 1998, although the data provides no evidence of any clear temporal trend.

[5]The scale chosen is the OECD modified equivalence scale, which assigns a value of 1 to the household head, of 0.5 to each additional adult member and of 0.3 to each child no older than 13.

Table 1.3: Kakwani index over time

	1998	**2000**	**2001**	**2002**	**2003**	**2004**	**2005**
Macro-region 1	-0.136**	-0.091**	-0.069**	-0.065**	-0.099**	-0.090**	-0.074**
	[1495]	[590]	[586]	[603]	[528]	[505]	[502]
Macro-region 2	-0.106**	-0.081**	-0.081**	-0.109**	-0.105**	-0.085**	-0.109**
	[2173]	[925]	[907]	[840]	[788]	[780]	[751]
Macro-region 3	-0.103**	-0.099**	-0.044*	-0.062**	-0.065**	-0.076**	-0.084**
	[1287]	[504]	[516]	[529]	[482]	[441]	[400]
Macro-region 4	-0.130**	-0.138**	-0.092**	-0.094**	-0.104**	-0.111**	-0.112**
	[1659]	[638]	[612]	[640]	[622]	[562]	[531]
Macro-region 5	-0.130**	-0.116**	-0.125**	-0.119**	-0.084**	-0.137**	-0.147**
	[1183]	[433]	[478]	[469]	[438]	[420]	[417]
Macro-region 6	-0.136**	-0.135**	-0.080*	-0.089**	-0.108**	-0.125**	-0.075**
	[806]	[329]	[300]	[313]	[289]	[287]	[245]
Macro-region 7	-0.087**	-0.074**	-0.078**	-0.034	-0.027**	-0.083**	-0.016**
	[692]	[223]	[341]	[332]	[328]	[275]	[241]
CH	**-0.130****	**-0.110****	**-0.084****	**-0.094****	**-0.098****	**-0.104****	**-0.101***
	[9295]	**[3642]**	**[3740]**	**[3726]**	**[3475]**	**[3270]**	**[3087]**

Significance level: $^{*}p < 0.05$, $^{**}p < 0.01$.

The values in brackets show the sample size.

Macro-region 7 (Ticino) seems to reach the best value in terms of equity, for all the years considered, since the Kakwani is not statistically different from zero in 2002, 2003 and 2005. Macro-region 4 (Zurich) is in line with the Swiss average, and macro-region 3 (Basel City, Basel Land, and Argovia) is, after Ticino, the least regressive group of cantons.

The weighted Kakwani index for the SHE in Switzerland ranged from -0.084, reached in 2001, and -0.13 in 1998.[6]

Table 1.4 presents the Kakwani indexes for the Swiss cantons and helps shed light on how fiscal federalism allows cantons to choose their preferred financing policy.

It is evident from the table that there are important differences in the regressivity of the financing system among cantons.

The Kakwani index for *federal taxes* is not statistically different from zero for all the cantons apart from canton Basel Land (BL), which has a slightly progressive value. This means that federal taxes are concentrated quite proportionally among the population. This result could be easily explained if we think that the tax amount considered here is the sum of a progressive federal direct tax (based on income) and an indirect tax (VAT) based on consumption that is normally regressive. Canton Basel Land is the only one in which the effect of the progressive direct tax more than offsets the effect of the regressive VAT.

The Kakwani indexes for the *cantonal and municipal taxes* are more difficult to explain. While we would expect a highly progressive value, most of them are not statistically different from zero. Only Solothurn (SO), Ticino (TI), and Basel Land (BL) present a slightly positive value of the Kakwani index, which means a slightly progressive tax. These values are smaller than the results expected, but there are at least two reasons that could explain them.

First of all, this financing source includes both cantonal and municipal tax. Very rich households may choose to live in jurisdictions that have a lower local (municipal) taxation. Therefore, a richer household may have to pay the same cantonal tax, but a different (lower) municipal tax than a poorer household that lives in a municipality with higher taxation. This could bias the progressivity results in favor of a score that cannot reject the hypothesis of proportionality.

A second reason could be the fact that there is a mismatching between household income that refers to the year of the interview and household taxes that are paid on income earned the year prior to the interview. There could be situations in which a person has reduced his or her revenues (perhaps because they have temporarily lost their job or have retired), but reports to pay high taxes because she/he earned a "regular" income in the previous year. This could worsen the situation of regressivity because the Kakwani index is computed as if tax paid and income

[6] As explained in section 4, data for 1998 was collected in a different manner that makes it possible to compute the Kakwani index for the total health expenditure, including the out-of-pocket payment and the complementary insurance. We computed this for all of Switzerland and found that the equity results are even worse (Kakwani -0.14).

Table 1.4: Kakwani index across cantons

	Number of observations	**Federal Taxes**	**Cantonal and municipal taxes**	**General social insurances**	**Mandatory health insurance (net premiums)**	**Weighted Kakwani index**
ZH	5264	0.0088	0.0389	0.0074	-0.2548**	-0.1085**
BE	4191	-0.0052	-0.0121	0.0330**	-0.20806**	-0.0947**
LU	1353	-0.0069	0.0181	0.0560**	-0.2298**	-0.0936**
SZ	450	-0.0578	-0.0417	0.0025	-0.2508**	-0.1344**
OW	129	0.0387	0.0744	0.0999**	-0.2323**	-0.0567**
NW	158	0.0236	0.1256	0.0594	-0.2738**	-0.0729**
GL	138	0.0063	0.0135	0.0366	-0.2304**	-0.0913**
ZG	375	0.1093	0.0540	-0.0230	-0.2603**	-0.0871**
FR	916	-0.0159	-0.0121	0.0246	-0.1863**	-0.0874**
SO	1013	-0.0311	0.0473**	0.0536**	-0.2109**	-0.0868**
BS	782	0.0485	0.0835	-0.0001	-0.2149**	-0.0582**
BL	1066	0.0557*	0.0933**	-0.0042	-0.2191**	-0.0774**
SH	295	-0.0963	-0.1310	-0.0059	-0.2239**	-0.1535**
SG	1648	0.0011	0.0407	0.0369*	-0.2571**	-0.1066**
GR	723	-0.0481	-0.0327	0.0231	-0.2605**	-0.1285**
AG	2311	0.0200	0.0308	0.0147	-0.2250**	-0.1065**
TG	806	0.0251	-0.0372	0.0406*	-0.2281**	-0.1131**
TI	2432	0.0122	0.0850**	0.0310*	-0.1800**	-0.0592**
VD	2531	-0.0148	0.0131	-0.0099	-0.2050**	-0.1004**
VS	1083	-0.0464	-0.0435	-0.0006	-0.2123**	-0.1114**
NE	776	-0.0526	0.0306	-0.0035	-0.2009**	-0.0904**
GE	1195	0.0277	0.0976	-0.0673**	-0.2482**	-0.0643*
JU	268	-0.0076	0.0799	0.0549	-0.1809**	-0.0566**
CH	**30232**	**0.0074**	**0.0153**	**0.0090***	**-0.2271****	**-0.0999****

Significance level: $*p < 0.05$, $**p < 0.01$.

earned referred always to the same year[7].

As expected, the Kakwani indexes for the general social insurance are not statistically different from zero. Since the tax rate is the same for every wage earned, regardless of its level, we expected a Kakwani value near to zero. Nevertheless, there are some exceptions. Four cantons (Bern, Lucerne, Obwald, and Solothurn) present a progressive value, while canton Geneva is the only one that has a regressive index for this financing source. This could be because the share of capital income over the total income in Geneva is higher than in the other cantons (and, consequently the wage income share is lower): the fact that the payment for the social insurance is only based on wage income could induce a regressive effect in the results.

The most interesting results are those for the *mandatory health insurance*. They are all significant for each canton and range from a value of -0.18 for canton Ticino to -0.27 for canton Nidwald.

It is quite notable that the values can be so different, even in neighboring cantons: the value of the index in canton Grisons (-0.26) is 0.08 less than that of its neighbor, Ticino (-0.18). Based on the Welch's t-test, we can show that this difference is statistically significant at the 95% level. This outcome probably hides a different choice in the subsidy policy that cantonal authorities are free to set-up. Previous studies (see Balthasar *et al.* (2008); Ortiz (2011)) have shown that on average the same family (e.g. parents with two children and a revenue of 70'000 CHF) has to devote 4.4% of disposable income if it lives in Obwald, but up to 16.4% if the canton of residence is Vaud. This heterogeneity jointly reflects the premium level and the generosity of the subsidy policy. In general cantons with a very high level of premiums prefer to distribute generous subsidies to a limited group of citizens, whereas those with lower premiums levels choose to reach more people by distributing lower subsidy amounts (OECD/WHO, 2011). There is one clear outlier-canton which distributes higher subsidies than the average to an above-average share of the population, namely Ticino (Crivelli *et al.* (2007) show that the subsidy policy in Ticino is 34.5 percent more generous in terms of CHF spent per capita than the Swiss average). This might explain why in Ticino the net premiums are less regressive than elsewhere. On the other hand Grisons distributes subsidies that are below the Swiss average and support only a limited number of people. This strategy seems to be in line with the computed Kakwani for MHI, which is in the case of Grisons the most regressive one.

Finally, the results on the *total* Kakwani indices are all statistically significant[8] and all

[7]There is another factor to consider: the tax reform. From 1990 until 2003 cantons, in turn, had to adapt their system of taxation to a new one. This passage provoked an asymmetry between the tax paid in one year and the income earned in the same year that may have brought some additional asymmetries also in our computations.

[8]The standard error for the total Kakwani index have been obtained summing the standard errors for the single components (i) according to the formula:
$se = \sqrt{\sum_i (w_i se_i^2) + 2\sum_{ik}(w_i w_j cov_{ij})} \quad \forall i \neq j$

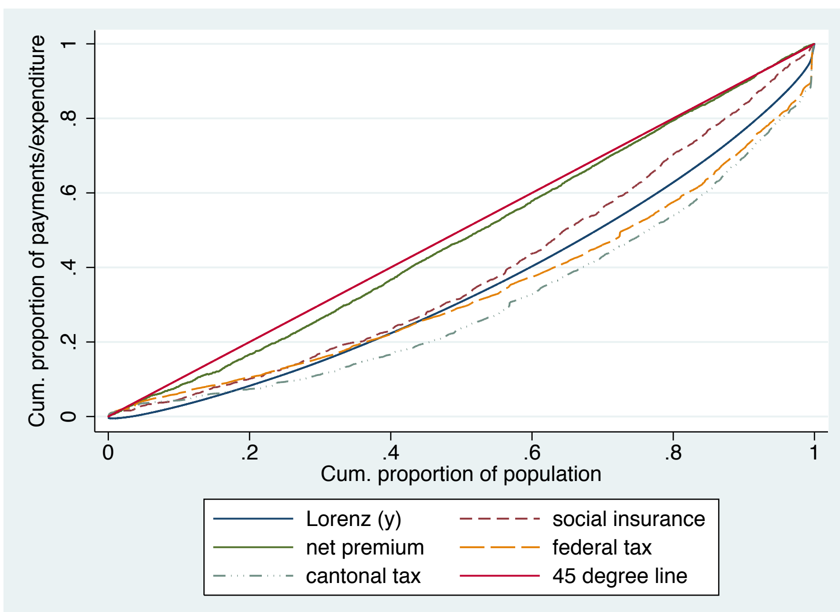
negative.

The cantons that appear to be the most regressive ones are canton Schaffhausen and Schwyz, with a Kakwani equal to -0.15 and -0.13, respectively. The cantons that reach the best results in terms of equity are Jura, Obwald and Ticino, all of which have a Kakwani value of around -0.05.

It is worth noting that the difference between the least (Jura) and the most regressive canton (Schaffhausen) is considerable (approximately 0.10) and significant at 99% level (according to the Welch's t-test); it is the same as the difference between the US (-0.13) and Sweden (-0.015) from the Wagstaff et al.'s (1999) study, countries that have two completely different health-care financing systems. Although Shaffausen and the Jura have greater absolute levels of regressivity than the US and Sweden respectively.

We now report the concentration curves and the dominance tests results.

Figure 1.1: Concentration curves for each financing source, canton Geneva.



A general summary comment, based only on the visual inspection of the concentration curves, would suggest that their order follows a similar pattern for all the cantons. Firstly, starting from the 45-degree line and going down, the net premium curve is always the first one, very close to the 45-degree line and not crossing with the others. Secondly, the social contribution curve is generally the next one, always very near to the Lorenz curve, which indicates proportionality. Thirdly, the federal tax line is also quite near to the Lorenz curve and, finally, the cantonal tax curve is always the most distant and lies under the Lorenz, indicating progressivity. Nevertheless, the two curves for taxes and the one relative to the social insurance often cross among them and with the Lorenz, which indicates an alternation between regressivity among the poorer and progressivity among the richer.

We have chosen to show the results only for canton Geneva, which is among the least regressive.

Figure 1.1 represents the concentration curves of each financing source. The concentration curve for the net premium is very close to the 45-degree line, indicating quite strong regressivity, and it dominates the Lorenz curve. In Table 1.5 we read that the poorest 30 percent of the population receives about the 15 percent of the total income and pays 26 percent of the net premium.

Table 1.5: Cumulative shares of income and health payments by income decile, canton Geneva

Quantile	**Cum. share of income**	**Cum. share of federal taxes**	**Cum. share of cantonal and municipal taxes**	**Cum. share of general social insurances**	**Cum. share of mandatory health insurance (net premiums)**
Q10	2.78%	6.26%**	4.32%	4.57%	8.08%**
Q20	8.24%	10.55%*	7.40%	10.06%	16.62%**
Q30	14.75%	15.61%	11.21%*	16.87%	26.16%**
Q40	22.35%	22.04%	16.60%**	23.15%	36.71%**
Q50	30.82%	29.29%	23.69%**	31.75%	47.10%**
Q60	40.31%	37.41%	32.80%**	43.70%*	57.73%**
Q70	50.92%	46.08%*	42.95%**	55.78%**	68.60%**
Q80	62.85%	57.53%*	53.94%**	70.27%**	79.40%**
Q90	76.98%	71.94%	69.61%*	83.72%**	89.47%**

Significance level: *$p < 0.05$, **$p < 0.01$. The significance level indicates whether the percentage share of financing source is significantly different from the *income_pre*.

The tests of dominance (not shown in the table) find evidence of dominance of the social contributions curve on the Lorenz curve and dominance of the Lorenz curve on the cantonal tax curve. These two last results tell us two important things. Firstly, the dominance for the social contribution curve confirms the result of the Kakwani index: the social contributions are regressive in Geneva, while the situation is different for the cantonal tax. This financing source seems to be progressive, but only starting from the second decile of the ranked population (i.e., for the richest 80 percent). For the poorest 20 percent of the ranked population, we cannot reject the hypothesis of proportionality (as the difference between the two curves is not statistically significant).

1.7 Discussion: lessons for a federal state

The most relevant evidence is that the regressivity of the MHI implies that the subsidy policy adopted by the state does not succeed in making the financing of this source progressive or, at least, proportional. Considering that the subsidy is the most important tool that cantons have to mitigate the regressive nature of the community-rated premiums, this is a very strong result for policy makers.

However, the value of Kakwani here includes something more than the effects of the earmarked subsidies chosen by the cantons. Also, the individual choices (premium and deductible) allowed by the competitive setting of the Swiss health insurance have a significant impact on equity. The Swiss system encourages competition between insurance companies, which plays out in the level of the premium and in the quality of service; twice per year, people have the option to switch to another insurance company (with lower premiums, e.g.). Moreover, citizens are allowed to obtain a discount on the monthly premiums by choosing a higher deductible; by assuming some of the financial consequences of getting sick themselves, they can pay a smaller monthly premium[9]. Finally, people who choose to be part of a managed care plan can also get a discount on premiums. These choice options are not constrained to a certain level of income, but there is evidence that not all citizens can manage the available information to make a rational choice in terms of switching (Ortiz, 2011) or of selecting the appropriate type of contracts. Accordingly, we cannot exclude the existence of a social gradient in the ability to manage information. In this case, the burden of solidarity with the sick implied by community rating may be partially shifted back from the good risk to the bad risk as well as from the better off to the worse off, with negative consequences on the equity level.

Similarly, the Kakwani index of taxation includes also an indirect (or endogenous) effect due to individual choices. For example, very rich people may decide to live in a canton (or municipality) with a less progressive taxation to pay less taxes and this may also have an impact on the regressivity level of the canton. If this effect increases or decreases the regressivity depends on the distribution of income among the population.

Another consideration needs to be specified. The Kakwani index considers only the relative financing burden of households with respect to others, not the absolute burden of the health financing. The poor in a canton with a slightly regressive financing may have to pay absolutely more for health care than people with the same income in a canton with highly regressive financing.

In other words, the Kakwani index is appropriate to measure the level of vertical inequity,

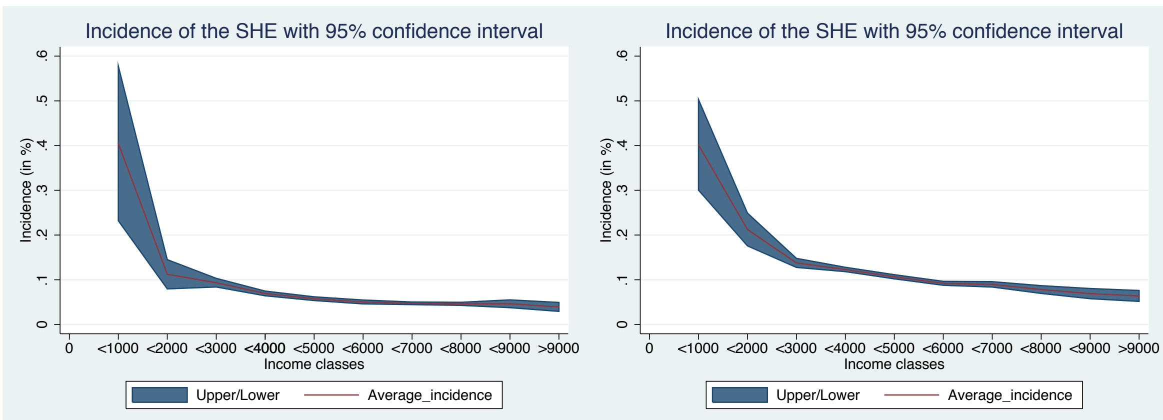
[9]Financially constrained households may 'underinsure' by choosing higher deductibles or switching to a more restrictive insurance plan (managed care). Hence they may incur higher out of pocket payments, which are not measured in the study and which would underestimate the inequity level.

but it is not able to capture the extent of horizontal inequity.

To have an idea of the dimension of horizontal inequity, we computed the incidence of health expenditure on the total income for different income classes. We found that there is a similar general trend for all cantons; specifically, the incidence for the lower-income classes is heavy and decreases for the richer classes.

To exemplify this, the graphic representations for the incidence of the financing burden for the cantons of Zug and Geneva are reported below (Figure 1.2).

Figure 1.2: Average incidence of the socialized health expenditure on income, Zug (left) and Geneva (right).



The red line represents the average incidence for each of the income classes considered, while the blue-shadow area indicates the confidence intervals. Interestingly, the two red lines have a similar shape, although the line for Geneva is shifted upwards with respect to the other.

With an exception for the first income class, the level of incidence is quite different: the share of income that people in Geneva pay is almost double that of citizens earning an equivalent income in Zug. This means that people in Geneva (one of the least inequitable cantons) pay more, not only in absolute value, but also as a percentage on their income than in Zug (one of the most inequitable cantons in terms of the Kakwani analysis).

The reason for such a result can easily be found by simply looking again at the absolute level of expenditure (Table 1.2): Geneva registers one of the highest levels (6578 CHF per capita), while Zug has one of the lowest ones (3885 CHF per capita).

This example clearly shows that the burden of health-care financing may be equally distributed, but may be too heavy to be affordable for lower- and middle-income people. To have a clear understanding of the equity issue, the importance of the absolute expenditure levels cannot be neglected.

The fact that people of the same income class show a different incidence level across cantons suggests the presence of horizontal inequity. We would like to highlight a final remark upon this issue. The Swiss federalism allows some form of autonomy to cantons, especially in deciding the subsidy policy. Therefore, subsidies are an important tool in reducing the vertical inequity within a canton, but, on the other hand, being administered differently in each canton, they may also lead to further horizontal inequity, i.e. to more inequity across cantons.

Finally, we cannot disregard that one possible criticism with respect to the equity issue is the potential trade-off between equity and efficiency, according to what Okun theorized in 1975 (Okun, 1975). To check for this, we considered the efficiency results found in a study by Widmer and Zweifel (2012) in which they performed a data envelopment analysis for the Swiss cantons for six categories of public goods, including health, using data relative to the years 2000-2004. We ranked the cantons first according to our equity results and then according to the efficiency results found by the two authors. We compared the two ranking lists through a Spearman test and a Kendall test of rank correlation, but the results do not reject the null hypothesis of independence. This leads to the conclusion that there is no direct link between equity and efficiency in this particular setting.

1.8 Conclusions

This study represents the first attempt to investigate the impact of federalism on the financing of a universal health insurance system in terms of equity.

We have analyzed the financing of the Swiss health-care system from an equity point of view, with particular attention to the differences across cantons. We used the Kakwani index for each financing source to see how each of them shifts from proportionality. The dominance test for the concentration curves has also been performed to exploit more the available information.

The general results suggest that Swiss health-care system financing is regressive in all cantons, although there are huge differences among them. The reason for this lies in the federal setting, which allows cantons to have some freedom in certain areas, such as how to design subsidy policies for lower-income people, the choice of tax rates, and the choice of how much of the total expenditure has to be financed through taxation and how much through mandatory health insurance. These factors, along with the characteristics of the Swiss health-care system based on competition among insurance companies and on the supply of premium discounts for people opting for higher deductibles and managed care contracts, bring different levels of inequity. The results highlight the fact that the level of regressivity of mandatory health insurance premiums (net of subsidies) varies significantly across cantons, whereas for federal taxes and general social insurance the hypothesis of proportionality cannot be rejected in the majority of cantons. Finally, the cantonal tax is progressive only in three cantons.

The main limitations of our paper are due to the dataset on which we relied; we had to aggregate data of seven different years to get robust estimates at the cantonal level. Moreover, we had to focus on SHE because the information on out-of-pocket expenditure was not reliable for six out of seven years. Since 2008, a new legislation has been enforced in Switzerland that assigns to cantonal authorities even more leeway regarding the subsidy policy (the matching grants transfer has been replaced by a lump-sum payment). It is therefore necessary to monitor the situation in more recent years in order to figure out how inequity developed in Swiss cantons after this reform.

The results contained in this work shed light on some aspects that should be considered in other federal states that are planning to use the regulation of private health insurance and premium subsidies to ensure universal coverage. Although the combination of community rating and premium subsidies might be in theory the best solution in terms of (vertical) equity (Kifmann and Roeder, 2011), in the daily practice of federal states the empirical evidence can point to a different outcome. In fact, tax competition and a "race to the bottom" in social policy (in order to avoid the outflow of good taxpayers and the inflow of people looking for social aid) may jointly lead to a lower level of vertical equity than a sovereign state would choose if the design of social policy would occur at the national level.[10]

[10]In 1991, through the HIA draft bill, the Swiss federal government suggested to fix the maximum incidence level for health insurance premiums at 8% of the taxable income. However, this 'social target' was not anchored in the law and cantons were entrusted with large autonomy in the design of their subsidy scheme. In actuality, in several cantons a large part of the population pays for premiums a share that significantly exceeds this threshold (Balthasar *et al.*, 2008; Ortiz, 2011).

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Chapter 2

Equity in the Swiss health care system: the role of managed care and higher deductible

Keywords
Equity, inequality, managed care, doctor visits.

2.1 Introduction

Switzerland, as many other OECD countries, has adopted an egalitarian view regarding the organization and the delivery of health care services. This means that the health care system is considered equitable in the utilization if people with the same health care needs are treated equally, irrespective of their socio–economic conditions. In line with the literature that investigates this topic, we refer to the concept of *horizontal inequity* any time the differences in the health care utilization are amenable to a different socio–economic condition. In this paper, we focus on the equity in utilization taking into account the recent Swiss health care system changes.

Equity in access to health care has not emerged as a big issue in Switzerland in the past years. Leu and Schellhorn (2004) analyze equity in access to health care in Switzerland over time. Their analysis is based on data from the "Socio Medical Indicators for the Population of Switzerland" for 1982 and on the "Swiss Health Survey" for the years 1992, 1997 and 2002. They assess the level of horizontal equity for each year and report separate estimates for visits to the general practitioner (GP) and to specialists. Results suggest that there is little or no inequity in GP consultations, whereas specialist visits are pro rich distributed. The authors find little or no variation over time.

Van Doorslaer *et al.* (2006) perform a similar analysis looking at inequality in the use of medical care in 21 OECD countries in 2000. They compute inequity in doctor utilization through the concentration index of the need–standardized use. Results for Switzerland suggest that specialist visits are slightly distributed in favor of rich people, while there is no evidence of horizontal inequity for GP visits.

Allin *et al.* (2009) compare the socioeconomic inequality in the use of health care services based on wealth and income, separately. The ultimate goal of the paper is to investigate the sensitivity of the measures of equity to the different socioeconomic indicators. Data comes from the Health and Retirement Study in the US and the Survey of Health Ageing, and Retirement (SHARE) in Europe, for the year 2004. Regarding Switzerland, the authors find that the probability of visiting a doctor is slightly pro–rich distributed, but only conditional on income, not on wealth. Nevertheless, the authors specify that results on Switzerland are not robust due to the small sample size.

We can assume from these studies that the inequity in health care utilization in Switzerland is not a source of big concern. Pro-rich inequity in the distribution of specialist visit is actually an issue, but it is compensated by the slightly pro–poor distribution of GP visits. Compared to the other European countries, on average Switzerland ranks in mid positions for the inequity in doctor visits.

Nevertheless, there are no recent studies able to attest that inequity in health care utilization has not become an issue after the recent evolutions of the health care system. The analysis of the equity in the access to medical services for the years 2007 and 2012 is of particular interest if considered in relation to the main challenges that the Swiss system has been facing in these years. The surge in health care costs and the consequent increase in insurance premiums, pushed a large share of population to adhere to new forms of contract that allow them to benefit from a substantial discount on the monthly paid premiums of statutory health insurance. These alternative forms of insurance are either contracts based on gatekeeping (such as managed care contracts), or contracts with a higher deductible. Given that these contractual features are generally adopted by healthier people (who are generally also the better–off), their larger diffusion may have translated in a worsening of the inequity of the financing system (as already discussed in chapter 1). On the other hand, we are not able to predict in advance their effect on the inequity in access to care. If also individuals belonging to the first income quintiles have enrolled in these plans, we cannot exclude that, on average, the inequity in the access to care worsened. These contracts may have introduced some forms of barriers that somehow affected the access to healthcare providers.

The goal of this paper is to explore the dimension of equality and equity in access to the Swiss health care system in 2007 and 2012, trying to assess whether there is a trend that may be traced back to the major changes of the Swiss health care system in the period considered.

The analysis is focused both on federal level and on linguistic areas. Moreover, with respect to the previous literature, we rely on a better empirical definition of needs. We use a large set of variables to capture the need for health care, and to compute the need–standardized number of visits.

Furthermore, it is worth highlighting that the time lapse in which our analysis is based (2007–2012) includes the delicate years of the global financial crisis (2008–2009). Although in Switzerland we do not find either evidence of significant cuts in the public sector due to the crisis, or shortages of health care supplies, a report by the WHO lists Switzerland among the countries that report a reduced use of health care due to financial barriers since 2008 (WHO, 2013). In a similar situation, it might occur that people who can rely on less economic resources are those who choose to reduce the number of doctor visits, leading to an increased inequity of health care system utilization. This leaves room for deeper research in this field.

In our analysis, we compute concentration indices for the distribution of visits to GPs and visits to specialists for the years 2007 and 2012. We analyze the total number of visits, and we also perform a two–step approach to separate the probability to consult a doctor from the conditional number of visits. Our results suggest not only that specialist visits are pro-rich distributed, in line with previous literature, but also that the pro–rich inequity worsened over time. By contrast, no pro–rich inequity emerges in GP visits. If compared to the other European countries, Switzerland's inequity in access to care is one of the highest. It is close to the values of countries such as Portugal or Italy, which rank in the second half of the list in terms of equity in specialist visits (van Doorslaer *et al.*, 2004).[1]

We also perform a decomposition analysis (on the probability of consulting a specialist) to disentangle the contribution of the different sources of inequity. We show that some forms of managed care contracts, and even more the choice for a higher deductible, are negatively correlated with the probability of a specialist visit. However, since they are generally chosen by richer and healthier people, their contribution ends up to be towards a more equitable distribution of care. Being enrolled in a plan with higher deductible seems to reduce inequity in access to specialist care. On the contrary, the analysis does not find any contribution to inequality due to being enrolled in a managed care plan, neither in 2007, or 2012. We have to highlight that our results are specific to the Swiss context, where the financing system assigns public subsidies to the worse–off. Therefore, people who receive financial aid have no interest in saving money on monthly premiums, and they generally opt for the basic insurance contract.

The paper is structured as follows. The next section will explain the major trends of the Swiss health care system in the years 2007–2012. Section 2.3 illustrates the empirical methods

[1] Note that we are comparing results of different years. Results for Switzerland refer to 2007 and 2012, while results for the other countries precede 2004.

we use. Section 2.4 defines the dataset and the variables we employed in the analysis. Section 2.5 shows and discusses the results of the study. Finally, the last section concludes.

2.2 Recent trends in the Swiss health care system

After the reform of 1996 with the enactment of the Health Insurance Law, all Swiss residents have to be covered by a statutory health insurance. Each insurer offers the same nationally defined benefit basket, and competes on the level of premiums. Premiums may differ according to the geographic area, three age categories (children, young adults between 18 and 25 and adults) and insurance company. Premiums do not depend on income: a system of earmarked subsidies, financed both by the Confederation and the cantons, provide financial aid for the worse–off. Insurance plans may vary by deductible level; the higher the deductible, the higher is the discount offered with respect to the basic premium. Furthermore, different forms of managed care contracts are also available and allow for a discount on premiums in exchange for a reduced freedom of choice[2].

In the next subsections we discuss two recent trends: the increased popularity of managed care plans and the further diffusion of higher deductibles.

Larger diffusion of managed care plans

Managed care contracts offer a discount on premium to the insured at the cost of a reduced freedom of choice. Generally, in this form of insurance a certain physician (or group of physicians) acts as gatekeeper: insurance holders need to have her referral for each (non–emergency) doctor visit. Managed care plans encountered a large popularity in the US since 1980, while in Europe they were completely unknown[3]. Switzerland was one of the first countries in Europe to allow pilot projects with managed care in 1990, and to introduce this type of contracts in its legislation in 1996, with the aim of containing health care expenditures. Nevertheless, this form of insurance was not successful at the beginning and remained quite marginal for more than a decade. In 2004 the share of managed care contracts was still less than 10%, a rather low percentage if compared to the US situation, where managed care contracts grew dramatically already in the Eighties'; in 1993, 70% of all the insured had opted for a managed care plan (Glied, 2000).

In Switzerland, the managed care contracts may assume different forms (Beck *et al.*, 2010). First, the so–called health maintenance organizations (HMO), i.e. organizations that provide (when physicians are independent) or arrange (if physicians are employed by the insurance)

[2]For a more detailed explanation of the Swiss health care system, refer to OECD/WHO (2011)

[3]We refer in this case to the European Bismarkian system, such as Switzerland, the Netherlands, Germany, France, Belgium. In fact, the NHS (in Italy or UK) may generally be considered as a sort of public managed care.

managed care for health insurances. Second, the general practitioner or family doctor model. Third, the preferred provider list, where the choice of the doctor is limited to a list of names chosen by the insurance. Also the call centers (known as "callmed option") can be considered a sort of managed care plan: a medical operator answers patients' questions and refers them to a physician, according to their need. Finally, the "insurance with bonus" is a contract that provides a discount on premium only if the individual has not sought care during the last year.

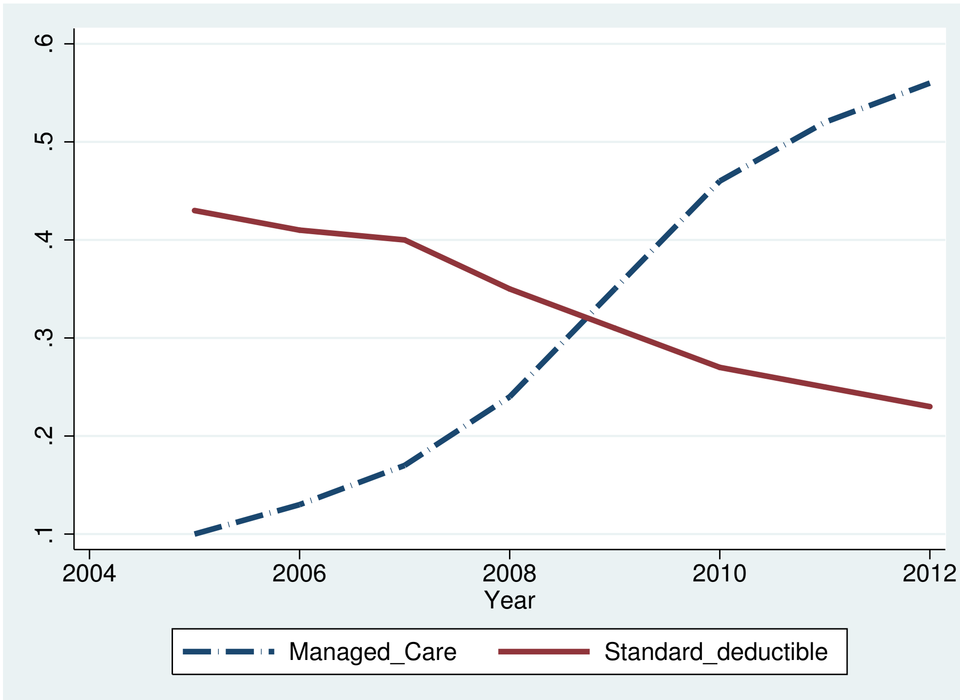
Different from the other managed care plans, HMOs are "models with capitation", meaning that the network is collectively responsible for a financial budget, and that the provider–network is remunerated by capitation for every individual enrolled in the plan. This financial responsibility is a strong incentive for the physicians not to prescribe unnecessary care. While some forms of family doctor plans may also be included among models with capitation, all the other forms of managed care are "non–contracted models", meaning that providers are not incorporated into a health insurance plan (Reich *et al.*, 2012).

Until 2007, managed care contracts remained quite uncommon: on average, they did not arrive to cover the 17% of total population. More precisely, their diffusion differed consistently among areas: in the north–east German cantons, the percentage of managed care was already close to 30% in the first years of 2000, while in the Romand cantons this percentage was significantly smaller (1–5%). However, since 2007, there has been a larger diffusion of this form of insurance in all areas of Switzerland. Their growing popularity may be explained also in the light of a change in the legislation. In fact, until 2007, a revision of the federal law allowed individuals to have only one exclusive reduction of premiums, choosing between a managed care plan or a higher deductible. Being the two options mutually exclusive, the majority of insured opted for a higher deductible, as it allowed for a larger discount. In 2007, a new law made it possible to combine both saving options (i.e. managed care plan and higher deductible), getting an even higher reduction on premiums. Pushed by the willingness to save money, a wide part of the Swiss population has opted for managed care contracts since 2007.

Graph 2.1 is based on administrative data and shows the increasing trend of managed care plans from 2005 to 2012 (dashed line). In only 7 years, their diffusion grew from 10% to more than half (56%) of the insured population. The other line represents the decline of contracts with standard deductible, as we will comment in the next subsection.

Managed care contracts became popular and appreciated both by insured and insurers. They allow insured to save money on premiums, and insurance companies (and institutions) to contain health care expenditures. Lehmann and Zweifel (2004) compute that the average treatment costs of HMO in Switzerland are about 62% lower than in the usual fee–for–service setting. In their analysis they find that 2/3 of savings (40 percentage points) are due to an increased efficiency and 1/3 of savings (20 percentage points) is due to the self–selection effect. A more recent research

Figure 2.1: The trends of managed care contracts and standard deductible, from 2005 to 2012.



Notes: Values from zero to one on the y-axis. Data come from the Federal Office of Public Health (2014).

by Reich *et al.* (2012) confirms these results for HMO and adds that the efficiency gains due to the family doctor model is about 15%, while the efficiency gain due to callmed model is less than 4%. Beck *et al.* (2010) use another methodology and find that the efficiency gain allowed by HMO is, on average, only 9%, but they detect a huge variation across HMOs.

Managed care plans are generally chosen by the "good risks", i.e. young and healthy people that are normally "cheaper" for insurance companies (Beck, 2004). This suggests that, for the overall equity of the health care system, these forms of insurance may lead to an increased inequity in the financing. Being the better–off induced to save money on premiums, they end up contributing to the total health care financing less than those who choose a standard contract (who are generally the worse–off and the middle–income class). This phenomenon leads to a more inequitable financing system. The first set of columns of Table 2.1 shows the share of people enrolled in the different plans of managed care by income quintiles, for the years 2007 and 2012. Data are extracted from the Swiss Health Survey, the dataset we use in the analysis. Family doctor is the most popular managed care plan and is especially chosen by people belonging to the central income quintiles. The HMO is slightly more widespread among the third, fourth and fifth quintiles. The shares of callmed and insurance with bonus are, on average, smaller than the other plans, the former being widespread quite equally among income quintiles and the latter being more diffused among the better–off. All in all, the shares of each of the managed care plans are bigger for 2012 with the only exception of insurance with bonus, which seems to have slightly reduced its popularity in 2102.

Table 2.1: Share of managed care and higher deductible for income quintiles. Years 2007 and 2012.

Income quintile	**Managed care**			**Higher deductible**		
		2007	*2012*		*2007*	*2012*
First	Insurance with bonus	0.5%	0.2%	Ded 1000 CHF	4.8%	5.4%
	HMO	2.1%	2.9%	Ded 1500 CHF	15.6%	10.6%
	Family doctor	9.5%	28.3%	Ded 2000 CHF	1.6%	1.5%
	Callmed	2.6%	5.1%	Ded 2500 CHF	8.4%	8.3%
Second	Insurance with bonus	0.8%	0.6%	Ded 1000 CHF	5.0%	5.2%
	HMO	2.3%	3.0%	Ded 1500 CHF	16.4%	13.6%
	Family doctor	9.5%	32.4%	Ded 2000 CHF	2.7%	3.7%
	Callmed	2.1%	6.5%	Ded 2500 CHF	9.3%	13.4%
Third	Insurance with bonus	1.0%	0.9%	Ded 1000 CHF	5.3%	6.0%
	HMO	3.4%	4.8%	Ded 1500 CHF	16.2%	15.3%
	Family doctor	8.7%	32.0%	Ded 2000 CHF	2.3%	3.7%
	Callmed	2.5%	7.3%	Ded 2500 CHF	9.9%	15.4%
Fourth	Insurance with bonus	1.5%	1.0%	Ded 1000 CHF	6.2%	7.0%
	HMO	2.7%	4.9%	Ded 1500 CHF	17.6%	15.9%
	Family doctor	8.5%	32.2%	Ded 2000 CHF	3.3%	5.1%
	Callmed	2.4%	6.8%	Ded 2500 CHF	12.1%	18.5%
Fifth	Insurance with bonus	1.7%	1.1%	Ded 1000 CHF	7.6%	6.9%
	HMO	3.2%	5.4%	Ded 1500 CHF	19.4%	17.9%
	Family doctor	7.1%	28.4%	Ded 2000 CHF	3.8%	5.9%
	Callmed	2.4%	8.7%	Ded 2500 CHF	16.9%	25.0%

Notes: The first set of columns reports the share of people enrolled in a managed care contract, by income quintiles. The second set of columns reports the share of people who chose a higher deductible, by income quintile. Data are extracted from the Swiss Health Survey for 2007 and 2012. Sample weights are used.

If, on the one hand, managed care contracts are an instrument to save money on individual premiums for the insured, on the other hand they might also act as a barrier to healthcare access. Gatekeeping itself, as well as the limited choice of doctors, may, in some cases, discourage people to seek care, particularly specialist care. The effect on equity in access is not evident. If managed care plans were mainly chosen by the better–off, this should not necessarily translate in an average worse inequity in access. Nevertheless, Table 2.1 does not reveal that managed care model meets only the preferences of the better–off. Therefore, we cannot exclude that part of the savings may actually hide an increase in inequity.

It has been demonstrated that in the US people with managed care contracts had lower ratings of care provided by their primary physician, but no association with a reduced access to specialists was found (Bambra *et al.*, 2014; Grembowski *et al.*, 2003)[4]. For Switzerland, no empirical evidence is available on this issue.

Increasing choice of higher deductible

Driven by the same reason of dealing with the increasing burden of health insurance premiums, many people opt for another solution: they take themselves part of the risk of becoming sick in order to get a discount on the monthly premiums. This is possible by choosing a contract with a higher deductible. With a standard insurance contract[5], the level of deductible is fixed at 300 CHF (about 330 US dollars). Yet insurance companies allow to choose among five other levels of deductible: 500 CHF, 1000 CHF, 1500 CHF, 2000 CHF or 2500 CHF. The higher the deductible, the higher is the discount on monthly premium for the insured. Similarly to managed care contracts, this has become quite a common choice, especially among young and healthy people, i.e. the categories of people that are generally considered "good risks". Assuming that their need for healthcare is small, they are prone to take themselves the risk of being sick in return for a discount on premium's amount. The red line of Graph 2.1 shows the percentage of people insured with the standard deductible of 300 CHF from 2005 to 2012. In only seven years, this percentage decreased by almost a half, changing from 43% to 23%.

The second set of columns of Table 2.1 shows, instead, the share of insured who chose a higher deductible in 2007 and 2012. More than the managed care plans, it is evident that higher deductibles are clearly preferred by the better–off. This holds for all the deductible options and both in 2007 and 2012. We also notice a larger diffusion of higher deductibles in 2012, with respect to 2007.

According to the literature, a reduction in the insurance coverage should often reduce the moral hazard effect (Gardiol *et al.*, 2005; van Vliet, 2004). Nevertheless, a smaller insurance coverage may also have a negative consequence. It may act as an entrance barrier to health care services, similarly to being enrolled in a managed care plan. It is true that, if we assume that

[4]The only exception is represented by the treatment of depression. In this case managed care in the US is associated with an inequitable access to specialist visits (Steiner and Robinson, 1998)

[5]We refer to standard contracts for adults and young adults (18–25 years old).

people are rational, they should choose the best contract for themselves using the information they have regarding their health status. Those who opt for a higher deductible should be those for whom it is convenient to have such a contract, i.e., those who do not expect to have a large need of health care. Nevertheless, it is also true that health is a particular good, not totally predictable. We cannot assume perfect information of individuals for themselves. Clearly, the unexpected need for care is not a concern for the better–off, who are willing to pay extra out–of–pocket expenditures. Different is the situation for the worse–off; it can be that they are discouraged to go to the doctor, despite their health care needs. Since the share of people in the lower income quintiles with a higher deductible is not negligible, this option may lead to increased inequity in access to care.

2.3 Methods: the estimation of inequity in health care services utilization

Previous research on this topic generally refers to the concepts of *horizontal equality* and *horizontal equity*. The former indicates a situation where people with the same economic resources have equal rates of health care use, while the latter is achieved when people in the same need for health care show equal rates of health care use. The measure of horizontal equality might become an index for equity only in the case in which the need for medical care does not vary with income, which is very unlikely.

The general approach in studying the inequality in health care utilization relies on the computation of the concentration index (CI), which is an extension of the Gini index. People are ranked according to their income and a concentration curve of the number of doctor consultations is computed. On the x axis the cumulative percentage of people ranked by their income is represented, while on the y axis there is the cumulative percentage of the number of doctor visits. The concentration index is two times the area between the 45–degree line (which represents perfect homogeneous distribution of the visits according to income) and the concentration curve. If the index is positive, the health care use is more distributed towards the rich, while a negative sign suggests that lower income people make a more extensive use of (or have more access to) the health care services (O'Donnell *et al.*, 2008).

Measuring the degree of horizontal inequity requires the same procedure but the CI is computed on the number of *need–standardized* doctor consultations.

The standardization is necessary to neutralize the effect of those variables that may influence the use of health services although they should not be taken into account in deciding the necessary number of visits. These factors, such as policy variables or personal socio–economic conditions, may be considered as illegitimate factors in the health care utilization decision process. The need–standardized health care use is the level of health care that everyone belonging to a similar health status should use. We can define it as the legitimate health care use associated to a determined health status. To get a need–standardized measure of health we use in this paper the approach proposed by Wagstaff and van Doorslaer (2000) known as "indirect standardization". The level of (indirectly) need–standardized care for individual i ($\hat{y}_i^{IS}$) is computed according to the following equation:

$$\hat{y}_i^{IS} = y_i - \hat{y}_i^x + \overline{y} \tag{2.1}$$

where y_i is the *actual* amount of utilization, $\hat{y}_i^x$ is the need–expected utilization and $\overline{y}$ is the sample mean.

$\hat{y}_i^x$ denotes for each individual the *predicted* amount of health care (defined as the health service utilization) that she would have used if she had been treated like other people with the same need characteristics. In our analysis it represents, in turns, the number of GP visits and the number of specialist visits. It is computed starting from the estimation of the following equation:

$$y_i = \alpha + \beta ln(inc_i) + \sum_k \gamma_k x_{k,i} + \sum_p \delta_p z_{p,i} + \varepsilon_i \tag{2.2}$$

where the health service utilization (i.e. the number of visits) is regressed on a set of variables including a vector of need variables $x_{k,i}$, (the logarithm of) income, and a set of non-need variables ($z_{p,i}$) that could have had a role in the decision whether to go to the doctor or not. We keep the actual values of the need variables and we replace the non–need variables – the income and the vector z – with their sample means, to isolate their effect on health care utilization. We include in the estimation also the means of these factors to avoid a problem of omitted variable bias, since they could contribute to explain health care utilization.[6] In the end, $\hat{y}_i^x$ is computed according to the following equation:

$$\hat{y}_i^x = \hat{\alpha} + \hat{\beta} ln\overline{(inc_i)} + \sum_k \hat{\gamma}_k x_{k,i} + \sum_p \hat{\delta}_p \overline{z}_{p,i} \tag{2.3}$$

The definition of needs relies on the health care that a person, on average, is expected to use, given some of her personal characteristics such as (generally self–assessed) health status, age and gender.

We compute the concentration index (CI) of y_i to assess the measure of inequality in the distribution of the actual number of visits. By contrast, the concentration index of the need–standardized health care use CI($\hat{y}_i^{IS}$) is a measure of the horizontal inequity and is called horizontal index (HI). Among others, van Doorslaer *et al.* (2004, 2006, 2000) compute this index

[6] As a sensitivity analysis, we performed all the computations both with and without the cantonal dummies. Since we got only negligible differences in the results, we decided to keep them out from our main analysis.

in a cross–sectional setting. The horizontal index suffers from a general limitation: it relies on the assumption that, after controlling for the need factors, any other difference in the use is to be attributed to non–need factors. The predicted value of doctor visits is indeed explained on the basis of need factors (and non–need factors are set equal to the sample average to neutralize their effect). This is a strong assumption if we consider the limited number of variables that are generally available to control for different health needs. In our paper we employ a better specification of need to mitigate this problem.

In this paper we focus the analysis on the total number of visits, and we also perform a two–step approach, splitting the probability to go to the doctor and the conditional number of visits (i.e. the frequency of visits after the first contact occurred). An advantage of this model is that it is easily interpretable within the framework of "principal–agent model": the patient (the principal) decides to seek care and the physician (the agent) decides the level of care she needs. This interpretation is very suitable for the Swiss context, where the deductible represents a strong entrance barrier. The decision to consult a doctor for a new episode of care can be considered as independent from the decision of how much health care to take, once the first contact occurred.

We use a simple OLS for the estimation of the total number of visits, while for the probability of visits we run a probit model and for the conditional number of visits we run a negative binomial model. The general functional form for a non–linear model follows equation 2.4:

$$y_i = G\Bigg(\alpha + \beta ln(inc_i) + \sum_k \gamma_k x_{k,i} + \sum_p \delta_p z_{p,i}\Bigg) + \varepsilon_i \tag{2.4}$$

Similarly to van Doorslaer *et al.* (2004), in the standardization process we set non–need variables equal to their sample means, by analogy with the linear case.

We compute the concentration index on the basis of a "convenient covariance" formula (Kakwani, 1980), namely:

$$CI = \frac{2}{\overline{y}} cov(y, r) \tag{2.5}$$

This formula relies on the fact that CI is equal to a weighted covariance between individual number of visits and individual's relative rank (i.e. her relative position in the income distribution). Given this relationship, it is possible to get to the following "convenient regression" formula (Kakwani *et al.*, 1997) of a (transformed) y_i on the individual weighted relative rank (r_i):

$$\frac{2\sigma_r^2}{\overline{y}} y_i = \alpha + \beta r_i + \varepsilon_i \tag{2.6}$$

where the coefficient β is equal to CI (O'Donnell *et al.*, 2008), σ_r denotes the variance of r_i, and the standard error of β is the standard error of the CI. The advantage of using this formula is the possibility to compute the CI standard errors. To correct standard errors for heteroscedasticity and autocorrelation we use the Neway–West variance–covariance matrix (Newey and West, 1994).

It has been noted (Wagstaff, 2005) that when the variable of interest is binary, the limits of CI are not necessarily -1 and +1. To deal with this problem, we use the Erreygers correction (Erreygers, 2009) for the estimation of the CI relative to the probability to go to the doctor, which adjusts the formula by the mean and the bounds of the variable in question. In this way, we also overcome the risk that the CI depends on the mean of the variable considered. We compute the Erreygers CI (ECI) according to the following formula:

$$ECI = \frac{4\overline{y}}{y_{max} - y_{min}} * CI(y) \tag{2.7}$$

Decomposition

Concentration index is a useful instrument to compare inequities across time and across countries, but it does not allow to explain the contribution of each factor associated with the inequity. The decomposition analysis is, instead, a descriptive tool useful to unravel the relative significance of each regressor on total inequity, and assess the relative magnitude of each source with respect to the others. From decomposition we can infer about the association of each factor with the dependent variable, but we cannot infer a causal interpretation. Decomposition only helps to shed light on causal association if equation 2.2 has a causal interpretation (van Doorslaer and van Ourti, 2011). Following the procedure by Wagstaff *et al.* (2003), we perform the decomposition analysis to assess the weight of each variable on the total concentration index of inequality. In the linear case the concentration index may be split in a weighted sum of the concentration indices of each regressor, where the weights are the elasticities of y with respect to that regressor. The CI can be rewritten according to the following formula:

$$CI = (\beta\overline{inc}/\overline{y})CI_{inc} + \sum_k(\beta_k\overline{x}_k/\overline{y})CI_k + \sum_p(\beta_p\overline{z}_p/\overline{y})CI_p + GC_\varepsilon/\overline{y} \tag{2.8}$$

where $\overline{y}$ represents the mean of y, $\overline{x}_k$ the mean of x_k, $\overline{z}_p$ the mean of z_p, GC_ε is the generalized concentration index of the error term ε. In this paper we perform the decomposition for the probability of going to the specialist that we previously estimated through a probit regression. Therefore we need to adjust the Equation 2.8 for a non–linear model. Similarly to van Doorslaer *et al.* (2004), we use the estimates of the partial effects evaluated at the means. We therefore rely on the following decomposition of CI:

$$CI = (\beta^m \overline{inc}/\overline{y})CI_{inc} + \sum_k (\beta_k^m \overline{x}_k/\overline{y})CI_k + \sum_p (\beta_p^m \overline{z}_p/\overline{y})CI_p + GC_\varepsilon/\overline{y} \tag{2.9}$$

where β^m, β_k^m, β_p^m, are the partial effects of each variable evaluated at the sample means. We are aware of the fact that the decomposition may lead to different results, according to the different choice of the reference group. Unfortunately this is the inevitable consequence of relying on a linear approximation. The approximation error is included in the generalized concentration index of the error term (together with the estimation error).

2.4 Dataset

We use data from the Swiss Health Survey (SHS), provided by the Swiss Federal Office of Statistics, for the years 2007 and 2012. This is a specific survey addressing health issues on the Swiss population. It is conducted every five years in the three main Swiss official languages (German, French and Italian) through assisted telephone interviews (CATI). Individuals are also required to fill a written questionnaire afterwords. SHS provides weights to make the sample representative of the Swiss population and to deal with non–response bias.
Our dependent variables are the answers to the question: "How many times have you consulted a GP/medical specialist in the last 12 months?"

Normally, studies on the inequity of health care use, measure the real health care need through the self–assessed health status. They often include also a dummy variable to control for individuals who suffer from a chronic illness or disability (van Doorslaer *et al.*, 2004).

The SHS contains more questions about the health status of individuals and we try to exploit all the available information. As proxies for the health status, we use the "standard" *self–assessed health status* divided in 5 dummies, one corresponding to each health level (very good, good, fair, bad or very bad), and a dummy for *chronic conditions*. We also add a set of dummies equal to one if the individual suffered from some health problems during the last 4 weeks, namely: *back problems*, *weakness*, *sleeping problems*, and *headaches*. We include a dummy if individual needed *informal assistance* in everyday activities in the last 12 months (available only for 2012) and, finally, a dummy for *obesity* (equal to one where the Body Mass Index is at least equal to 30).

Following the standard approach in the literature, we chose to consider *age* and *sex* as proxies for need and to include them in the equation as need–factors. More precisely, we assigned to each individual one dummy variable according to the sex and the age category (*female18–24*, *female 25–34*, *female 35–44*, *female 45–54*, *female 55–64*, *female 65+*, *male18–24*, *male 25–34*, *male 35–44*, *male 45–54*, *male 55–64*, *male 65+*).

Several forms of income are available in the SHS. We refer to the *household income* per equivalent adult, equalized through the OECD equivalence scale.

Among the non–need variables, i.e. variables that may affect health care but do not indicate need for health care, we include *education* and *marital status*, that could increase the probability to seek care, and activity status that could instead impact on the time price of health care use (van Doorslaer *et al.*, 2004). We use three dummies relative to education: one if the highest level of education obtained is compulsory education, one if the highest level of education obtained is secondary education, and one dummy if the individual got a degree. The dummy marital status is equal to one if the person is married or if she lives in a registered partnership and zero otherwise, while we control for the activity status with three dummies (*inactive*, *employed*, *unemployed*).

We also include in our analysis the variables relative to the choice for higher deductibles and the other insurance plans' options. In particular, we add one dummy variable for each different amount of deductible: *300 CHF*, *500 CHF*, *1000 CHF*, *1500 CHF*, *2000 CHF*, *2500 CHF*. We also add a dummy for each insurance contract available in the dataset: *standard contract*, *insurance with bonus*, *HMO*, *family doctor*, *callmed option*.

The total number of observations is 18.760 for 2007 and 21.597 for 2012. We drop from the sample young people up to 18 years old and missing values. We end up with a sample of 16.185 observations for 2007 and 16.449 observations for 2012. We apply cross–sectional sample weights at the individual level in all analysis.

2.5 Results

We now show the results of the inequity analysis for the GP and for the specialist visits. Results are presented for the whole of Switzerland and for the three main linguistic areas of the country: German, French and Italian. Differentiating results for linguistic areas is interesting for two reasons. First, from a health system point of view: it is worth reminding that Switzerland is a particular federal state where lower levels of government (i.e. cantons), have large autonomy in the organization of health care and other aspects of the social security system (e.g. the allocation of subsidies). Second, from an individual point of view: many of the personal choices depend also on the cultural influence, which is often specific to each linguistic area. For example, the individual choice of other forms of insurance than the standard one has been more popular in the German area, than in the other two. Both these reasons may have an impact on the equity in health care utilization.

Table 2.2 reports concentration indices of actual visits (CI) and horizontal index of inequity (HI), computed on the basis of the need–standardized visits.

GP visits seem to be pro–poor distributed (second column) in both years. However, the concentration index does not take into account the real healthcare needs of people. It is very

Table 2.2: Inequality (CI) and inequity (HI) in GP and specialist visits for Switzerland. Years 2007 and 2012.

	Total visit		**Probability**		**Conditional number**	
	CI	*HI*	*CI*	*HI*	*CI*	*HI*
GP visits						
Year 2007	**-0.0777**	-0.0321	-0.0028	0.0127	**-0.0776**	**-0.0399**
Year 2012	**-0.0956**	**-0.0025**	**-0.0638**	-0.0075	**-0.0710**	**-0.0339**
Specialist visits						
Year 2007	0.0230	**0.0417**	**0.0441**	**0.0855**	-0.0006	0.0028
Year 2012	0.0110	**0.0699**	**0.0480**	**0.1159**	-0.0106	0.0221

Notes: Inequality and inequity indices computed with an OLS model for the total number of visits, a probit for the probability and a negative binomial for the conditional number of visits. Significant indices reported in bold (p–value < 0.05).

likely that GP visits are more concentrated among the worse–off, them being also the less healthy. To check whether GP visits are pro-poor distributed, we have to look at the concentration index of the need–standardized number of visits, i.e. the horizontal index (second column of results). It is significant (although very close to zero) only in 2012, meaning that once standardized for needs, there is no evidence of pro–poor concentration. If we split the estimation into the probability of going to the doctor and the conditional number of visits, we find no significant concentration in the probability neither pro poor, or pro rich. By contrast, the conditional number of visits is more concentrated among poor people in both years. This result could be explained by the tendency of better–off people to go to the specialist after the first visit to the GP. This is partially confirmed by results for specialists, as we find evidence of a pro–rich concentration at least in the probability of a visit.

The simple distribution of specialist visits does not reveal any evidence of inequality neither pro–poor nor pro–rich (the coefficient is not significantly different from zero). But the concentration indices for need–standardized visits are positive, suggesting a pro–rich inequity.

Interestingly, it seems that the pro-rich inequity occurs only in the probability of a specialist visit. Rich people are more likely to go to the specialist doctor than poor people, but after the first contact, there is no inequity in the conditional distribution of visits (with the only exception of German areas that show a pro–rich inequity also in the conditional distribution, as we will see in Table 2.3). Considering the trend 2007–2012, we find evidence of a strongly increased inequity favoring the rich in the access to the specialist. The horizontal index increases by 36% changing from 0.08 to 0.11. This interesting result may indicate the existence of a sort of barrier for the worse–off to go to the specialist.[7]

Let us now discuss the concentration indices for GP and specialist visits in the German, French and Italian areas, separately. As Table 2.3 shows, we detect that the pro–poor inequity in the conditional number of GP visits is especially driven by the German regions, both in 2007 and 2012. It is worth highlighting that in both years the French area has an unexpected pro–rich (albeit not significant) inequity in the probability of going to the GP. If we consider the concentration of specialist visits, we notice that the need–standardized probability to go to the specialist is pro–rich distributed for all the three areas. In 2007 this inequity is stronger in the French area than in the German and Italian ones. In 2012 the three areas converge to a more similar result, everywhere bigger than the result of five years earlier. When we look at the concentration index of the conditional number of visits, we find that only in the German area the conditional number of specialist visits is pro–rich distributed.

If we compare our results to the previous findings for Switzerland (Leu and Schellhorn, 2004), we can claim that the inequity distribution of specialist visits has been following an increasing trend since 1997.[8] To frame Switzerland in an international context, we can refer to the study by van Doorslaer *et al.* (2004) where the authors compute the income–related inequalities in doctor utilization in 12 European member States. Apart from three exceptions (Belgium, the UK and the Netherlands), on average there seems to be no pro rich concentration in the probability to see a GP in the European countries. The conditional number of visits is generally pro poor, instead. Switzerland presents a situation similar to Italy or Greece, with a non-significant inequity in the probability to see a GP and concentration index of about -0.03 for the conditional number of visits. As far as the probability of seeing a specialist is concerned, our analysis shows a value of inequality similar to results for Portugal and Italy and anyway higher, whereas Austria, Denmark and Portugal perform worse for the level of inequity in the conditional

[7]Besides the success of managed care plans and higher deductible, it is worth mentioning also another possible factor of inequity. During the years 2006–2010 some people started to be excluded from the statutory coverage due to their economic inability (or unwillingness) to pay the insurance premium. In 2006 a reform of the Federal Health Insurance Act came into force and remained valid until the end of 2010. It stated that insurers could suspend their coverage of health care services to those who gave up paying the premium. For these people only emergency health care was guaranteed, with the obvious consequence that going to the doctor when needed (particularly to the specialist) was much more difficult for them.The number of people excluded from the statutory coverage largely grew in the years 2009–2011, exceeding the 130.000 units. This may also have brought inequity into the system. However, this phenomenon does not contribute to explain our results, since the sample we use does not include these individuals.

[8]It is worth mentioning that Leu and Schellhorn (2004) use a different specification of the health demand equation. This makes their results not completely comparable with ours.

Table 2.3: Inequality (CI) and inequity (HI) in GP and specialist visits for the three linguistic areas. Years 2007 and 2012

	Total visit		**Probability**		**Conditional Number**	
	CI	*HI*	*CI*	*HI*	*CI*	*HI*
GP visits						
Year 2007						
German area	**-0.0928**	**-0.0488**	-0.0101	0.0067	**-0.0896**	**-0.0544**
French area	**-0.0368**	0.0091	0.0270	0.0324	**-0.0475**	-0.0076
Italian area	**-0.0702**	-0.0127	**-0.0311**	0.0012	**-0.0599**	-0.0068
Year 2012						
German area	**-0.1015**	**-0.0024**	**-0.0726**	-0.0136	**-0.0739**	**-0.0399**
French area	**-0.0766**	-0.0016	-0.0296	0.0247	**-0.0643**	-0.0225
Italian area	**-0.1117**	0.0022	**-0.0597**	-0.0119	**-0.0898**	-0.0372
Specialist visits						
Year 2007						
German area	0.0212	0.0402	**0.0482**	**0.0877**	-0.0040	0.0011
French area	0.0460	**0.0677**	**0.0589**	**0.1092**	0.0178	0.0203
Italian area	-0.0238	0.0301	0.0247	**0.0656**	-0.0337	0.0006
Year 2012						
German area	0.0166	**0.0801**	**0.0537**	**0.1209**	-0.0073	**0.0291**
French area	0.0249	**0.0729**	0.0589	**0.1199**	0.0011	0.0234
Italian area	-0.0719	-0.0180	**0.0432**	**0.1167**	-0.0901	-0.0587

Notes: Inequality and inequity indices computed with an OLS model for the total number of visits, a probit for the probability and a negative binomial for the conditional number of visits. Significant indices reported in bold (p–value < 0.05).

number of specialist visits. A more recent work by Devaux (2013) computes the inequity index for Switzerland and other European countries, not included in the previous research. Results regarding inequity in specialist visits place Switzerland only before Belgium, Poland, Spain and France.

Decomposition analysis

Looking only at the concentration indices, one could think that the effect of increased inequality in the probability of going to a specialist in the years 2007–2012, may actually be due to a sort of entrance barrier for the worse off. Different forms of managed care plans, as well as a higher deductible, may induce the insured not to go to the specialist, maybe in favor of a less expensive GP visit. The decomposition analysis may clarify if these forms of insurance and the option for a higher deductible have a role in the dimension of inequality. Even though it does not allow to assess any causal relationships, it assign a relative magnitude to each contribution for each source of inequality (both legitimate and illegitimate). Table 2.4 shows for each variable the marginal effects obtained from the probit regression, its concentration index and its contribution in explaining the total inequality, based on Formula 2.9. All the values refer to the the probability of visiting a specialist.

Before discussing the results obtained from the decomposition analysis, we look at the marginal effects from the probit regression (first and fourth columns of Table 2.4).

Among the managed care plans, the family doctor is the only one able to reduce the probability to go to the specialist, but only in 2007. In 2012 the only significant result is given by the callmed option, but with a positive sign, suggesting that this plan increases the probability of going to a specialist. The results relative to the higher deductible are stronger and always significant (with the exception of the deductible equal to 500 CHF in 2007). We find that the higher the amount of the deductible, the lower is the probability of consulting a specialist. In 2012 a deductible of 1000 CHF reduces the probability to go to the specialist of 6%, while the effect of a deductible of 2500 CHF is close to 13%. Results are similar for the two years considered.

Among the other non–need variables, being richer increases the probability of a visit of 11% in 2012, similarly to having a degree (14%). This suggests that richer and more educated insured are more likely to go to the specialist than a less educated and worse–off individual with the same healthcare needs. Also being married slightly increases the probability to go to the specialist. The opposite is true for being employed (or unemployed in 2012) that reduces the probability of a specialist visit, compared to being inactive.

Results from the decomposition analysis show that in both years the most important contribution to inequality is given by the income itself, immediately followed by education, which is strongly pro–rich distributed. A more educated person is likely to be more informed about

Table 2.4: Marginal effects from a probit regression on the probability of a specialist visit and results from decomposition analysis. Years 2007 and 2012.

	Specialist visits: year 2007			**Specialist visits: year 2012**		
	Marg. Eff.	*CI*	*Contrib.*	*Marg. Eff.*	*CI*	*Contrib.*
SAH good	**0.0642**	0.0003	0.0000	**0.0463**	-0.0252	-0.0009
SAH fair	**0.1024**	-0.1110	-0.0026	**0.1068**	-0.1862	-0.0047
SAH poor	**0.1683**	-0.0020	-0.0901	**0.1877**	-0.2514	-0.0021
SAH very poor	**0.2456**	-0.2862	-0.0008	**0.2913**	-0.2026	-0.0005
chronic	**0.0828**	-0.0001	-0.0025	**0.1237**	-0.0374	-0.0027
back	**0.0446**	-0.0335	-0.0015	**0.0329**	-0.0473	-0.0011
weak	**0.0350**	-0.0508	-0.0016	**0.0238**	-0.0729	-0.0012
sleeping	**0.0416**	-0.0089	-0.0003	0.0249	-0.0359	-0.0005
head	0.0078	-0.0444	-0.0003	0.0126	-0.0233	-0.0002
obesity	**-0.0424**	-0.0946	0.0006	-0.0071	0.0001	0.0126
assist	-	-	-	**0.1671**	-0.0762	-0.0028
female1824	**0.4298**	-0.2456	-0.0065	**0.2487**	-0.2009	-0.0037
female2534	**0.4814**	-0.0481	-0.0040	**0.2529**	0.0472	0.0015
female3544	**0.4222**	-0.1891	-0.0206	**0.2178**	-0.0539	-0.0020
female4554	**0.3647**	0.0233	0.0017	**0.1994**	-0.0083	-0.0004
female5564	**0.2600**	0.1209	0.0067	**0.1098**	0.0659	0.0011
male1824	**0.1443**	-0.1898	-0.0016	**-0.1664**	-0.0538	0.0006
male2534	**0.1423**	0.1610	0.0032	**-0.2736**	0.1708	-0.0050
male3544	**0.1038**	-0.0578	-0.0014	**-0.2217**	0.0834	-0.0029
male4554	**0.1541**	0.0740	0.0020	**-0.2078**	0.0634	-0.0025
male5564	**0.1592**	0.2283	0.0064	**-0.1736**	0.1682	-0.0042
male65plus	**0.1796**	0.0664	0.0025	**-0.1044**	-0.0446	0.0009
log (income)	**0.0779**	0.0374	0.0527	**0.1082**	0.0325	0.0526
marital status	**0.0409**	-0.0802	-0.0038	**0.0311**	-0.0283	-0.0009
secondary edu	**0.0529**	-0.0704	-0.0050	**0.0728**	-0.0308	-0.0028
degree	**0.1035**	0.2567	0.0167	**0.1371**	0.3426	0.0163
insurance bonus	0.0570	0.2237	0.0003	-0.0945	0.2007	-0.0003
HMO	-0.0109	0.1149	-0.0001	-0.0092	0.1208	-0.0001
family doctor	**-0.0660**	-0.0473	0.0006	0.0123	-0.0047	0.0000
callmed	0.0345	0.0027	0.0000	**0.0690**	0.0830	0.0007
ded 500 CHF	-0.0041	-0.0127	0.0000	**-0.0264**	-0.0573	0.0004
ded 1000 CHF	**-0.0807**	0.0919	-0.0009	**-0.0612**	0.0634	-0.0004
ded 1500 CHF	**-0.0929**	0.0423	-0.0014	**-0.0879**	0.0830	-0.0019
ded 2000 CHF	**-0.1039**	0.1533	-0.0009	**-0.1182**	0.1895	-0.0016
ded 2500 CHF	- **0.1298**	0.1533	-0.0044	**-0.1287**	0.1837	-0.0067
unemployed	0.0010	-0.0493	-0.0016	**-0.1130**	-0.2993	0.0000
employed	**-0.0491**	-0.0986	-0.0062	**-0.0440**	0.0539	-0.0038
residual			-0.0059		0.0006	

Notes: The first and the fourth columns show the marginal effects from a probit regression on the probability of consulting a specialist. The second and the fifth columns (CI) report the concentration indices, while the third and the last column report the contribution to inequality. Significant indices reported in bold (p–value $<$ 0.05).

the overall health care system features and also about her health problems. Thanks to her information advantage, she may decide herself to go directly to the specialist without passing first through a GP.

All forms of managed care plans summed together give a contribution to inequality close to zero, both in 2007 and 2012. It means that they do not have a role in increasing inequity in access to care.

Summing together the contributions to inequality of the higher deductibles, we get a value equal to -0.0075 in 2007 and equal to -0.0103 in 2012. Therefore, the total contribution of this option is bigger in 2012, and it has a negative sign in both years. This means that the total contribution of higher deductible in explaining inequity is actually significant, but it goes in an unexpected direction. This option reduces inequity instead of increasing it.

Our results can be explained if considered together with the distribution of these forms of insurance among income quintiles (see Table 2.1). As already remarked in Section 2.2, higher deductible have found more success among richer people and in 2012 even more than in 2007 (with the only exception of deductible equal to 500 CHF that is pro–poor distributed, on average). Therefore, on average, a higher deductible may act as en entrance barrier more for the better–off than for the worse–off. It is likely that a high deductible prevents insured from going to the doctor when is not needed and the effect is stronger, the higher the deductible. A high deductible seems then to have a very positive effect. It seems to be effective in avoiding unnecessary care, reducing moral hazard behaviors. Moreover, it does not seem to be responsible for the worsening of equity in access, on average.

However, we also have to mention that people in the first quintile often receive a public financial aid (earmarked subsidy), thus they do not have an interest in getting a discount on premiums. This is a particular feature of the Swiss healthcare system, essential to explain our results. If the worse–off could not rely on a subsidy, many more individuals belonging to the first income quintile would have likely opted for a plan with a discount on premiums. In this scenario, the contribution to inequity of managed care and high deductible might have been different. Unfortunately, SHS data does not collect information on financial subsidies, therefore we cannot include it in our analysis.[9]

To sum up, our findings suggest that the choice of a higher deductible acts as an entrance barrier for insured, as expected. Nevertheless, being them more widespread among the better–off, this barrier prevents from going to the specialist more people belonging to the middle–rich income

[9]We computed the decomposition analysis on a subsample of individuals, as robustness check. We dropped from the sample the individuals belonging to the first income quintile, as a proxy for those receiving the earmarked subsidy. For the managed care plans we get results similar to the main ones, while the contributions to inequality of higher deductible maintain a negative sign, but their relative magnitude slightly decreases with respect to the main results. Therefore, our results are robust even though we are not able to control for the subsidies.

class. This option seems to have two positive effects. First, it rebalances the concentration of visits towards a smaller pro–rich inequity and second, it discourages moral hazard behaviors. On the contrary, our analysis does not reveal any contribution to inequality due to being enrolled in a managed care plan, neither in 2007, or 2012.

2.6 Conclusion

In this paper, we focus on the inequity in doctor utilization in Switzerland in 2007 and 2012. This spell of time coincides with a larger diffusion of managed care plans, as well as contracts with higher deductible. Both these options allow insured to get a discount on monthly premium. Managed care comes at the cost of less freedom of choice, while a higher deductible provides less insurance coverage. We aim to assess whether these forms of insurance may be responsible for an increased horizontal inequity in the health care utilization.

We investigate this topic empirically, computing the concentration indices of *actual* and *need–standardized* number of GP and specialist visits. We do this for the total number of visits, for the probability to seek care and for the conditional number of visits. Results suggest that, in line with previous literature, there is evidence of a pro–rich inequity in the distribution of specialist visits. Moreover, this inequity strongly increased in the five years considered, growing from 0.04 to 0.07. This increase is mostly reflected on the probability to go to the specialist. In all the three linguistic areas, the probability to go to the specialist is very inequitable, reaching values that are among the highest in the European context.

We also perform a decomposition analysis, which assigns to each factor a magnitude of their contribution towards inequity. We find that inequity in the distribution of specialist visits is mainly due to income and educational attainment.

Our results are an interesting contribution to assess to what extent managed care and the higher deductible impact on inequity in health care services utilization. The role that managed care plays on total inequality is close to zero. We get different results as far as the contribution of a higher deductible is concerned. This has a significant weight in the decomposition of inequality in access, but in the opposite direction. A higher deductible seems to mitigate the inequity in the distribution of specialist visits. This is due to the fact that this option is more widespread in the group of "good risks", who generally belong to the middle–high income class. Furthermore, the worse–off often get an earmarked subsidy, so they have no incentive to enroll in a plan that guarantees a discount on premiums. Therefore, we find that the plans with a high deductible seem to reduce the inequity in access to care. Moreover, managed care and higher deductible are even likely to reduce the moral hazard effect, preventing people from going to the specialist when is not really needed. The increased inequity in access seems to be mainly due to a much larger contribution of education in 2012 than in 2007.

This work presents some limitations. Results from the decomposition analysis help to shed light on the role of the different insurance plans on total inequality, but do not explain why we observe a worsening in the equity in the access to specialist visits. Our analysis does not consider the evolution of income distribution over time, nor the price of specialist visits. Both are factors that may have an impact on equity. Further research on this topic would be needed.

Equity in access is a principle valid in itself and it is even more significant if considered as a necessary step to reach the ultimate principle of equity in health. In this paper, we do not assess the trend in inequality in health for the years 2007–2012. It may be that more educated people get the best treatments by the best specialists and this might translate in a better health outcome for them. There is space for research to investigate whether and to what extent the inequities we found translate into inequities in health outcomes.

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Chapter 3

Short term effects of public smoking bans on health

Keywords
Smoking ban, AMI, ETS, health inequality.

3.1 Introduction

The negative effects of smoking—both primary and second hand—are well known and subject of many studies. In a recent review study, Cawley and Ruhm (2011) show that smoking is the most important modifiable risk factor related to both mortality and disability and the most damaging health behavior in high-income countries. A recent report by the American Hearth Association (Go *et al.*, 2014) estimates a total cost associated with smoking of $193 billion per year, 50% due to direct medical costs and 50% due to productivity costs. Smoke also contributes to health inequality since smoking prevalence and environmental tobacco smoke (ETS) exposure[1] are higher among people from low socio-economic status (SES) (e.g. Cutler and Lleras-Muney (2010)). The public awareness about the large negative effects of smoking induced several governments all over the world to enact a series of measures to reduce tobacco consumption and ETS exposure. Among them, smoking ban in public places is one of the most important public health intervention implemented in many developed countries in the last decades. However, several concerns have been raised—in particular among economists (e.g. Adda and Cornaglia (2010))—about the real health benefits deriving from this public health intervention.

This paper investigates the short term effect of public smoking bans on the incidence of acute

[1] In this paper ETS refers only to second-hand smoke, namely being exposed to someone else's cigarette, cigar, or pipe smoke.

myocardial infarction (AMI), which represents the main cause of death in almost all OECD countries (OECD 2013). In particular, exploiting the time and geographical variation in the introduction of smoking ban in public places in Switzerland, we show that the incidence of AMI decreases on average by about 10-12% immediately after the law implementation with large heterogeneity across population sub-groups. Moreover, using this quasi-natural experiment, we provide evidence of a causal association between ETS and AMI.

Previous economic studies mainly analyze the effect of smoking ban on smoking behavior, while they are limited in the ability of studying the effects of smoking bans on objective health outcomes. This might be explained by the fact that the negative effects of smoking on most of the associated diseases may generally require long times of observation.

For instance, the effect of smoking on lung cancer may be detected only in the long run. A change in smoking behavior —such as the individual decision to quit smoking— is, by contrast, easier to measure. Therefore, it often represents the only feasible outcome to assess the effect of the anti-smoking policy in the short run. However, the evidence from these studies is quite mixed. In the US, Evans *et al.* (1999) show that smoking ban on workplaces reduced smoking prevalence by 5%. On the other hand, Adda and Cornaglia (2010) report evidence of displacement of ETS from public to private places and no effect on smoking prevalence. In Canada, (Carpenter *et al.*, 2011) find that these policies had no effect on smoking prevalence, but induced large reduction in public-places ETS exposure without any private displacement. The introduction of smoking bans in public places had limited or no effects on smoking prevalence also in Germany (Anger *et al.*, 2011) and UK (Jones *et al.*, 2013), while there is contrasting evidence regarding Italy (Bono *et al.*, 2014; Buonanno and Ranzani, 2013). Given the well-know negative association between smoking and body weight (e.g. Cawley *et al.* (2004)), some studies evaluate the effect of this policy also on other health behaviors, showing some evidence of weight gains and increasing alcohol consumption after the introduction of the smoking ban in Italy (Pieroni *et al.*, 2013) and in the US (Liu *et al.*, 2010). All in all, the risk of negative spillovers on other behaviors, together with the evidence of displacement from public to private places, might explain the previously mentioned concerns about the health benefits of this policy.

We contribute to the literature by focusing on AMI hospitalizations, which allow us to measure the short-run effects of smoking bans on an objective metric of health. A wide strand of the literature argues that smoking is the most important avoidable cause of coronary diseases (Glantz and Parmley, 1991). In particular, several studies show that it is associated with a greater risk of AMI, especially among men (Teo *et al.*, 2006). From a medical point of view, exposure to smoke is associated with an endothelium-dependent dilatation, which may lead to an early arterial damage (Celermajer *et al.*, 1996), as well as a tendency towards an acute enhanced platelet aggregation (Sargent *et al.*, 2004) and reduced health variability (Pope *et al.*, 2001).

All these factors occur rapidly, within 30 minutes after exposure to smoke, increasing the risk of myocardial infarction (Barnoya and Glantz, 2005). At the same time, these acute negative effects disappear in a few hours after the end of the exposure (Argacha *et al.*, 2008). Therefore, a reduction in the number of AMIs should be quantifiable immediately after the introduction of the smoking ban, starting from the initial days after the reduction of smoking exposure (Richiardi *et al.*, 2009). This means that we can evaluate the short-term effect of this policy.

We are fully aware of the previous epidemiological literature investigating the effect of smoking bans on cardiovascular diseases. For instance, in a meta-analysis of the existing literature, Meyers *et al.* (2009) argue that after the implementation of the smoking ban, AMI risk decreases by 17% overall, with a bigger effect among younger individuals and non-smokers. However, there are fundamental concerns regarding the identification assumptions behind the estimation strategies implemented in these studies. Most of them are based on a simple pre and post comparison in a single city or country (e.g. Di Valentino *et al.* (2011) on the Swiss canton Ticino), without accounting for all potential confounders, such as preexisting trends or the effect of contemporaneous policies. To this end, Shetty *et al.* (2011) find that the relationship between smoking bans and AMI in the US is sensitive to the choice of the city and specification. For instance, using data about US-Medicare enrollees aged 66+ for the years 1999–2008, Barr *et al.* (2012) show that the negative effect of smoking bans on AMI hospital admissions disappears as they account for non-linear adjustment for secular trend and random effects at the county level. Finally, some of these studies evaluate the effect of smoking bans in very small cities and this may lead to problems of external validity of their results (e.g. Sargent *et al.* (2004)).

In our estimation strategy we address all these concerns. In particular, we exploit the Swiss quasi-natural experiment to assess empirically whether there is a causal effect of the smoking ban on AMI hospitalization. Switzerland is a perfect environment to study, since the 26 Swiss cantons implemented the public smoking ban at different points in time, making it possible to exploit both time and geographical variation. We use monthly data on the universe of Swiss hospitals admissions observed from 2004 to 2012, broken down by age and sex. This allows us to evaluate the heterogeneity of our results. In particular, the reduction in AMI hospitalization has been larger in absolute terms for men aged 50–65 and 66+, while we find some effects only on women 66+, the only subgroup among women at risk of AMI. It is worth noting that in relative terms (compared with the mean incidence of the disease in the population) the larger effect is found among men under 66 (both under 50 and 50–65). To explain such result, we use survey data from the Swiss Household Panel showing that younger people are more likely to go to bars and restaurants and so more exposed to ETS before the policy implementation. Such intuition is corroborated by self-reported information on ETS exposure, that we extract from the Swiss Health Survey, that show a large age gradient in ETS exposure in 2007 before the policy was enacted.

Even more interesting is the evidence we report of a dramatic drop in the average ETS exposure in 2012 (after the implementation of the policy) that was not followed by any reduction in smoking prevalence and cigarette consumption. This interesting result —consistent with most of the literature so far mentioned— suggests that the reduction in AMI hospitalization is only due to a reduction in ETS exposure and not to a reduction in smoking consumption. After all, the reduction of ETS exposure, and its negative consequences (i.e. AMI), represents the main goal of the policy, while the reduction of smoking prevalence and consumption might be at most considered a potential positive spillover. Hence, we believe that our paper can better evaluate the effectiveness of this kind of policy than most of the previous economic literature that only analyzes its effects on smoking prevalence and consumption.

In our analysis we also exploit the residential information on the hospitalized patients to evaluate the heterogeneity of our results across smaller areas (106 smaller geographical units) with different level of income and education, holding constant the confounding effect of different institutional settings at cantonal level. In particular, we find that the effect of the policy on AMIs has been stronger in low income and low education areas. As before we use survey data on smoking prevalence and ETS to explain our results. As expected among low SES people, there is the largest incidence of both smokers and the largest (self-reported) exposure to ETS.

We evaluate the robustness of our results to alternative model specifications. In our estimation strategy we account for many factors that might confound our results, such as area (canton or region) and time fixed effects, as well as area specific time trends. We also control for local unemployment rate to take into account the well-know pro-cyclical behavior of smoking and AMIs (Ruhm, 2005, 2007) and to take into account the effect of the last financial and economic crisis.

We indirectly test our main identification assumption by checking whether the policy also affected the incidence of lung, prostate and breast cancer which are known not to have been affected by the policy in the short run. We also show that the reduction was larger during the winter period, when people are more likely to seat inside public places. Moreover, we implement a "placebo" test (by using one year lag and lead values for the treatment) to assess concerns about the potential endogeneity of the policy (Carpenter *et al.*, 2011) and to rule out the possibility of other confounders or external shocks driving our results.

Our results are particularly relevant from a health policy perspective. First they shed a light on the doubts raised by previous literature on the effectiveness of this policy. We report evidence of large positive benefits on health already in the short-term, by showing a reduction in AMI incidence, the most important cause of death related to ETS together with lung cancer. Second, our results suggest potential benefits also on health inequality, since we show that the ban had stronger effects in areas characterized by lower SES (low income and education).

The next section describes how and when the smoking bans were implemented in the Swiss cantons. Section 3.3 presents the data and some descriptive evidence. Section 3.4 explains the choice of the model and discusses some empirical issues. Section 3.5 shows and compares our main results and introduces some robustness checks. Finally, section 3.6 concludes.

3.2 Swiss institutional setting

Smoking bans in Switzerland were implemented only after a long series of anti-smoking initiatives undertaken some years before and mostly defined at federal level. In 2001 the Swiss Federal Council adopted the National Tobacco Control Program, aimed at sensitizing people towards the bad effects of smoke and at changing smoking behavior. The Program devoted economic resources for anti-smoking campaigns and other measures and was meant to last until 2005, but it was extended until 2008 and subsequently until 2016 (under the new name of National Tobacco Program). To guarantee the necessary economic resources for anti-tobacco initiatives, in 2004 the Federation set up also the Tobacco Control Fund, financing it through a tax of 2.6 centimes on each cigarette packet sold. It supported measures aimed to assist people who wanted to give up smoking and to protect non-smokers against second-hand smoke. It was in this setting that in April 2007 canton Ticino (TI) decided to implement a smoking ban in all public places (i.e. bars, restaurants and clubs), with the aim of contrasting smoke and protecting the population against second-hand smoke. Ticino was the first canton opting for such a policy. Canton Graubunden (GR) followed Ticino one year later.

In May 2010 a federal law regarding the implementation of the smoking ban came into force (it was enacted already in October 2008): smoking in bars and restaurants was prohibited in each canton. The federal law was less restrictive than the one approved in canton Ticino. It allowed restaurants to have some rooms for smokers, and even the service may be guaranteed in some cases. However, cantons were free to set the federal law only as a lower bound and to choose more restrictive regulations. Anticipating the implementation of the federal law, most of the other cantons adapted and introduced the public smoking ban during 2009 and 2010. The last canton to introduce the ban was Appenzell O.Rh. (AR) at the beginning of 2011.

Table 3.1 shows for each canton (first column) when the smoking ban in public places started to be in force (second column) and the strictness of the law that they chose to adopt (third column).

The majority of cantons decided for a more severe regulation, meaning that they chose not to dedicate special rooms for smokers. It is worth remarking that there seems to be a sort of "linguistic" trend in the timing of smoking bans implementation: the Italian speaking cantons were the first cantons to implement this policy, followed by most of the French speaking cantons and lately, by the German speaking cantons. It is likely that the choice of a canton whether to implement the smoking ban was also affected by the choice of neighbor territories, both Swiss and non-Swiss.

Finally, it is important to stress that only the smoking ban in public places was set at cantonal level, while all the other anti-smoking policies, such as the excise tax, were set at federal level.

3.3 Data

Administrative data

Our main analysis is based on data from the "Swiss Hospital Statistics" (HOST), provided by the Swiss Federal Statistical Office. It is a wide dataset collecting information on each patient for the universe of the Swiss hospitals from 1999 to 2013; after 1998, it became compulsory for each hospital to participate at this statistics. HOST includes basic demographic information (age, sex and residence) for each patient, and information on the main diagnosis, as well as secondary diagnoses. For our analysis we use data from 2004 to 2012, due to the lack of some hospitals' data for the years before 2004. Since the dataset is based on individual data, we are able to explore the heterogeneity across different classes of population by carrying out separated estimates for different groups of patients (according to age and sex). Concerning patient's geographical residence, Switzerland was split in more than 700 geographical regions called "Medstat regions": each of them counting from 3.500 to 10.000 inhabitants and including territories similar from a socio-economic and geographical point of view. In 2008 the Swiss Federal Statistical Office implemented a major redefinition of the Medstat regions, changing substantially the agglomeration of municipalities that form each region. The rationale behind this redefinition was to have more homogeneous and comparable areas across the country. Unfortunately, for the aim of our research, this change does not allow us to analyze the impact of the smoking ban policy at this geographical level because the same areas are not comparable before and after the redefinition. For this reason, we consider two other geographical levels for our analysis: the 26 cantons and the so-called "Spatial Mobility" (SM) regions, 106 geographical units that are smaller than cantons but larger than the Medstat regions.[2]

[2]Knowing the zip codes grouped in each Medstat and in each SM region, we are able to group the Medstat regions into the larger SM regions. Nevertheless, in some cases the Medstat regions cannot be uniquely assigned to one SM region. In these cases we split the number of AMIs of the Medstat region and we assign it to the SM

Table 3.1: Implementation of smoking bans (SB) in the Swiss cantons

Canton	**SB implemented in**	**More severe regulation**
AG	05.2010	
AI	05.2010	
AR	01.2011	X
BE	07.2009	X
BL	05.2010	X
BS	04.2010	X
FR	01.2010	X
GE	11.2009	X
GL	05.2010	
GR	03.2008	X
JU	05.2010	
LU	05.2010	
NE	04.2009	X
NW	05.2010	
OW	05.2010	
SG	07.2010	X
SH	05.2010	
SO	01.2009	X
SZ	05.2010	
TG	05.2010	
TI	04.2007	X
UR	09.2009	X
VD	09.2009	X
VS	07.2009	X
ZG	03.2010	
ZH	05.2010	X

Carrying out the analysis at cantonal level is an advantage because it coincides with the political area of the introduction of the smoking bans. On the other hand, focusing on the SM regions (hereafter regions) allows us to exploit the heterogeneity of the effect across income and education levels (see also Section 3.4) holding costant the effect of institutional factors at cantonal level. To be able to exploit these potential sources of heterogeneity, we added to our dataset some socio-demographic information: population, income, education and unemployment rate. Data on population for cantons and municipalities are provided by the Swiss Federal Statistical Office. We aggregate data of municipalities to obtain data at regional level. However, the population breakdown by sex and age is available only for cantons, while for municipalities this information is provided only in the 2000 census. To obtain a proxy of the regional population by age and sex for the years 2004–2012, we multiply the value of the regional population of the year 2000 (from the municipal census) for the rate of variation of the population of the canton to which the region belongs. We do this for each age and sex category.

Monthly data is available only for unemployment rates, while for income and education we have data only for few years. We extract data on unemployed from the website of the State Secretariat for Economic Affairs (sector of job market statistics) for the years 2004–2012. Data is taken from the official registers of unemployed. It is monthly updated and available both at cantonal and regional level.

Data on education is available in the Statistical Atlas of Switzerland for the year 2000. For each region we compute the share of people with only compulsory primary education, the share of people with a degree of second level (high school) and the share of people with a degree of third level (college). Finally, we use data on income from the Swiss Federal Tax Administration website for the year 2000–2010.[3]

Survey data

In this paper we also use survey data from the Swiss Health Survey (SHS) and the Swiss Household Panel (SHP) to investigate smoking habits, ETS exposure and life style of various subgroups of the population.

The SHS was meant to provide measures of the Swiss population's health status and health behaviors. The SHS has been carried out once every five years since 1992 by the Federal Statistical Office (FSO). The last survey took place in 2012 as part of the data collection programme of the Swiss population census. The survey is carried out in German, French and Italian using

regions to which the Medstat region belongs. We use as a weight the number of zip codes of the Medstat that belong also to the SM region divided by the total number of zip codes that compose that Medstat.

[3]We cannot insert the income in our estimation as a time varying variable because up to now the last available year is 2010. For this reason we use a mean of the median income in each region in the period 2004–2010.

computer assisted telephone interviews (CATI) followed by a written questionnaire. The target population consists of the resident population in private households aged 15+. Overall, 11'314 women and 10'283 men were questioned. The survey provides sample weights to take into account the sample design and non-response. See OFS (2013) for additional information on the survey design and methodology. From the 2007 and 2012 survey of the SHS, we extract information on smoking behavior (smoking prevalence and cigarette consumption) and ETS (average minutes of passive smoke per day) broken down by education and age.

The SHP constitutes a unique longitudinal database for Switzerland. The survey covers a broad range of topics including information on leisure time and life style that we use to investigate who are the people that are more likely to attend public places (i.e. bars, pubs and restaurants) and so more likely to be affected by the smoking ban policy. Data collection started in 1999 with a sample of 5,074 households containing 12,931 household members. A refreshment sample was added in 2004 and since 2013 the SHP contains a third refreshment sample of 4'093 households and 9'945 individuals. See Voorpostel *et al.* (2014) for additional information on the SHP.

3.3.1 Descriptive statistics and graphical evidence

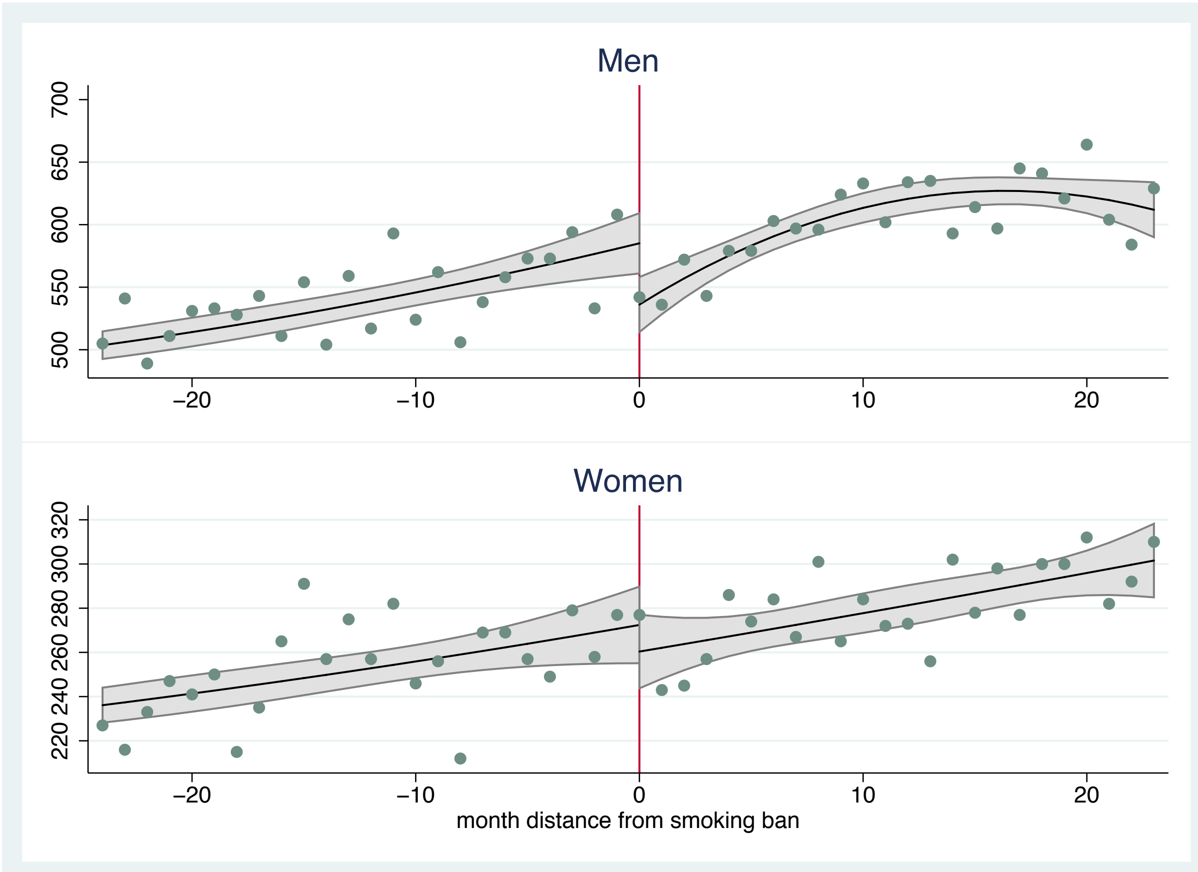
Table 3.2 shows in the first two columns the monthly average and standard deviation of AMI hospital admissions for the whole population by area and by sex in the years 2004–2012.

In the third column we report the corresponding incidence per 100'000 inhabitants by dividing the number of AMI cases in each canton for the corresponding population. Similarly, the last column reports the incidence based on regions. Consistent with the medical literature, two thirds of the AMI hospital admissions are men, in particular among those in the age class 50–65 and 66+, while for women the incidence of AMIs is almost only among women aged 66+. There is evidence of substantial variability in the number of AMI hospital admissions not only across cantons but also over time. We observe an increasing trend in AMI hospital admissions over the years under investigation: this may depend on better measures of prevention, on better access to health care and on changes in the definition of the diagnosis, especially after a recent reclassification of some angina patients into AMI cases (Insam *et al.*, 2013). We crosschecked our dataset with data from the AMIs Plus – the National Registry of AMI in Switzerland – and we confirmed the consistency of HOST data and the presence of the increasing trend. There is also evidence of seasonality with a larger incidence of AMI during the holiday season. This descriptive inspection suggests that taking such spatial and time heterogeneity into account in our estimation strategy is crucial. HOST collects about 90% of the total estimated number of AMI cases: the remaining 10% are people who die from AMI before arriving to the hospital (Meyer *et al.* (2007) perform this estimation for the year 2004). This small percentage is expected to further decrease over time, due to the easier access to hospitals in many areas. Notice that the

Table 3.2: Average number of AMI hospitalizations per month at cantonal and regional level and incidence per 100'000 inhabitants, by age and sex

	CANTONAL VALUES			**REGIONAL VALUES**		
	# AMIs	**(S.D.)**	**Incidence**	**# AMIs**	**(S.D.)**	**Incidence**
All						
Total	29.556	(31.328)	10.094	7.250	(7.544)	10.324
Under 50	3.580	(4.266)	1.798	0.878	(1.316)	1.882
50–65	8.654	(9.752)	15.721	2.123	(2.524)	15.629
66+	17.321	(18.557)	37.630	4.249	(4.758)	37.874
Men						
Total	20.227	(21.598)	14.124	4.961	(5.242)	14.366
Under50	3.069	(3.733)	3.022	0.753	(1.186)	3.165
50–65	7.032	(8.006)	25.481	1.725	(2.124)	25.169
66+	10.126	(11.007)	55.150	2.484	(2.908)	52.020
Women						
Total	9.329	(10.291)	6.182	2.288	(2.826)	6.335
Under 50	0.511	(0.941)	0.530	0.125	(.395)	0.553
50–65	1.623	(2.204)	5.941	0.398	(.781)	5.835
66+	7.195	(8.061)	25.604	1.765	(2.312)	26.988
N	2808			11448		

Figure 3.1: Discontinuity in AMI hospital admissions after the smoking ban by sex

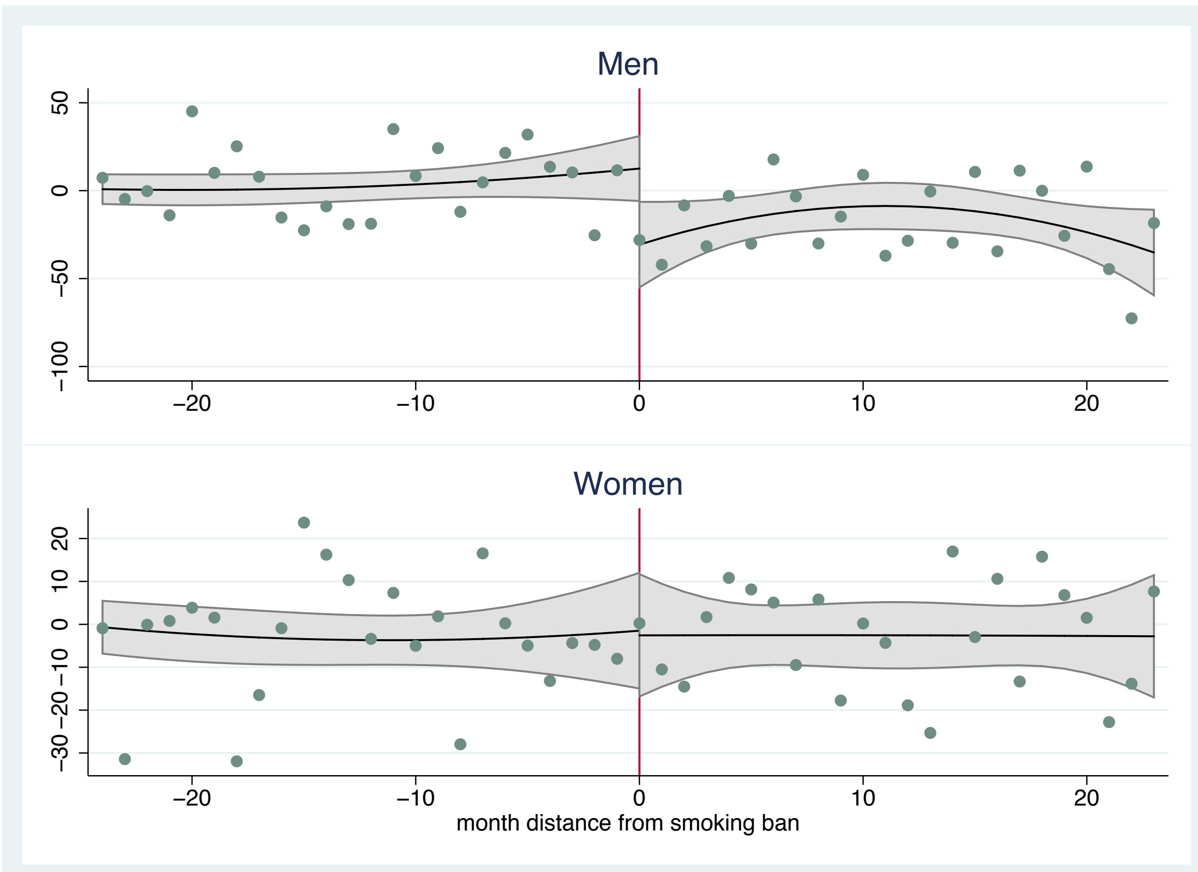


mean incidence of AMI hospitalizations is slightly larger at regional level than at cantonal level. This difference arises because the cantonal mean gives the same value to all cantons, regardless of the population's size. By contrast, the mean incidence at regional level is closer to the true national value.

In Figures 3.1 and 3.2, we show first evidence of a statistically significant discontinuity in the number of AMI hospital admissions at national level in correspondence of the smoking ban implementation for men but not for women.

Since each canton implemented the policy at different points in time, we aggregate AMI cases across different cantons based on the distance in months from the policy implementation, as in a sort of Regression Discontinuity Design (RDD). More specifically, in Figure 3.1 we show the raw data, while in Figure 3.2 we report the aggregate residuals from a regression in which we control for year and canton fixed effects. This second figure is meant to take into account the time and geographical heterogeneity described before. The vertical line represents the "month

Figure 3.2: Discontinuity in AMI hospital admissions after the smoking ban by sex (residuals)



zero", namely the month during which each canton introduced the ban. Purely for descriptive purposes, the fitted lines are based on a quadratic fit on the two years before and after the implementation of the policy, while the shaded areas correspond to the 95% confidence interval.

For men, Figure 3.1 already shows evidence of a discontinuity in AMI cases after the introduction of the smoking ban, although we are not controlling for temporal trends and cantonal differences. This discontinuity is confirmed in Figure 3.2, where we take into account these potential confounding factors. In particular, the residuals are slightly above zero in the two years before the discontinuity, while they rapidly decrease towards –40 in the months immediately after the smoking ban implementation. For women, instead, we do not have evidence of any discontinuity at the cut-off date.

Unfortunately, we are not able to show similar evidence for cantons since the number of AMI cases each month is too small to show graphical evidence of a significant discontinuity. In the next section, we investigate the effect of smoking bans on AMI hospitalizations trying to take into account all the potential factors that might confound the effect of the policy. It is worth reminding that, since Figure 3.2 nets out canton and time fixed effects, we are already controlling for the most important confounding factors in the relationship under investigation.

3.4 Empirical strategy

The empirical strategy exploits the geographical and time variation in the implementation of the smoking ban. In the analysis that follows, the outcome of interest is represented by the AMI incidence rate, AMI_{kt} (as reported in column 3 of Table 3.2), that is obtained dividing the absolute number of patients hospitalized for AMI in area k (it can be either canton or region) and at time t by the corresponding population of the area k at time t. We prefer this variable to the absolute number of AMI cases because it allows us to control for the different dimension (and population structure) of our Swiss regions. Moreover, it nets outs the effect of differential population trends across regions.[4] More specifically, we estimate the following equation:

$$AMI_{kt} = \beta_{0k} + \beta_1 SB_{kt} + \beta_2 X_{kt} + \delta_t + f(k,t) + \epsilon_{kt} \tag{3.1}$$

where SB_{kt} is a dummy equal to one if the smoking ban was enacted in area k at time t; the coefficient β_1 measures the effect of the policy. Our specification also includes area, β_{0k}, and time, δ_t, fixed effects and, as additional robustness check, area specific time trends, $f(k,t)$. More specifically, we model the function $f(k,t)$ as a linear function (i.e. multiplying the geographical dummies for a linear time trend). Finally, X_{kt} includes monthly data on unemployment rate.

[4]The other advantage of using the incidence rate is that we can rely on linear regression models and avoid the issue of estimating count data models.

Ruhm (2005, 2007) show that both smoking and AMI discharges are pro-cyclical, so we might expect a negative correlation between unemployment and AMIs.

Time is measured in months (108 months starting from January 2004 to December 2012). As already mentioned in Section 3.3, the geographical unit under investigation, the area k, can be either the canton or the region. Since the policy is implemented at cantonal level it is natural to consider the canton as the reference geographical unit. However, we believe that it is still worth focussing the analysis on a smaller geographical area for two main reasons. First, it allows to analyze the heterogeneity in the effect of interest across educational and income levels, holding constant the confounding effect of institutional differences at cantonal level. In other words, the analysis at the level of regions exploits the geographical variation within cantons in the level of education and income. Second, the increase in the number of observations allows us to increase the power of our estimation strategy and to reduce the relative weight of smaller cantons. As discussed in Section 3.3.1, the analysis at cantonal level put the same weights to cantons of different dimensions.[5]

Given the panel nature of the dataset, the baseline estimation method for this model is the within estimator (fixed effects), which allows to remove all time-invariant confounders at area level. Our estimates are separated by sex and age group (below age 50, age 50–65 and age 66+) to take into account the large differences in the incidence of AMIs across these population subgroups. We also implement separated estimates according to the severity of the policy implemented by the different cantons (severe vs. mild policy, see Section 3.2 for further details).

3.4.1 Identification issues

As discussed throughout the paper, the identification of the effect of the smoking ban on AMIs comes from the time and geographical variation in the implementation of the policy across cantons. Notice that different from a standard "diff-in-diff" strategy, we do not need the classic common trend assumption, since the policy has been implemented in all cantons but at different points in time. Therefore, it becomes crucial that there are no other policy changes or cantonal shocks coinciding with the introduction of the smoking ban and affecting the number of AMIs. The specification proposed in equation (3.1) should be sufficiently flexible to account for the potential confounding factors at geographical level since it also includes area specific time trends. In addition, our specification includes local unemployment rates to address the concern that the policy implementation is somewhat contemporaneous with the last financial and economic crisis. However, this is unlikely since the policy implementation has a very large temporal variation

[5] As robustness check we implemented our cantonal estimates using weights based on population size. The results we found are slightly larger than those reported in the main text and so very close to the effect estimated using the regions.

that does not coincide with the timing of the crisis (2008–2009).

More generally, we believe that it is very unlikely to have any confounding factors correlated with the policy implementation in each canton and able to affect the number of AMIs at the same time. Unfortunately, we cannot test it directly, but we implement a battery of robustness checks that supports our identification assumption. First, we implement a sort of "unconfoundedness test" in which we test whether the policy had an effect also on the incidence of lung, prostate and breast cancer, that should not be affected by the policy in the short run. Second, as in Carpenter *et al.* (2011), we implement a placebo test in which we assess concerns about the potential endogeneity of the policy by controlling for one year–lead of the smoking ban. Such a test is implemented to verify whether we are confounding the effect of the smoking ban law with other unobservable trends that vary across cantons and in particular whether "*large shocks to outcomes systematically precede rather than follow*" smoking ban laws.[6] We test for the presence of jumps in the AMI hospitalizations distribution one year lag and lead the real implementation of the policy in each canton. Another robustness check we implement is to check whether the effect of public smoking ban is larger during the winter season. The intuition is that during these months people that go to pubs and restaurants are more often inside public places because of the adverse weather conditions. Whereas during the other seasons people should stay more often outside drinking and eating and so they should be less exposed to ETS. So, if we are not confounding the effect of the ban with that of other factors, we should observe a larger reduction in AMIs during the winter season.

Even tough all these tests support our identification strategy, some concerns may still arise regarding the estimation of the standard errors, since in panel data errors are often correlated both across time and regions. In this case it might not be sufficient to cluster standard errors only at the statistical area level (as suggested by Bertrand *et al.* (2004)). It is indeed very likely that the number of AMI cases is correlated also spatially across cantons. The decision of one canton to introduce a smoking ban may have an influence also on the choice of a neighbor canton. In our case, it is quite evident that the timing in the implementation of the policy is somewhat affected by some geo-cultural factors. In particular, Ticino was the first canton to implement the policy in 2007 (probably affected by the Italian smoking ban implemented since 2003), followed by Graubunden (always at the border with Italy and with a quite large Italian community). Then, the policy has been implemented since mid 2009 in most of the French cantons, while most of the German cantons implemented the policy only starting from 2010. This geo-cultural factor in the timing of the policy implementation should not affect the estimated effect of the smoking bans since we control for both canton fixed and canton specific trends. However, such spatial correlation might affect the standard errors of our estimates. To account for this second potential

[6]Carpenter *et al.* (2011) page 49.

source of correlation we use a two-way clustering strategy implemented by Cameron *et al.* (2011). As robustness check, we also use the OLS fixed effects estimator with panel-corrected standard errors (PCSE) based on the work by Beck and Katz (1995). It estimates the parameters by OLS and estimates the standard errors and the variance-covariance matrix assuming that the disturbances are by default heteroskedastic and contemporaneously correlated across panels. We also assumed that errors are first-order auto correlated and that the coefficient of the AR(1) process is specific to each panel.

3.5 Results

This section reports the main results of our analysis. In Section 3.5.1 we show results from the estimation of the effect of the smoking ban on the incidence of AMI hospitalizations using the identification strategy presented in Section 3.4. Section 3.5.2 investigates the mechanisms that lead to the observed reduction in the incidence of hospitalizations after the implementation of the smoking ban. This analysis is conducted using survey data from the SHS and the SHP. Finally, Section 3.5.3 reports the results of a number of robustness checks already discussed in Section 3.4.1.

3.5.1 The effect of smoking ban on AMI

We start showing in Table 3.3 the estimates of the effect of the public smoking ban on the incidence of AMI hospitalizations on total population and by sex.

Each column represents a different specification.

Specifically, Model A includes area and time fixed effects; Model B adds monthly data on area specific unemployment rates; as a further robustness check, Model C includes area specific linear time trends. Estimates are implemented at cantonal (first set of rows) and regional level (second set of rows). Consistent with the graphical evidence in Figures 3.1 and 3.2 (that aggregate data at national level), the table shows robust evidence of a reduction in the incidence of AMI mainly for men. For women, instead, we have evidence of a small and slightly significant effect only in the regional estimates. This result should not be surprising given the very low incidence of AMI among the total women population (see Table 3.2).

The estimated reduction in the incidence of AMI is robust across all specifications. The estimated effect only slightly decreases in size (but not significantly) when we include area specific time trends. This reduction should be due to the effect of imposing the same linear trend before and after the reform in each area (for this reason, from now on we will use Model B as reference). Most importantly, estimates are larger in size—although not significantly— at the regional level. In the total population, it corresponds to a monthly reduction of roughly one

Table 3.3: Effect of public smoking ban on AMI incidence, by sex

	A	B	C
Panel A: cantonal estimates			
Total	-0.869 ***	-0.854 ***	-0.795 **
	(0.317)	(0.328)	(0.346)
Men	-1.489 ***	-1.495 ***	-1.364 ***
	(0.423)	(0.430)	(0.495)
Women	-0.279	-0.252	-0.263
	(0.426)	(0.439)	(0.443)
Obs.	2808	2808	2808
N	26	26	26
Panel B: regional estimates			
Total	-1.231 ***	-1.216 ***	-1.113 ***
	(0.354)	(0.339)	(0.359)
Men	-2.016 ***	-2.020 ***	-1.908 ***
	(0.609)	(0.599)	(0.659)
Women	-0.453 *	-0.459 **	-0.373
	(0.234)	(0.233)	(0.256)
Obs.	11448	11448	11448
N	106	106	106
Area & time f.e.	X	X	X
Monthly unemployment		X	X
Area specific trends			X

Notes: The table shows the coefficient on the effect of the public smoking ban on the incidence of AMI hospitalizations by sex.
Standard errors are robust and clustered at canton and month level (two-way clustering).
Significance levels: ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

AMI case every 100'000 inhabitants (so a reduction of about 1'000 cases per year in Switzerland). Looking at the mean incidence reported in Table 3.2, the estimated coefficients correspond to roughly 8–9% reduction in the incidence of AMI according to the cantonal estimates and 10–12% reduction for the regions. This difference is not surprising since, as aforementioned, cantonal and regional estimates weight the cantons differently. In particular, the cantonal estimates give the same weight to small and big cantons, while in the regional estimates bigger cantons get more weight since they are split in more regions. If the effect is homogeneous across the population the two estimates should converge. However, since most of the smaller cantons implemented only a mild regulation against smoking in public places (see Table 3.1), we might expect a smaller effect on these cantons. This is confirmed by the results provided in Table 3.4, where we show larger effects in cantons that implemented the more severe regulation, and by a further robustness check in which we use population weights for the cantonal estimates.[7] For this reason, in the analysis that follows we only show results for the regions (estimates at cantonal level are always very similar).

Table 3.4: Strictness of the law: mild vs. severe law

	Total	Men	Women
Regional estimates			
Mild	-0.357	-0.552	-0.204
	(0.609)	(1.015)	(0.508)
Severe	-1.356 ***	-2.257 ***	-0.500 **
	(0.323)	(0.535)	(0.207)
Obs.	11448	11448	11448
N	106	106	106

Notes: The table shows the coefficient on the effect of the public smoking ban on the incidence of AMI hospitalizations using two dummies for the ban (severe and mild implementation) using model B as in Table 3.3. Standard errors are robust and clustered at canton and month level (two-way clustering).
Significance levels: ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

Having assessed the impact of the public smoking ban on the total population, we investigate the heterogeneity of our results across age (Table 3.5), income and education (Table 3.6) subgroups. Table 3.5 shows the results for three different age groups, under 50, 50–65 and 66+.

[7]The results, available upon request, are larger than those reported in Table 3.3, and very close to the regional estimates.

Table 3.5: Effect of public smoking ban on AMI incidence, by age and sex

	under 50	50–65	66+
Regional estimates			
Total	-0.386***	-1.879**	-3.599***
	(0.117)	(0.907)	(1.384)
Men	-0.673***	-4.263***	-4.513**
	(0.238)	(1.346)	(2.032)
Women	-0.105	0.495	-2.932**
	(0.075)	(0.836)	(1.412)
Obs.	11448	11448	11448
N	106	106	106

Notes: The table shows the coefficient on the effect of the public smoking ban on the incidence of AMI hospitalizations by age and sex using model B as in Table 3.3. Standard errors are robust and clustered at canton and month level (two-way clustering).
Significance levels: ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

Table 3.6: Heterogeneity by income and education

	Low Income	High Income	Low Educ	High Educ
Regional estimates				
Total	-1.875***	-0.486	-1.470***	-0.945**
	(0.512)	(0.463)	(0.396)	(0.368)
Men	-2.870***	-1.198	-2.809***	-0.822**
	(0.929)	(0.783)	(0.846)	(0.298)
Women	-0.911***	0.201	-0.169	-1.092
	(0.272)	(0.308)	(0.518)	(0.661)
Obs.	5724	5724	5724	5724
N	53	53	53	53

Notes: The table shows the coefficient on the effect of public smoking ban on the incidence of AMI hospitalizations by education and income groups using model B as in Table 3.3. Standard errors are robust and clustered at canton and month level (two-way clustering).

Significance levels: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

In absolute terms, larger effects are estimated in population subgroups more at risk of AMI, namely men in the age group 50–65 and both men and women 66+. More specifically, for men aged 50–65 and 66+ we observe a reduction in the incidence of about 5 cases every 100'000 inhabitants per month. Notice that for women, the effect in the age group more at risk of AMI—women aged 66+—is now large and significant.

In relative terms, the story is quite different since for men the effect is larger for the subgroup under 50 (20% reduction with respect to the mean incidence) and 50–65 (17%) . The difference between absolute and relative terms is due to the difference in the risk of AMI across subgroups but also to the different likelihood of each subgroup to attend bars and restaurants. In the next section, using survey data from the SHP and the SHS, we show that younger people (under 66) are more likely to go to bars and restaurants and were more exposed to ETS before of the policy implementation.

The next step is to investigate whether the public smoking ban had a heterogeneous effect across income and education groups. We argue that this analysis is particularly relevant from a policy perspective since it concerns the potential impact of the policy on health inequality. Indeed, a differential impact of the policy across socio-economic group would exacerbate (in the case of a larger reduction among richer and high educated) or reduce the level of health inequality in the incidence of AMI. Ideally, one should use individual data for this kind of analysis, but the hospital register provides us only geographical information on patients' residence. For this reason, we implement the analysis at the smallest possible level of aggregation, the region. As explained in Section 3.4, this level of aggregation allows us to exploit the within canton variation in the level of income and education, and then control for institutional confounders at cantonal level. In particular, using data on the average median income between 2004 and 2010, we split the regions in two groups high vs. low income. We do the same for education using 2000 census data on education attainment. Table 3.6 clearly shows evidence of heterogeneity across both dimensions, but in particular across income groups. It indicates that the effect of the ban affected mainly low income and low education regions. In these regions, the estimated reduction in AMI incidence is almost twice the average effect estimated in Table 3.3, while the reduction in the richest and more educated regions is in most cases small and not statistically significant. As for the previous evidence, in the next section we show that there was a clear SES gradient in ETS exposure before the policy implementation that is able to explain our results.

Finally, we test whether the effect of the public smoking ban has been larger in cantons that implemented the more severe regulation, as described in Section 3.2. We test this by including in the regression two dummies that capture the effect of the two different regulations, mild and severe. Table 3.4 shows the results by age. The table clearly indicates that the public smoking ban had a significant impact on AMI only in regions where a more severe regulation was implemented. As previously discussed, mild regulation has been implemented mainly in smaller cantons, also charachterized by the absence of metropolitan areas. On the other hand, we might also expect that the public smoking ban had larger effects in metropolitan areas charachterized by a large number of bars and restaurants. Therefore, we avoid speculating on the results reported in Table 3.4, saying that mild regulation are not effective in reducing the incidence of AMI hospitalizations.

3.5.2 How did the smoking ban reduce AMI incidence?

In the previous section we show evidence of a sharp and heterogeneous (by age and SES) reduction in AMI hospitalizations after the implementation of the smoking ban. In this section we want to investigate how this effect takes place. First, we want to figure out whether this effect comes mainly from a reduction of smoking habits (both prevalence and cigarette consumption) and/or a reduction in ETS. Second, we want to understand why the effect was larger among people from low SES. Unfortunately, HOST data does not contain information on patient's smoking behaviors and on important socio-economic characteristics (i.e. income and education). We overcome this problem by using data from two Swiss surveys, namely the SHS and the SHP. As already discussed in Section 3.3, the SHS has been carried out every five years since 1992 and provides important information on population health and health behaviors including smoking and ETS. Using this dataset, Table 3.7 reports smoking prevalence, average cigarette consumption and average (self-reported) minutes of ETS exposure by age and survey year.

In particular, we use the same age categories used in our main analysis and report statistics for the survey years 2007 and 2012, namely the years before and after the implementation of the smoking ban in public places across Switzerland. Consistent with most of the literature on the effect of smoking bans on smoking habits (e.g. Anger *et al.* (2011); Jones *et al.* (2013)), the table shows that between 2007 and 2012 there was no evidence of any reduction in smoking prevalence and in smoking consumption. If ever, we observe a slight increase in consumption among smokers aged 50+. At the same time, however, we observe a dramatic reduction in the self-reported ETS over the 5 years for all age groups. The reduction is particularly large, more than 50%, among people under 50 and people 50–65. The table also shows an age gradient in both smoking prevalence and ETS. In effect, older people, especially those age 66+, smoke less and are less exposed to ETS.[8]

The evidence reported in Table 3.7 is important for our analysis for two main reasons. First, it reveals us that the reduction in AMI hospitalization should be due only to reduction in ETS exposure and not to a reduction in smoking prevalence. Second, the evidence of an age gradient

[8] We checked whether this gradient is driven by the fact that there are more women among the older people. However, even if we split the sample by both age and gender the gradient remains almost unaffected.

Table 3.7: Smoking prevalence, cigarettes consumption and ETS exposure (minutes per day), by age and survey year

	Smokers		Cig. per day		ETS	
	2007	2012	2007	2012	2007	2012
under 50	31.39%	31.24%	9.35	10.00	38.40	16.54
	(0.50)	(0.49)	(0.18)	(0.20)	(1.25)	(0.91)
50–65	26.82%	26.18%	18.85	19.78	25.16	12.19
	(0.66)	(0.66)	(0.32)	(0.31)	(1.27)	(1.066)
66+	11.37%	12.38%	25.14	27.80	9.42	6.07
	(0.51)	(0.55)	(0.42)	(0.40)	(0.72)	(0.65)

Notes: The first set of columns reports the percentage of smokers by age group in the survey years 2007 and 2012; the second set of columns reports the average number of cigarettes smoked per day (conditional on smoking); while the third set of columns reports the average self-reported minutes of ETS exposure per day. Data are extracted from the Swiss Health Survey (2007 and 2012). Sample weights are used to take into account the sample design and non-response.

in ETS exposure (and also in smoking prevalence) is consistent with the evidence reported in the previous section. In particular, the smaller reduction in the incidence AMI hospitalization —in relative terms— among people 66+ seems to be confirmed by the smaller exposure of this age group to ETS. We argue that this is due to the lower probability of this age group to go out in pub, bars and restaurants. We verify this hypothesis by using data from SHP that includes in the questionnaire information regarding leisure activities, including pub, bar and restaurant attendance. Using data from the 2007 wave of the SHP, Table 3.8 shows a clear age gradient also in the likelihood of attending pubs, bars and restaurants.

Table 3.8: Going to bar, pub and restaurant, by age and sex

	Men		Women	
	weekly	monthly	weekly	monthly
under 50	60.37%	89.87%	44.18%	83.38%
	(0.36)	(0.22)	(0.33)	(0.24)
50–65	54.55%	84.13%	40.17%	77.30%
	(0.57)	(0.42)	(0.49)	(0.42)
66+	50.10%	76.58%	34.05%	65.44%
	(0.76)	(0.64)	(0.61)	(0.61)

Notes: This table shows the percentage of respondents that declare to go at least weekly or monthly in bar, pub and restaurant by age and sex. Data are extracted from the 2007 wave of the Swiss Household Panel. Sample weights are used to take into account the sample design and non-response.

Since statistics for 2012 are very similar, we prefer to split the sample also by sex. For both men and women, the share of older people attending these public places at least weekly or monthly is significantly lower than for the younger subgroups (under 50 and 50–65). Unfortunately the survey does not allow to divide the attendance of restaurants from that of pub and bars, where we might expect a larger exposure to ETS and a larger attendance of people under 50.

Finally, using the same approach, we investigate whether the socio-economic gradient in the effect of the policy that we show in Table 3.6 can be explained by different levels of ETS exposure across socio-economic groups. For this reason, in Table 3.9 we report smoking prevalence, cigarette consumption and ETS by education and survey year.

We consider only two main education groups. The high educated group includes people with

Table 3.9: Smoking prevalence, cigarettes consumption and ETS exposure (minutes per day), by education and survey year

	Smokers		Cigarettes per day		ETS	
	2007	2012	2007	2012	2007	2012
Low Educ	28.26%	29.53%	16.38	18.22	35.53	21.83
	(0.41)	(0.41)	(0.25)	(0.26)	(0.92)	(0.80)
High Educ	25.83%	25.47%	16.05	17.43	26.42	12.23
	(0.60)	(0.49)	(0.38)	(0.30)	(1.02)	(0.71)

Notes: The first set of columns reports the percentage of smokers by education in the survey years 2007 and 2012; the second set of columns reports the average number of cigarettes smoked per day (conditional on smoking); while the third set of columns reports the average self-reported minutes of ETS exposure per day. Data are extracted from the Swiss Health Survey (2007 and 2012). Sample weights are used to take into account the sample design and non-response.

at least a high school diploma, while the low educated group includes people with vocational education or below. As expected, smoking prevalence and especially ETS are larger among low educated people in 2007. In effect, the average difference in ETS exposure is roughly 9 minutes. This might explain why we found larger effects of the policy on people from low SES region. Over the 5 years, the reduction in ETS exposure was dramatic for both education groups.

All in all, the results reported in this section are consistent with those reported in the previous section and provide reasonable explanations and mechanisms for the estimated effect of the policy. It is important to stress that even thought the evidence provided in this section is often based on simple bivariate analysis, we are confident that they are not affected by composition effects.[9]

3.5.3 Robustness checks

In this section we implement a set of robustness checks as described in Section 3.4.1. The first set of robustness checks are aimed to test the identification strategy of this paper. First we implement what we call an "unconfoundedness test", in which we test whether the policy affected a set of health outcomes, namely the incidence of lung, prostate and breast cancer, that should not be affected by the policy at least in the short run. The lung cancer affects both men and women, while prostate and breast cancer are sex specific. Even though the lung cancer is

[9]We test it by running a set of multivariate regressions including all the observables characteristics (results are provided upon request).

affected by ETS, we should not be able to observe effects in the short term. This test is meant to verify whether we are confounding the effect of the policy with some other unobservable trend or confounder factors. The test reported in Table 3.10 shows no effect of the public smoking ban on the incidence of all these health outcomes.

Table 3.10: Unconfoundness test: effect of the ban on the incidence of lung, prostate and breast cancer

	Lung cancer	prostate cancer	breast cancer
MS region estimates			
Smoking Ban	-0.002	0.002	-0.020
	(.005)	(.017)	(.043)
Obs.	11448	11448	11448
N	106	106	106

Notes: The table shows the coefficient on the effect of the smoking ban using model B as in Table 3.3. Standard errors are robust and clustered at canton and monthly level (two-way clustering).
Significance levels: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Second, we implement a sort of placebo tests to verify that there are no other jumps in the AMI distribution before or after the real implementation of the policy. As argued by Carpenter *et al.* (2011) this test is particularly suitable to test the endogeneity of the public policy because it allows to control for one–year lead of the smoking ban. Table 3.11 shows the results of this test.

The table clearly shows that there is no evidence of any significant discontinuity in the distribution of the AMI hospitalizations before or after the real implementation.

Another robustness check that indirectly tests our identification strategy is to verify weather the reduction in AMI hospitalizations was larger during the winter season, when people stay more often inside the public place because of the adverse weather conditions. We show the results in Table 3.12, where we report the average effect at the baseline (from April to October), and the interaction with the winter season (from November to March).

As expected, the reduction in AMIs during the winter period is almost twice the reduction we observe during the rest of the year (compare the size of the two coefficients). Again, the results support our identification strategy.

The last robustness check concerns the estimations of the standard errors. As already discussed, to take into account the two potential sources of correlation, within (over time) and

Table 3.11: Placebo test

	Total	Men	Women
Regional estimates			
1 year lag	-.099	.098	-.304
	(.514)	(.902)	(.371)
Smoking Ban	-1.176 **	-1.981 **	-.382
	(.470)	(.839)	(.303)
1 year lead	.604	.722	.488
	(.658)	(.872)	(.504)
Obs.	11448	11448	11448
N	106	106	106

Notes: The table shows the coefficient on the effect of the smoking ban using model B as in Table 3.3 and controlling for the effect of two placebo reforms one year before and after the real implementation. Standard errors are robust and clustered at canton and month level (two-way clustering).
Significance levels: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 3.12: Seasonal heterogeneity

	Total	Men	Women
Regional estimates			
Smoking ban	-.822 **	-1.410 **	-.329
	(.301)	(.578)	(.233)
Smoking ban*winter	-.690	-1.177 **	-.104
	(.428)	(.508)	(.660)
Obs.	11448	11448	11448
N	106	106	106

Notes: The table shows the coefficient on the effect of the smoking ban using model B as in Table 3.3 plus an interaction term for the winter season (November to March). Standard errors are robust and clustered at canton and month level (two-way clustering).
Significance levels: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

between regions, we use a two-way clustering strategy. As robustness check we use an OLS fixed effects estimator with panel–corrected standard errors (PCSE). It estimates the variance-covariance matrix assuming that the disturbances are by default heteroskedastic and contemporaneously correlated across panels, and it allows errors to be autocorrelated (first order autocorrelation specific to each panel). Since this model estimates the full variance-covariance matrix and so a very large number of time parameters (107), we avoid to include a full set of time fixed effects. We substitute them with year and month fixed effects (18 parameters).

Table 3.13 shows that results from PCSE are very similar to those reported in the main text.

Table 3.13: Panel corrected standard errors estimates

	Total	Men	Women
Regional estimates			
Smoking ban	-1.086***	-1.815***	-.380
	(.394)	(.626)	(.385)
Obs.	11448	11448	11448
N	106	106	106

Notes: The table shows the coefficient on the effect of the smoking ban controlling for time and month fixed effects and monthly unemployment rate. The effect is estimated using panel corrected standard errors assuming first order autocorrelation, AR(1). Significance levels: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

3.6 Conclusion

In this paper we assess the causal effect of the introduction of public smoking bans in Switzerland on AMI hospitalizations. We show that this anti-smoking policy implemented in Switzerland was effective in reducing the cases of AMI. Our results are robust to different model specifications and the large battery of tests implemented in this paper never casts doubt on our identification strategy.

On average, the introduction of public smoking bans leads to a reduction of about 10-12% of AMI hospitalization. However, we found large heterogeneity in the effect of interest across sex and age groups. In absolute terms, we found evidence of larger reductions in AMI hospitalizations among men 50+, the subgroup of the population with the largest incidence of AMI. Among women, we found evidence of a significant reduction in AMI hospitalizations only in the age group 66+, being them the only women subgroup really at risk of AMI. In relative terms, the largest reduction (roughly –20%) was observed among men under 50 and 50–65, the subgroups of the population—among those at risk of AMI—that more often should go to bars and restaurant.

We also exploit the spatial heterogeneity in the effect of public smoking bans by focusing our analysis on very small regions (spatial mobility regions), which allows us to hold constant the confounding effect of institutional factors at cantonal level. Such analysis shows that the policy was more effective in poorer and less educated regions and in regions that implement a more severe regulation. The validity of our results is also supported by the available survey data information on smoking prevalence, cigarette consumption and ETS.

Our analysis is of particular relevance from a policy perspective for at least two important reasons. First, from a public health perspective, our paper shows that improving the quality of air in public places through smoking bans has a positive impact on population health. Additionally, our results demonstrate that the impact of the public smoking ban is even more effective in the areas with low income and low education. In these areas the AMIs' reduction is almost twice as large as the average effect.

We also present evidence from survey data that support our results. In particular, we show that younger and less educated people were more affected by ETS before the implementation of the policy and so more likely to be affected by the policy. This is also supported by the large drop in self-reported ETS exposure between 2007 and 2012 based on the data from the SHS. Considering the well-know positive correlation between health, income and education, our results suggest that such a public health intervention might also be able to reduce inequalities in health.

Secondly, this policy leads to a significant reduction of health related costs. In the USA, the American Hearth Association (Go *et al.*, 2014) estimates for 2010 a total cost of 204.4 billions dollars due to hearth diseases. Even in Switzerland, the second country in the world for life expectancy at birth (OECD, 2013), AMI is the leading cause of death and it represents a huge economic burden. For instance, Wieser *et al.* (2012) estimate that in 2008 the total direct costs of acute coronary syndrome (of which AMIs represent more than 80% of the costs) were more than 1 billion Swiss Francs. To this large cost we should add 500 millions of indirect costs due to production losses. Therefore a reduction of 10–12% of AMI hospitalizations means saving at least 150 millions per year.

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