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**Working Paper**

## How Health Plan Enrollees Value Prices Relative to Supplemental Benefits and Service Quality

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# RUHR

ECONOMIC PAPERS

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## How Health Plan Enrollees Value Prices Relative to Supplemental Benefits and Service Quality

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Christian Bünnings, Hendrik Schmitz, Harald Tauchmann,  
and Nicolas R. Ziebarth<sup>1</sup>

## How Health Plan Enrollees Value Prices Relative to Supplemental Benefits and Service Quality

### Abstract

*This paper empirically assesses the relative role of health plan prices, service quality and optional benefits in the decision to choose a health plan. We link representative German SOEP panel data from 2007 to 2010 to (i) health plan service quality indicators, (ii) measures of voluntary benefit provision on top of federally mandated benefits, and (iii) health plan prices for almost all German health plans. Mixed logit models incorporate a total of 1,700 health plan choices with more than 50 choice sets for each individual. The findings suggest that, compared to prices, health plan service quality and supplemental benefits play a minor role in making a health plan choice.*

*JEL Classification: D12, H51, I11, I13, I18*

*Keywords: Service quality; non-essential benefits; prices; health plan switching; German sickness funds; SOEP*

*March 2015*

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<sup>1</sup> Christian Bünnings, UDE and CINCH; Hendrik Schmitz, Universität Paderborn, CINCH and RWI; Harald Tauchmann, Universität Erlangen-Nürnberg; Nicolas R. Ziebarth, Cornell University, USA. – We would like to thank Normann Lorenz and participants of the iHEA 2014 in Dublin for valuable comments and suggestions. Moreover, we would like to thank Thomas Adolph for providing us with the data on health plan characteristics and Philip Susser for editing this work. Financial support from the BMBF (Förderkenzeichen 01EH102A) is gratefully acknowledged. We take responsibility for all remaining errors in and shortcomings of the article. – All correspondence to: Hendrik Schmitz, University of Paderborn, Department of Economics, Warburger Str. 100, 33098 Paderborn, Germany, e-mail: hendrik.schmitz@uni-paderborn.de

# 1 Introduction

Health plan literacy has become a buzzword in the current health care debate. It refers to the provision of extensive information and education on health care and health plans, in addition to a set of available choices. Theory postulates that a high degree of health care literacy leads to behavioral changes—e.g., health plan switching—which would make the market more efficient and improve quality. Increasing consumer health literacy is generally seen as a promising road to gradually improving the efficiency and quality of health care systems around the world. Yet it remains controversial to what extent consumers are already using available information when making important choices, for example in choosing health plans.

Health plan choice essentially depends on the factors *(i)* price—typically a non-linear trade-off between premiums and cost-sharing amounts—, *(ii)* benefits covered, *(iii)* clinical health care quality—e.g., via provider networks or managed care—as well as the *(iv)* service quality of the insurer. Empirically estimating the impact of these four factors is challenging, particularly in the US setting where we observe a fragmented health care landscape with hundreds of thousands of different health plan parameters. Most US employees are limited to a choice of two or three plans—mostly being HMO or PPO—making it challenging to disentangle the generalizable impact of single determinants. That being said, most empirical studies on health plan choice determinants exploit the US setting (Dowd and Feldman, 1994; Cutler and Reber, 1998; Royalty and Solomon, 1999; Strombom et al., 2002; Atherly et al., 2004; Buchmueller, 2006; Buchmueller et al., 2013). In single payer markets such as Canada or the UK, people do not have any choice with regard to their health plans. This limitation explains the absence of empirical studies on the determinants of health plan choice for these countries.

This study focuses on the German case which is, for several reasons, a particularly interesting one to analyze. The German statutory health insurance (SHI) represents a “third way” between government run single payer systems without any choice and the US, where health care is pre-dominantly offered through less regulated private entities. Although the US system has moved towards an increasingly regulated system under the Affordable Care Act (ACA) (or *ObamaCare*), the German market is still more heavily regulated and standardized. However, Germany combines this heavy regulation with a relatively high degree of

health plan choice. In the German SHI, about 130 sickness funds (=health plans<sup>1</sup>) compete for mostly mandatorily insured customers. Most of these health plans operate nationwide although several are solely offered in some of the 16 German states. An interesting feature of the German health care market is that, unlike the US, managed care is legally prohibited. Furthermore, selective contracting does not exist. This implies the absence of provider networks and the uniformity of reimbursement rates leads to uniformity of clinical health care quality across all 130 health plans. Providers do not know or care about patients' SHI sickness funds, which eliminates the relevance of health plan determinant (iii) above—variation in health care quality. German social legislation also prohibits deductibles and coinsurance rates and only allows small copayments for inpatient and outpatient care. Those small copayments for inpatient and outpatient care do not vary across plans either.<sup>2</sup> This regulation shuts down the non-linear trade-off between premiums and cost-sharing in factor (i) above. Finally, German social legislation establishes a very generous “essential benefit package” similar to the one under the ACA in the US. Essentially all medically necessary inpatient and outpatient treatments are covered.<sup>3</sup> However, sickness funds may “voluntarily” offer the coverage of additional benefits such as alternative treatments or immunizations for tropical diseases to differentiate their product.

This study empirically exploits the standardization and the extensive health plan choice set in the German market. We link representative enrollee panel data to publicly available health plan prices, as well as standardized health plan quality information, and exploit changes in these health plan characteristics across 130 plans and over 4 years. We exploit standardized supply-side information from a well-respected private company that consistently surveys and ranks all German health plans. Thus, our empirical approach exploits the same standardized supply side information that German consumers can access in online portals and magazine rankings in order to select health plans.

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<sup>1</sup>We use the terms “health insurance (company),” “sickness fund,” and “health plan” as interchangeably.

<sup>2</sup>For the time period under consideration, the copayments were €10 (\$13) per day for a hospital day as well as €10 per calendar quarter for outpatient visits. Total cost-sharing is capped at 2% of the annual income, for chronically ill at 1%.

<sup>3</sup>As in other countries, the coverage of dental care and eyeglasses is limited.



This paper estimates the impact of the three health plan choice parameters (i) price, (ii) “non-essential” supplemental benefits, and (iii) service quality on the decision to select health plans. *Service quality* is mostly defined by health plan accessibility (via physical branches, hotlines or the internet) and the quality of information provided to customers looking for help. While a substantial body of empirical literature analyzes the price impact of health plan choice<sup>4</sup> (e.g. Strombom et al., 2002; Atherly et al., 2004; Schut and Hassink, 2002; Buchmueller, 2006; Tamm et al., 2007; Frank and Lamiraud, 2009; Buchmueller et al., 2013; Schmitz and Ziebarth, 2011; Wuppermann et al., 2014) there exist only a few studies on the role of benefits and quality.

Using employer data from General Motors and accounting for health plan fixed effects, Scanlon et al. (2002) estimate changes in health plan market shares due to the introduction of quality report cards. They observe that employees avoid subscribing to health plans with below average ratings. Chernew et al. (2008) use the same data and apply a Bayesian learning model to show that only 3% of enrollees switch health plans due to report cards. Estimating a cross-sectional conditional logit model on health plan choices of Harvard employees, Beaulieu (2002) finds a positive relationship between higher quality ratings and the probability of health plan choice. And exploiting data on federal US employees, Wedig and Tai-Seale (2002) use a nested multinomial logit model to show how these report cards increase price elasticity. Harris (2002), in contrast, conduct a discrete choice experiment in West Los Angeles and conclude that large quality differences would be required for consumers to accept provider access restrictions. Dafny and Dranove (2008) analyze the response of federal retirees to public quality ratings while controlling for market-based learning and find that both public and nonpublic information play a modest role in health plan decision making.<sup>5</sup> Finally, Abraham et al. (2006) do not find that information about higher-quality alternatives affects switching behavior.

This paper is one of the rare studies on the determinants of health plan choice—in particular when taking into account health plan benefits and services—outside the US. Linking

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<sup>4</sup>Overviews are provided by Kolstad and Chernew (2009) and Gaynor and Town (2012).

<sup>5</sup>This is in line with Jin and Sorensen (2006) who exploit public and nonpublic health plan ratings and find evidence that both influence individuals’ decisions but only moderately.

representative individual-level health plan switching information from the German Socio-Economic Panel Study (SOEP) to detailed objective health plan data from 2007 to 2010, this paper investigates the relative roles of prices, non-essential benefits, and service quality. As discussed, by construction, the German institutional framework eliminates important confounding channels such as additional non-linear variation in cost-sharing dimensions or differences in provider networks and reimbursement rates and thus, perhaps most importantly, health care quality (Ziebarth, 2012; Gruber and McKnight, 2014). A major strength of our analysis is that we are able to reproduce an almost complete picture of each enrollee's SHI health plan choice set and are not restricted to single regions, employers or certain subgroups of the population. The empirical specifications employ mixed logit models that take heterogeneity in individual preferences as well as unobserved health plan characteristics into account. Our findings show a significantly negative price effect on health plan choice but no indication that supplementary benefits and service quality play an important role in the decision to choose health plans. Heterogeneity analyses with respect to individuals' age, gender and health status indicate that only modest effect heterogeneity exists in this market.

The remainder of the paper is organized as follows: The next section covers the institutional details of the German health insurance market. Section 3 outlines the empirical specification and section 4 presents the data used for estimation. The estimation results are presented in section 5. Section 6 concludes.

## **2 Institutional Background**

The German health insurance system is characterized by the coexistence of statutory health insurance (SHI) and substitutive private health insurance (PHI). This paper focuses on the SHI, which covers roughly 90% of the population most of whom are compulsorily insured. Insurance under the SHI is mandatory for employees with gross wage earnings below a defined threshold (in 2014: €53,550 per year). Nonworking spouses and dependent children under 25 years are covered at no additional costs by SHI family insurance. Further regulations also apply to specific groups of the population, such as students and the unemployed,

although most of them are covered by SHI. High-income employees, self-employed individuals and civil servants may opt out of the SHI and buy substitutive PHI or stay under the SHI as voluntary members. Currently the SHI market comprises around 130 not-for-profit health insurance companies, also called “sickness funds”, roughly half of which are operating nationwide, while the remaining ones solely operate in some federal states. Switching sickness funds is uncomplicated: the minimum contract period is 18 months and there is no enrollment period; guaranteed issue exists and several specific search engine websites help consumers to compare and switch health plans.<sup>6</sup> Yet, health plan switching is a rare event among SHI enrollees. In a given year only about 5% of all SHI insured switch health plans (Schmitz and Ziebarth, 2011).

About 95% of the SHI benefit package is predetermined by social legislation at the federal level. The federally mandated minimum benefit package is very generous relative to international standards, basically including all medically necessary treatments in addition to prescription drugs, birth control, preventive and rehabilitation care as well as rest cures (cf. Ziebarth (2010a)). Albeit more generous, this minimum benefit package is comparable to the new Essential Health Benefits under the ACA. However, German social legislation additionally heavily restricts cost-sharing such that only small copayments exist that are identical across health plans. Yet, to differentiate their product and attract enrollees, sickness funds have the opportunity to voluntarily offer additional benefits, which are not part of standard package under the SHI. These **optional supplemental benefits** can be subdivided into (i) alternative medicine and (ii) further supplementary benefits.

*Alternative medicine* covers complementary treatments such as ayurveda, homeopathy, osteopathy, and urine therapy.<sup>7</sup> Although the effectiveness of alternative medicine is discussed controversially, demand for such treatments seems to be increasing. Along with these alternative medical treatments, sickness funds may also offer conventional “non-essential”

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<sup>6</sup>The minimum contract period for those who are subscribed in special health plans (“optional tariffs”) ranges from one to three years. If sickness funds plan to increase prices, insurees have—independent of the enrollment length—an extraordinary right to cancel the contract and switch funds.

<sup>7</sup>More precisely, sickness funds are allowed to provide alternative medical treatments only if they fulfill the efficiency principle. According to the German Social Code Book V (paragraph 12), treatments fulfill this principle if they are sufficient as well as medically and economically appropriate, which is, however, a rather vague legal concept.

medical treatments. Examples for these *supplemental benefits* could be preventive check-ups (e.g. the “J2” check-up for adolescents) and certain types of immunizations (e.g. malaria prophylaxis). Typically these benefit differences result only in very small expenditure differences, e.g., a single combined vaccination shot against diphtheria and typhoid fever costs about €15. However, supplemental benefits may also comprise more expensive medical treatments, such as additional subsidies for in-vitro fertilizations. However, these more expensive supplemental benefits usually solely apply to a very small group of enrollees.

**Health plan premiums** are calculated in form of social insurance contributions. To calculate the employee share of the premium, a sickness fund specific contribution rate is applied to the gross wage, including all fringe benefits, up to a defined contribution ceiling (in 2014: 48,600 € per year). One half of the contribution rate is formally paid by the employee and the other half by the employer. In January 2009 and as part of a health policy reform (*GKV-Wettbewerbsstärkungsgesetz*), SHI financing was reorganized. Prior to January 2009, health insurance premiums were a function of gross wage earnings and the contribution rate. The latter was set independently by each sickness fund, resulting in a variety of contribution rates, ranging from 12.2 to 16.9% of individual’s gross wage earnings in 2008. The reform equalized the contribution rates to 15.5% across all health plans.<sup>8</sup> After 2009, if allocated revenues from the 15.5% standardized contribution rate did not cover the health plan’s expenses, sickness funds had to charge an additional premium in form of an absolute monthly Euro amount from their members. If allocated revenues exceeded expenses, sickness funds could pay out a bonus to their members. Hence, post reform, price differences were expressed in absolute rather than relative terms, which increased switching behavior significantly (Schmitz and Ziebarth, 2011; Wuppermann et al., 2014). However, we convert all monthly health plan premiums for each enrollee into euro amounts. Moreover, Section 5.4.1 looks at potential influences of the reform on the sickness fund choice behavior.

Apart from price and benefit differentiation, sickness funds compete on **service quality**. We define service quality as the general accessibility and the quality of information provided to enrollees. Most sickness funds operate a network of physical branches but also offer hotline

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<sup>8</sup>Due to the financial crisis, the contribution rate was temporarily decreased to 14.9% from July 2009 to December 2010. Since January 2011, the initial contribution rate of 15.5 percent applies.

services. Running a large number of branches may be preferable to (some) members—e.g., the elderly—but also implies higher operational costs. In order to reduce administrative expenditures, some sickness funds reduced their branch network significantly over time. A minority of sickness funds do not run any physically accessible branch but are exclusively available by telephone or the internet (“Direktversicherer”). Improved accessibility by phone or the internet was also enforced by the health care reform of 2000 (*GKV-Gesundheitsreformgesetz*). According to this legislation, all sickness funds had to improve their service and consulting. As a result, health plans started to operate different types of hotlines. While some were general, aimed to help insurance members with questions related to membership issues, others provided more detailed information, such as information about drugs and their side effects.

### 3 Empirical Specification

Following the contemporary literature that investigates the effect of health plan’s characteristics on health plan choice (e.g. Beaulieu, 2002; Wedig and Tai-Seale, 2002; Jin and Sorensen, 2006; Dafny and Dranove, 2008), we apply discrete choice methods. More precisely, we opt for a random parameters model (Revelt and Train, 1998; McFadden and Train, 2000). The random parameters model (RPL), also called mixed logit model, is a generalization of the conditional logit model (McFadden, 1973) and has two important advantages over the traditional conditional logit that makes it especially attractive in the present analysis.

First, several studies (e.g., Beaulieu, 2002) provide indications for the presence of heterogeneity in preferences with respect to health plan characteristics. Considering alternative medicine, for instance, preferences for such treatments are likely heterogeneously distributed across the population. While some individuals may value such treatments very highly, others may not care about them at all. The mixed logit model explicitly allows for introducing heterogeneity in consumer preferences by modeling the preference parameters as random variables.

Second, the RPL does not rely on the restrictive independence of irrelevant alternatives (IIA) assumption. The IIA assumption requires that the odds between two alternatives do

not depend on which other alternatives are available. As individual choice sets consist of a large number of alternatives (health plans) that can be considered as close substitutes, the IIA assumption is at least questionable.<sup>9</sup> We specify the linear index, measuring individual  $i$ 's inclination to choose health plan  $j$  as

$$\gamma_i' Premium_{ij} + \delta_i' Benefits_j + \zeta_i' Service_j + \alpha_j + \varepsilon_{ij} \quad (1)$$

$Premium_{ij}$  is the monthly health insurance premium in Euro. The vectors  $Benefits_j$  and  $Service_j$  includes measures for additional benefits and service quality characteristics, which are covered in more detail in the subsequent section. Since the premium is income-dependent—unlike  $Benefits_j$  and  $Service_j$ — $Premium_{ij}$  varies across individuals and sickness funds. To account for time- and individual-invariant unobservable health plan characteristics that might be correlated with our explanatory variables, we include a set of sickness fund fixed effects ( $\alpha_j$ ). Essentially, we assume that the unobserved part of utility consists of a sickness fund-specific fixed effect and a random error term.<sup>10</sup> If, however, time varying unobservable health plan factors exist that are correlated with the characteristics under scrutiny, endogeneity would still be an issue in our empirical analysis. We assume the  $\varepsilon_{ij}$  to be iid and to follow a type I extreme value distribution and, thus, arrive at the familiar conditional logit model.

The coefficient vectors  $\gamma_i$ ,  $\delta_i$  and  $\zeta_i$  comprise the preference parameters of interest. As indicated by the subscript  $i$ , these preferences are allowed to vary across individuals but are assumed to be constant for the same individual over time. We choose the most common distributional form for the coefficients and assume that individual preferences are normally

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<sup>9</sup>The IIA assumption is arguably less problematic for studies in the U.S., as choice sets in employer-sponsored settings typically do not comprise more than five alternatives. However, Wedig and Tai-Seale (2002) restrict their estimation sample to choice sets with five or fewer alternatives to make the IIA assumption more plausible. Moreover, Harris (2002) find statistically significant evidence that preference heterogeneity exists, suggesting that the IIA assumption is violated.

<sup>10</sup>This is similar to the approach adopted by Chernew et al. (2004). In general, one could allow the alternative-specific fixed effects to vary across individuals, similar to the other explanatory variables. Due to the large number of alternatives in our choice sets, this would require to estimate distribution parameters of more than 100 additional random variables. Given the already high dimensional optimization problem, this approach is not feasible in the present analysis.

distributed.<sup>11</sup> Since the likelihood function has no closed-form solution, we apply maximum simulated likelihood methods to estimate the parameters.<sup>12</sup>

Finally, although we have substantial information on individual socio-economic characteristics, we do not directly include them in the model. Considering that they are alternative invariant, including these variables would require interacting them with each alternative in the respective choice set (since we are estimating the relationship between alternative varying characteristics on health plan choice). Yet, individual choice sets are quite large (ranging from 41 to 73 alternatives) in our analysis. An approach that considers socioeconomic characteristics would inflate the number of coefficients enormously and would be impractical for computational reasons.

## 4 Data

We make use of two different data sources, which are described in more detail below, and combine them to a dataset that mirrors in detail the health plan choice sets of the SHI insured in Germany.

### 4.1 Individual Level Data

Individual-level data is taken from the German Socio-Economic Panel Study (SOEP). The SOEP is a representative longitudinal survey that started in 1984 and collects annual information on both the household and the individual level. Currently, the SOEP comprises more than 20,000 individuals from more than 10,000 households (Wagner et al., 2007). We use the waves 2008 to 2011.

The estimation sample exploits information on enrollees' current sickness fund as well as their insurance status. First, we exclude those individuals who are covered by PHI. Second, we restrict the sample to sickness fund enrollees, not the total number of insured. The latter would also include family members insured at no cost under SHI family insurance. We

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<sup>11</sup>We assume a diagonal variance-covariance matrix of the coefficients, and hence uncorrelated coefficients.

<sup>12</sup>We use the Add-On package `mixlogit` for Stata (Hole, 2007). Estimation results are based on 50 Halton draws. Any data or computational errors are our own.

focus on the paying members so we have exactly one observation per health plan choice decision. Paying members are the insurance holder, gainfully employed and earn more than €400 gross per month.<sup>13</sup> Third, to ensure that the empirical results are not driven by the high degree of state dependence—as only 5.3% of all individuals switch their health plans—we focus on sickness fund switchers and those who opt out of the family insurance to become a paying member. The intuition behind the latter is that these individuals are likely to inform themselves carefully about their new health plan options. Restricting the sample to ‘switchers’ is similar to estimating a model that also includes non-switchers and explicitly controls for state dependence. However, due to the large number of alternatives in individuals’ choice sets, estimating a mixed logit model that accounts for state dependence is both computationally expensive and instable. Since, in both approaches, the effects are identified by those who switch health plans, we opt for the simpler approach. The final estimation sample consists of 1,726 choices from 1,594 different individuals (that is, about 100 individuals change their sickness fund twice in the observation period).

To construct individual choice sets, we exploit information on the enrollees’ state of residence. Since the majority of, but not all, sickness funds operate nationwide it ensures that only relevant health plans enter an individual’s choice set. As can be seen in Panel A of Table 1, the choice sets include on average 58 sickness funds, ranging from 41 to 73 health plan alternatives. Choice sets are smallest for individuals living in Mecklenburg-Western Pomerania, ranging from 41 different health plans in 2011 to 53 in 2008. Individuals residing in North Rhine-Westphalia have the largest choice sets, ranging from 55 (2011) to 73 (2008).<sup>14</sup> Overall, the empirical identification relies on 400 to 500 observed health plan decisions per year, where each individual has 50 to 60 health plans to choose from. Thus, we annually observe about 25,000 potential options, totaling 100,000 options over the four years under consideration.

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<sup>13</sup>This excludes all those insured under SHI family insurance, the unemployed for some of whom social security pays the health insurance premium, full-time students who just pay an income-independent flat premium (2014: €64,77 per month) or who are insured under their parents’ family insurance, pensioners as well as special population groups, such as draft soldiers or low-income earners.

<sup>14</sup>Due to several mergers of sickness funds, the number of active health plans is decreasing over time.



Table 1: Descriptive Statistics

Panel A: Sample Characteristics	2008	2009	2010	2011	Total	
Observations (=choice sets)	25,920	25,476	27,352	21,140	99,888	
Choice Occasions	388	447	477	414	1,726	
Switchers	231	293	318	285	1,127	
Exit Family Insurance	157	154	159	129	599	
# Choice Sets per individual						
Mean	66.8	57.0	57.3	51.1	57.9	
S.D.	5.4	4.4	4.7	3.9	7.1	
Min.	53.0	48.0	48.0	41.0	41.0	
Max.	73.0	63.0	63.0	55.0	73.0	
Panel B: Individual Characteristics	Mean	S.D.	Med.	Min.	Max.	Obs.
Self-Assessed Health	2.34	0.84	2	1	5	1,724
Very Good	0.13	0.34	0	0	1	1,724
Good	0.49	0.50	0	0	1	1,724
Satisfactory	0.29	0.45	0	0	1	1,724
Poor	0.08	0.27	0	0	1	1,724
Bad	0.01	0.10	0	0	1	1,724
Age	37.03	12.63	37	18	80	1,726
Female	0.55	0.50	1	0	1	1,726
Monthly Gross Income [EUR]	1,959	1,413	1,700	400	12,885	1,726
Panel C: Sickness Fund Characteristics	Mean	S.D.	Med.	Min.	Max.	Obs.
Premium [EUR] <sup>+</sup>	152.98	8.04	158	133	178	323
Branch Network	0.00	1.00	-0.15	-2.42	2.48	323
Access by Phone	0.00	1.00	0.28	-2.95	1.39	323
Alternative Medicine	0.00	1.00	-0.17	-2.16	3.16	323
Supplemental Benefits	0.00	1.00	0.06	-4.73	1.93	323

Sources: Müller and Lange (2010); Lange (2011), German Federal (Social) Insurance Office, National Association of Statutory Health Insurance Funds, annual reports of the sickness funds, information by sickness funds. SOEP.v28. Authors' calculation based on the SOEP. <sup>+</sup> denotes fictive absolute EUR premiums based on a monthly gross income of €2,000 for the sake of illustration. The reason is that the premium depends on three parts: the sickness fund contribution rate, the individual income, and the sickness fund add-on premium (after 2009). In order to show average Euro amounts on the sickness fund level, we a hypothetical monthly gross income of €2,000. In the regression models, exact premiums are calculated based on enrollee and sickness fund level information.

SOEP interviews are typically carried out in the first quarter of the year, while sickness fund characteristics were collected at the end of a calendar year in November and December. Thus we link the respondents' health plan choice sets at the time of the interview in the first months of a year with the health plan information as provided at the end of the prior calendar year. Hence, we make use of switching and health plan data for the years 2007 to 2010.

We use individual's wage earnings reported in the SOEP to calculate health plan premiums for all potential choices in Euro amounts at each point in time. In addition we make use of information on age, gender and self-reported health status (SAH) later in a subsection on heterogeneity by demographic groups. As shown in Panel B of Table 1, slightly more than half the sample is female and the average age is 37 years. Thirteen percent self-rate their health as 'very good' and 49% as 'good'.

## 4.2 Health Plan Level Data

Information on the health plan characteristics (contribution rate, optional benefits and service quality) is provided by a private company (*Kassensuche GmbH*). It is collected through questionnaires that are sent out annually to all active sickness funds. Sickness funds have a strong incentive to participate in the survey, as *Kassensuche GmbH* operates a large German web portal where consumers can compare a broad range of characteristics of existing sickness funds. Moreover, at the end of each year, a popular weekly business magazine (*Focus Money*) publishes a detailed overview and ranking of the best 50 health plans as surveyed by *Kassensuche GmbH*. This health plan ranking is comprised of sub-scores for several subcategories, measured on continuous scales, which provide the basis for the benefit and service quality characteristics used in our regression model.<sup>15</sup> Since this is the same information that consumers obtain, directly exploiting this information is one main advantage of our approach. The main drawback of using this data is that interpretation of the computed scores is not straightforward as we discuss below. To account for slight differences in the calculation of the sub-scores over time, the regression models make use of the z-transformed sub-scores.<sup>16</sup> In to-

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<sup>15</sup>The number of subcategories has slightly changed over time, therefore we use only those scores which were part of the survey in each of the four years.

<sup>16</sup>The z-transformation is conducted for each year separately.

tal, we have information on 115 different sickness funds covering the years 2007 to 2010. The health plans included every year have a total market share of around 80% and also represent 80% of all existing plans (Müller and Lange, 2010).

**Price information** for each sickness fund is mainly based on pre-2008 publicly available health plan-specific contribution rates. Post-2008, we additionally take sickness fund specific add-on premiums and refunds into account (see Section 2). We link this publicly available information to both, individuals' gross income and the federally fixed yearly contribution ceiling, in order to calculate the exact monthly health plan premiums (in euros) that each enrollee pays. We do not consider the employer share, which is legally fixed at 50% of the total premium.<sup>17</sup> We disregard the employers' share since employees typically believe that they solely pay the employee share as premium and are thus very likely to make decisions solely based on their share. As seen in Panel C of Table 1, the share of the monthly premium that individuals carry—based on average monthly gross earnings of €2,000—ranges between €133 and €178, with a mean value of €153.

**Optional supplemental benefits** cover additional health care services that are not part of the standard SHI benefit package as defined by the federal regulator, such as immunizations for tropical diseases or preventive screenings for breast or skin cancer in younger ages. Moreover—in being more generous—sickness funds can deviate from the mandated minimum benefit package for domestic help and rooming-in. For example, federal legislation requires sickness funds to pay for domestic help for children aged 12 and younger with parents who are institutionalized for medical treatment (and no other household member or relative is available). More generous plans extend this coverage to children aged 14 and younger.

The optional supplemental benefits that sickness funds provide in addition to the federally mandated benefits (*Kassenleistungen*) enter the empirical model in terms of two variables (sub-scores): (i) *alternative medicine* and (ii) *supplemental benefits*. The score of *alternative medicine* is mainly based on the number of different alternative medical treatments offered

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<sup>17</sup>Effective July 1, 2005, the strict equal sharing of contributions was altered. Between 2005 and 2015, the employees' share was  $[0.9 + 0.5 \times (cr - 0.9)]$  percent of their gross wage up to the contribution ceiling, where  $cr$  denotes the overall contribution rate. In the example above, this amounts to an employee share of 7.45 percent and an employer share of 6.55 percent of the gross wage. If the incidence of the full health insurance contribution was fully on employees, this would simply reduce the price coefficients.

by each sickness fund. Sickness funds are not entirely free to offer any additional treatment, but may choose from a list of around 20 approved treatments (e.g. ayurveda or homeopathy). Additionally, the score takes into account whether these treatments are restricted to certain regions or physicians. More restrictions lead, *ceteris paribus*, to a lower score. Table Panel C of Table 1 provides some key descriptive statistics on *alternative medicine* and the other scores. While the z-transformation renders the mean uninformative, the negative median value indicates that the distribution of *alternative medicine* is skewed to the right.

**Health plan service quality** is measured by two variables: (i) *branch network* and (ii) *accessibility by phone*. The score of *branch network* measures the density of the branch network and takes into account that roughly half of all sickness funds are only active in certain federal states and, hence, are likely to have fewer branches than those who operate nationwide. More precisely, the original score before the z-transformation is derived from the log of the total number of branches divided by the number of federal states in which the sickness fund operates. *Accessibility by phone* considers the different types of available hotlines (medical, non-medical) and how many hours these hotlines are staffed. Differences in staff quality—i.e., the share of staff with special qualifications, such as social insurance clerks (“*Sozialversicherungsfachangestellte*”), physicians, nurses or pharmacist—are accounted for by weighting the hotline’s operating hours accordingly, where higher staff quality receives a higher weight.

The joint distribution of health plan characteristics exhibits positive correlations between all considered variables, ranging from 0.102 (‘supplemental benefits’ and ‘branch network’) to 0.549 (‘alternative medicine’ and ‘access by phone’). This indicates that sickness funds seem to position themselves either at the high or low benefit and service plan end in the market, rather than trying to build a specific reputation by offering very specific extra benefits or boosting specific quality indicators. This corresponds with the observation that the scores for service quality and additional benefits are positively correlated with prices; the correlations range from 0.053 (supplemental benefits) to 0.228 (branch network). Hence high benefit and service plans are on average the more expensive plans.

The positive correlation of premiums and service quality/extra benefits seems to indicate that sickness funds are aware of the quality-price trade off when they differentiate their product. Another interesting descriptive result highlights the relevance of health plan literacy: Considering the five plan characteristics above, the majority of plans in the market are dominated by at least one competitor. That is, there is at least one competitor that is better in all five characteristics. This holds for every single year with the shares of dominated plans ranging from 0.67 (2008) to 0.90 (2010).

## 5 Results

### 5.1 Descriptives

To obtain a first impression of whether prices, optional benefits, or service quality matter in an individual's decision to choose a health plan, we compare means of sickness fund characteristics of new and old plans for switchers. The results are based on 729 switches with information on both, the new and the old health plan.<sup>18</sup> As can be seen in Table 2, average health insurance premiums are lower for the new health plan chosen as compared to the old health plan. The average difference is €3.56 per month, indicating that price is a relevant determinant in health plan choice. For service quality characteristics, we do not observe significant differences between the new and the old health plans. The same applies to supplemental benefits, where we find no significant differences either for alternative medicine or for optional benefits. The quality differences between old and new plans are very small and below 0.05. Overall, the descriptive results provide first evidence that premium differences may have an effect on switching behavior while optional benefits and service quality seem to play a minor role in health plan choice.

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<sup>18</sup>Note that for the main analysis we only need information on the new health plan which is available for more individuals.

Table 2: New and Old Health Plan Characteristics for Switchers

Characteristic	$\bar{X}^{new}$	$\bar{X}^{old}$	$\Delta$	$s.e._{\Delta}$
Premium	162.47	166.04	-3.56***	(0.30)
Branch Network	0.91	0.88	0.04	(0.05)
Access by Phone	0.75	0.77	-0.02	(0.03)
Alternative Medicine	0.88	0.85	0.03	(0.05)
Supplementary Benefits	0.37	0.34	0.03	(0.04)

*Notes:* Authors' calculation. The table shows means of sickness fund characteristics for new and old health plans. The calculations are based on 729 observations. Standard errors of the differences in means are in parentheses. \*\*\*  $p < 0.01$ ; \*\*  $p < 0.05$ ; \*  $p < 0.1$

## 5.2 Main Results from Mixed Logit Models

Table 3 shows two different specifications of the model outlined in Section 3. It reports the estimated means and standard deviations of the random coefficients. Model 1 solely includes health plan premiums, optional benefits, and service quality, while Model 2 adds a set of sickness fund fixed effects.

According to Model 1, individuals prefer sickness funds with lower insurance premiums, as indicated by the significantly negative mean coefficient of *premium*. The corresponding estimated standard deviation is close to zero and not significant, suggesting that the relationship between premiums and individual sickness fund choice is largely homogeneous across individuals.

With respect to service quality, we observe significantly positive mean coefficients for both, *branch network* and *accessibility by phone*. In addition, the estimated standard deviations point towards considerable heterogeneity. Considering health plan benefits, the results suggest that alternative medical treatments are not significantly associated with insurance choice. Both the estimated mean and the standard deviation are close to zero and not significantly different from zero. In contrast, the estimated mean parameter of *supplemental benefits* exhibits the expected positive sign and is highly significant.

Since our models impose that the coefficients follow a normal distribution, we can use the estimated mean and standard deviation to calculate the share of individuals that place a positive value on optional supplemental benefits (Train, 2009). The corresponding probability is given by  $P(X > 0) = 1 - F_X(0) = 1 - \Phi(0 - 0.433/0.755) \approx 0.717$ , where  $\Phi$  represents the

Table 3: Main Estimation Results—The Role of Price, Non-Essential Benefits, and Service Quality in the Health Plan Choice Decision

	Model 1		Model 2	
	Mean	SD	Mean	SD
Price				
Premium	-0.092***(0.006)	0.003 (0.016)	-0.061***(0.008)	0.002 (0.014)
Service Quality				
Branch Network	1.600***(0.052)	0.774***(0.065)	0.063 (0.110)	0.354*** (0.128)
Access by Phone	0.964***(0.108)	0.618***(0.119)	0.052 (0.152)	0.210 (0.335)
Benefits				
Alt. Medicine	-0.044 (0.041)	0.039 (0.135)	0.069 (0.071)	0.371*** (0.108)
Suppl. Benefits	0.433***(0.055)	0.755***(0.071)	0.058 (0.064)	0.182 (0.150)
SF Specific Cons.	no		yes	
# Observations	99,888		99,888	
# Choice Occasions	1,726		1,726	

Notes: Authors' calculation. The table shows estimated coefficients of both, mean and standard deviation of the random parameters. Estimated standard errors are in parentheses. \*\*\*  $p < 0.01$ ; \*\*  $p < 0.05$ ; \*  $p < 0.1$

cumulative distribution function of the standard normal distribution. The calculation would suggest that, according to Model 1, around 70% of all enrollees place a positive value on optional supplemental benefits.

It is likely, however, that the variables on the sickness fund level capture other unobserved health plan characteristics, such as brand loyalty, awareness, or the general reputation of the fund. Therefore, Model 2 adds a full set of sickness fund fixed effects. As for the impact of *premium*, the estimated mean coefficient and standard deviation are close to what we find in Model 1 suggesting a minor role of unobserved confounding factors. Although slightly smaller in absolute magnitude, the estimated mean coefficient exhibits the expected negative sign and remains highly significant. As in Model 1, the estimated standard deviation is virtually zero, indicating a homogeneous relationship between premiums and health plan choice across individuals. Virtually everyone in the sample places a negative value on higher prices.

In contrast, when considering time-invariant health plan level effects, all mean coefficients of service quality and optional benefits dramatically shrink in size and become insignificant. In absolute terms, the magnitudes of these coefficients fall within the same range as the estimated *premium* coefficient. However, not only are the standard errors much larger, the variables are also measured on different scales, which needs to be taken into account when in-

terpreting the results. Comparing the effect of an increase by one standard deviation—which is about eight for the premium and one for the other variables (cf. Table 1, Panel C)—the response to a price increase is eightfold the response to an increase in service quality or optional benefits. Thus, we conclude that including health plan fixed effects is essential here, and that this is our preferred specification.

Next, while the standard deviations of the coefficients for *accessibility by phone* and *supplemental benefits* are small and insignificant, those of *branch network* and *alternative medicine* suggest significant heterogeneity in preferences for these factors. Around 40 to 60% of all enrollees either place a positive or negative value on these health plan characteristics (using the formula above). While some individuals seem to prefer more benefits and a better service, others actually want less. Individuals who do not need or want to physically visit their sickness fund branches may think that a large *branch network* would be a waste of money. Moreover, given that the usefulness of alternative medicines is not proven, the same argument may hold for *alternative medicine*.

Overall, the estimation results are in line with the simple descriptive results and intuition. They suggest that differences in health plan premiums are the main determinant in sickness fund choice. Medically non-necessary supplemental health plan benefits and service quality play, on average, a negligible role in consumers' health plan choice decisions. One explanation for this finding would be that quantitative premium differences are easy to understand and process for most people. The monetary trade-off to service quality parameters is much more abstract and may only become apparent when customers actually need help (Schram and Sonnemans, 2011). Research on Medicare Part D has also demonstrated that the insured do not always enroll in the optimal health plan and that they learn to improve their selection over time (Heiss et al., 2006, 2013; Abaluck and Gruber, 2011; Ketcham et al., 2012). Another explanation for our finding may be a low awareness of health plan differences in terms of service quality and optional benefits.



### 5.3 Quantifying the Tradeoff between Prices, Service Quality, and Optional Benefits

We now have a closer look on the relative importance of the different health plan characteristics for consumer health plan choice. We can exploit this information in any discrete choice model—which is based on a linear index as in equation (1)—the ratios of the coefficients represent marginal rates of substitution (*MRS*).<sup>19</sup> We focus on how consumers value health plan differences in supplemental benefits and service quality, compared to premium differences. More specifically, we are interested in  $-\delta_i^k/\gamma_i$  and  $-\zeta_i^k/\gamma_i$  that indicate at which rate enrollee  $i$  is willing to trade-off additional benefits or better quality against a premium decrease, where  $k$  indexes benefit and quality characteristics. In other words, if sickness fund  $j$  increases the premium by one unit, the probability the plan being chosen by enrollee  $i$  does not change if, at the same time, its phone accessibility improves by  $-\zeta_i^{\text{access}}/\gamma_i$  units. Interpreting these ratios, however, requires scales that measure a one unit change. As no natural scale is available for optional benefits and service quality, we define a unit as one standard deviation of the relevant characteristic in the sample distribution. Similarly, we define one premium unit. Yet, unlike for the service and quality measures, an increase in the monthly premium by one standard deviation can also be expressed as an absolute €8 increase (cf. Table 1, Panel C). This value has some intuitive appeal, as it represents the typical premium differential between health plans after 2008. For health plans that increased prices, €8 is also close to the average price increase in 2007 and 2008 (Schmitz and Ziebarth, 2011).

Mixed-logit estimation does not yield estimates for  $-\delta_i^k/\gamma_i$  and  $-\zeta_i^k/\gamma_i$  at the individual level and, for this reason, does not allow for the calculation of individual marginal rates of substitution. What one obtains from the estimation are normally distributed population parameter estimates  $\mu_k$  and  $\sigma_k$ . As  $MRS_k$  is a ratio of normal random variables, its distribution involves a Cauchy-component rendering the mean (and higher-order moments) undefined which cannot consistently be estimated (cf. Marsaglia, 1965; Cohen Freue, 2007). Hence, we have to base the discussion on quantiles rather than the mean of  $MRS_k$ . In particular, we focus on the median of  $MRS_k$ . Rather than directly interpreting  $-\hat{\mu}_k/\hat{\mu}_{\text{premium}}$  as *ML*-estimate for the median

<sup>19</sup>One may just as well refer to this ratio as ‘rate of substitution’ since it does not change due to linearity.

of  $MRS_k$ , we simulate the percentiles of the  $MRS$ -distributions—along with corresponding 95%-confidence bands—on basis of the results for Model 2 (cf. Table 3, right panel).<sup>20</sup>

### The Density of the Branch Network vs. a Lower Premium

Let us start with  $MRS_{branch}$ . The simulated median of  $MRS_{branch}$  is 0.128, indicating that the median enrollee trades-off lower premiums against a higher density network of branches at a rather small rate. More precisely, an increase in the branch network by one standard deviation (S.D.) is just valued as one-eighth S.D. of the premium, which translates to €1 per month. However, the simulated 95%-confidence interval of  $[-0.324, 0.564]$  indicates that this value is imprecisely estimated.<sup>21</sup> Nevertheless, even the upper confidence bound is only 0.56 and lets us exclude—with 95% statistical certainty—that consumers would trade an increase in the branch density network by one S.D. for more than €8.50 per month.

This picture somewhat changes when we consider the estimated heterogeneity in the  $MRS$ . Although the exact shape of the estimated  $MRS$  distribution depends heavily on distributional assumptions (and should be interpreted with caution), assessing other quantiles may provide insights in the variation of enrollees preferences. At the 95<sup>th</sup> percentile, the rate of substitution is more than ten times larger than at the median (point estimate 1.307). This means that, according to our estimates, those five percent of enrollees who have the strongest preferences for face-to-face services and a high branch density are willing to accept a 1.3 S.D. increase in the monthly premium (€10) for a one S.D. increase in the branch density. However, this number carries a lot of uncertainty (confidence interval:  $[0.286, 3.064]$ ). On the other hand, according to the estimated distribution of  $MRS_{branch}$ , 42% of enrollees would not be willing to accept any increase in premium in exchange for more physical branches. Taking sampling

<sup>20</sup>The reason is that the ratio of the means  $\mu_k/\mu_l$  does not provide an accurate approximation of the median of the corresponding ratio distribution—if the denominator distribution has a mean close to zero and a non-vanishing density at zero. To estimate the percentiles, we draw 2 million random numbers from the relevant ratio distributions, with the point estimates  $\hat{\mu}_k$ ,  $\hat{\mu}_{premium}$ ,  $\hat{\sigma}_k$ , and  $\hat{\sigma}_{premium}$  entering the involved normal distributions, and then average the simulated quantiles over 2,000 replications. Due to the large size of the pseudo sample, the estimated percentiles exhibit very little sampling variability and averaging has almost no effect. To simulate the confidence bands, we also sample 2,000 times where, in each replication, the four relevant parameters are drawn from the (estimated) jointly-normal distribution of the  $ML$ -estimator.

<sup>21</sup>Using  $-\hat{\mu}_{branch}/\hat{\mu}_{premium}$  directly as estimate for  $\text{med}(MRS_{branch})$  and applying the delta-method for calculating confidence intervals yields results (point estimate: 0.128, confidence interval:  $[-0.306, 0.561]$ ) that just marginally deviate from the simulation-based counterpart. This can be explained by the small value of  $\hat{\sigma}_{premium}$  that lets the density of the denominator almost vanish at zero.

error into account, one cannot even reject a number as high as 77% with 0.95% certainty. All in all, a large fraction of German enrollees find a sickness fund’s branch network as playing a marginal—if any—role in the health plan choice when compared to the premium. Yet, due to large heterogeneity, a small fraction of enrollees seem to value personal customer-to-customer services a great deal.

### Telephone Access and Supplemental Benefits vs. a Lower Premium

Turning to the remaining optional benefit and service characteristics, the pattern of estimated *MRSs* is similar to what we found and discussed for ‘branch network’ above. The estimated median *MRSs* range from  $1/7$  to  $1/10$  (access by phone: 0.106; alternative medicine: 0.140; supplemental benefits: 0.118), indicating a rather low median willingness to pay for optional supplemental benefits like immunization shots, alternative medicine, and service quality. The point estimates, however, are accompanied by rather wide confidence intervals. At the 95<sup>th</sup> percentile, however, the estimated *MRSs* (access by phone: 0.808; alternative medicine: 1.379; supplemental benefits: 0.723) are 6 to 10 times larger than the median. About 40% of all enrollees are not willing to accept any premium increase in return for more optional benefits or a better service.

When individuals decide to enroll in a health plan, its premium will be compared to *all* of the services and benefits offered, which has not been considered by the discussion above that focused on single services and benefits. Thus we finally analyze how consumers would trade-off a simultaneous increase by one S.D. for all four characteristics. That is, we consider  $-(\delta_i^{branch} + \delta_i^{access} + \zeta_i^{alt. med.} + \zeta_i^{sup. benefits})/\gamma_i$ , for which we obtain estimates as for any *MRS<sub>k</sub>* above. This calculation reflects an estimated median value of 0.491 (confidence interval: [-0.181, 1.181]). Although—not surprisingly—the median willingness to pay for joint improvements in quality and benefits exceeds the median willingness to pay for single improvements, the value of the point estimate is still less than one. Like the *MRSs* for particular quality characteristics, the joint rate of substitution exhibits considerable heterogeneity. At the 95<sup>th</sup> percentile, the rate is 2.438, i.e., five times larger than at the median. However, considering the other tail of

the distribution, 33% of the insured are not willing to accept any premium increase even if all considered services and benefits would simultaneously improve by one unit.

## 5.4 Heterogeneity Analysis

### 5.4.1 The Reform of 2009

As mentioned in Section 2, our sample period covers a price setting reform that became effective in January 2009. Prior to 2009, price differences between sickness funds were expressed in contribution rate differences. After 2009, contribution rates were fixed across all funds. Furthermore, price differences between health plans were expressed as flat euro add-on premium or refund. Schmitz and Ziebarth (2011) and Wuppermann et al. (2014) find substantial effects of this reform on health plan switching behavior.

*A-priori*, it is unclear whether differences in non-essential benefits and service quality became more or less salient after this price framing reform. On the one hand, because health price differences became more salient post reform, price differences could have increased in importance relative to other plan characteristics. On the other hand, because the market price dispersion decreased after 2009, a decrease in price variation across funds could have made quality differences more relevant to consumers.

To test whether the impact of prices, optional supplemental benefits, and service quality structurally changed post-reform, we split our sample into a pre- (2008/2009) and post-reform (2010/2011) period.<sup>22</sup> As can be seen in Table 4, the estimated mean coefficient of *premium* is about twice as large post-reform, which is in line with Schmitz and Ziebarth (2011) and Wuppermann et al. (2014). The estimated coefficients for *alternative medicine* and *supplemental benefits* are almost identical in pre and post reform years. All in all, it seems to be justified to not distinguish by the pre- and post-reform time periods in the main specification. This conclusion is also supported by an LR-test that fails ( $p$ -value 0.599) to reject the null hypothesis of equal distributions.

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<sup>22</sup>We opt for assigning the wave 2009 to the pre-reform period, as SOEP interviews are typically carried out at the beginning of a year and, hence, switching most likely refers to the previous year.

Table 4: Heterogeneity I—Reform 2009

	Mean		SD	
<b>Price</b>				
Premium × pre	−0.056***	(0.009)	0.001	(0.014)
Premium × post	−0.092***	(0.031)	0.069	(0.042)
<b>Service Quality</b>				
Branch Network × pre	−0.010	(0.119)	0.276	(0.200)
Branch Network × post	0.203	(0.139)	0.475***	(0.167)
Access by Phone × pre	0.152	(0.189)	0.428*	(0.223)
Access by Phone × post	−0.008	(0.171)	0.125	(0.341)
<b>Benefits</b>				
Alt. Medicine × pre	0.073	(0.115)	0.260	(0.168)
Alt. Medicine × post	0.093	(0.107)	0.651***	(0.177)
Suppl. Benefits × pre	0.010	(0.068)	0.012	(0.180)
Suppl. Benefits × post	0.076	(0.102)	0.136	(0.167)
<b>SF Specific Constants</b>			yes	
# Observations	99,888			
# Choice Occasions	1,726			

Notes: Authors' calculation. The table shows estimated coefficients of both, mean and standard deviation. The binary indicator "pre" is one for the SOEP waves 2008 and 2009, while "post" is one for the waves of 2010 and 2011. Estimated standard errors are in parentheses. \*\*\*  $p < 0.01$ ; \*\*  $p < 0.05$ ; \*  $p < 0.1$

## Health and the Potential for Cream Skimming

The German social health insurance system combines guaranteed issue with income-dependent contribution rates. Individual risk rating is strictly prohibited. This regulation, however, creates an incentive for sickness funds to engage in active or passive health risk selection. To minimize this incentive, a comprehensive risk adjustment scheme exists. The scheme is based on age, gender, a reduced earnings capacity, and 80 chronic illnesses, such as diabetes or cancer. However, because the risk adjustment scheme does not perfectly adjust for health risks, allocated revenues for high risk enrollees may still be smaller than actual costs. Still, it is unclear to what extent risk selection exists in the German market. While Bauhoff (2012) finds some evidence for direct risk selection based on the state of residence of the insured—which is found to be very small, however, in quantitative terms—Nuscheler and Knaus (2005) find no indication for risk selection in the German SHI.

Sickness fund diversity is important for sufficient competition in the market. If health plans are not allowed to diversify their products at all, enrollees will not have the ability or an incentive to search for and switch to the best health plans in the market. On the other hand, such diversity could also be exploited for indirect risk selection. For example, there is the notion that better educated individuals are more likely to choose health plans that include alternative medicine. Offering alternative medicine could therefore be a mechanism for sickness funds to attract lower risks.

This paper cannot directly test whether cream skimming exists in the German SHI. However, we indirectly assess the *potential* for indirect risk selection by analyzing heterogeneous responses to differences in prices, service quality and optional benefits. Risk selection strategies could be employed if healthy and unhealthy individuals were attracted differently by different health plan characteristics.

To classify individuals into different risk types, we use the self-assessed health (SAH) measure, age, and gender. Even though the German risk-adjustment is based on health, age, and sex, we argue that this simplistic approach is useful for several reasons. First, SAH is a more comprehensive measure and includes more information than the 80 illnesses considered in the risk equalization scheme. Second, SAH also includes more up-to-date information as enrollees' illnesses of the previous year enter the risk adjustment formula, but not current ones. Third, despite its simplicity, SAH has been shown to be an excellent predictor of true health (McGee et al., 1999). Issues related to reporting heterogeneity seem to be mostly limited to age and gender (Ziebarth, 2010b). Fourth, even though age and sex are included in the risk-adjustment formula, these indicators are not perfect risk adjusters. We stratify the results based on these easy-to-observe socio-demographics since they likely carry other correlated important health information. Also, gender and age can be easily observed and then considered by insurers for active or passive risk selection.

We cannot include the variables in the most flexible way for computational reasons. Estimation time and tractability of the model and stability of the estimation results are the main factors behind this restriction. Therefore, for each of the variables, we construct a mutually exclusive subset of two dichotomous indicators  $G^1$  (group 1) and  $G^2$  (group 2) that represent

Table 5: Heterogeneity II—Risk Types

	Age ( $G^1$ : age < 50)				Gender ( $G^1$ : males)				SAH ( $G^1$ : SAH < 3)			
	Mean		SD		Mean		SD		Mean		SD	
Price												
Premium $\times G^1$	-0.063***	(0.009)	0.001	(0.015)	-0.054***	(0.010)	0.015	(0.028)	-0.059***	(0.009)	0.005	(0.020)
Premium $\times G^2$	-0.053***	(0.013)	0.000	(0.028)	-0.070***	(0.010)	0.004	(0.015)	-0.069***	(0.011)	0.002	(0.022)
Service Quality												
Branches $\times G^1$	0.056	(0.112)	0.350**	(0.145)	0.083	(0.122)	0.317*	(0.177)	0.041	(0.119)	0.302*	(0.183)
Branches $\times G^2$	0.004	(0.134)	0.154	(0.398)	-0.008	(0.112)	0.264	(0.192)	0.054	(0.124)	0.422***	(0.147)
Phone $\times G^1$	0.187	(0.154)	0.442***	(0.158)	0.128	(0.167)	0.384**	(0.178)	0.307*	(0.166)	0.560**	(0.155)
Phone $\times G^2$	-0.069	(0.172)	0.153	(0.277)	0.236	(0.174)	0.519**	(0.172)	-0.098	(0.149)	0.114	(0.329)
Benefits												
Alt. Med. $\times G^1$	0.064	(0.074)	0.325***	(0.123)	0.058	(0.087)	0.523**	(0.105)	0.023	(0.080)	0.499**	(0.112)
Alt. Med. $\times G^2$	0.021	(0.114)	0.272	(0.262)	0.064	(0.074)	0.030	(0.219)	0.170*	(0.090)	0.164	(0.222)
Benefits $\times G^1$	0.060	(0.065)	0.143	(0.158)	0.121	(0.077)	0.178	(0.156)	0.083	(0.069)	0.150	(0.142)
Benefits $\times G^2$	0.072	(0.114)	0.357**	(0.216)	0.042	(0.077)	0.319**	(0.141)	0.058	(0.089)	0.375**	(0.168)
SF Specific Cons.	yes				yes				yes			
# Observations	99,888				99,888				99,790			
# Choice Occasions	1,726				1,726				1,724			

Notes: Authors' calculation. The table shows estimated coefficients of both, the mean and the standard deviation of the random parameters.  $G^1$  and  $G^2$  denote good health risks (age < 50; males, SAH < 3) and bad health risks (age  $\geq$  50; females, SAH  $\geq$  3), respectively. Estimated standard errors are in parentheses. \*\*\*  $p < 0.01$ ; \*\*  $p < 0.05$ ; \*  $p < 0.1$ . Significant differences between the distributions of estimated preference parameters are highlighted in gray.

different health risks. SAH is reported on a five point scale, ranging from 1 (very good) to 5 (bad). We require “good health risks” to report at least good health (SAH category 1 or 2), while bad health risks fall into categories 3 to 5. Age is collapsed into two binary variables marking individuals aged younger than 50 ( $G^1$ ) and older than 50 ( $G^2$ ). We also use separate indicators for males ( $G^1$ ) and females ( $G^2$ ); females use more health care and have higher expenditures.

We run one regression for each of the three separating variables and interact health plan characteristics with the corresponding indicators for both groups. As we have imposed a normal distribution for the random coefficients, conducting a joint test on equal coefficients (mean and S.D.) allows us to test whether the distributions of preferences differ significantly between high and low risks. This is essentially the approach adopted by Beaulieu (2002) and Wang et al. (2011), who run conditional logit and mixed logit models on different subsamples and compare the distributions of the estimated parameters across subsamples. Recall that we cannot include baseline levels of socio-economic controls that do not vary over the choice sets for each enrollees (see Section 3).

Table 5 presents the results of the heterogeneity analysis with respect to age, gender, and SAH. Significant differences between the distributions of the preference parameters are highlighted in gray. First, we compare the share of enrollees who place a positive value on certain health plan characteristics. Using age as a separating variable, we find no significant differences between the distributions of taste parameters.

With respect to health plan premium, the estimated distributional parameters are close to what we observe in the main specification without interaction terms (cf. Model 2 in Table 3). No differences are found when we stratify by age, gender, or health status. This means that the sick and the healthy, the young and the old, as well as males and females all value health plan prices in a similar fashion. The same holds for the factor *branch network*.

With respect to *accessibility by phone*, the null hypothesis is rejected when we stratify by gender and health status ( $p$ -values: 0.0059/0.0179). While the estimated shares of enrollees who value a good hotline service is almost equal in the gender specification ( $G^1$ : 63%,  $G^2$ : 67%), the difference seems to be substantial in the SAH specification ( $G^1$  (good health): 71%,  $G^2$  (bad health): 19%). This indicates that good health risks value good phone accessibility *more* than bad health risks.

The opposite holds for *alternative medicine*. Again, the hypothesis of equal distributions for gender and health is rejected ( $p$ -values: 0.0823 / 0.0309), but the share of those who value these treatments is much larger among bad health risks in both specifications (SAH:  $G^1$ : 52%,  $G^2$ : 85%; also gender:  $G^1$  (men): 54%,  $G^2$  (women): 98%).

Finally, with respect to *supplemental benefits*, the null hypothesis of equal distributions is likewise rejected for gender and health ( $p$ -values: 0.0823 / 0.0309), indicating that those in good health value supplemental benefits more than bad health risks (SAH:  $G^1$ : 71%,  $G^2$ : 56%; also gender:  $G^1$  (men): 75%,  $G^2$  (women): 55%).

In total, preference differences between age, gender and health status are rather small. When they exist, they do not point into one clear direction. Thus, we conclude that—in the current German public health insurance—supplemental benefits and services do not seem to be powerful tools for indirect cherry picking because different risk types do not seem to systematically react to special health plan features.



## 6 Conclusion

This paper exploits a unique institutional setting and linked individual and health plan level data to assess the relative roles of prices, non-essential benefits, and service quality in the decision to choose health plans. Individuals' health plan choices are modeled using a random parameters model which accounts for health plan heterogeneity and time-invariant unobserved factors. In total, the empirical setting exploits 1,724 health plan choices and almost 100,000 potential choice sets between 2007 and 2010.

We find that prices play the dominant role in the decision to choose health plans and only see limited effects of the provision of non-essential benefits and service quality. In quantitative terms, for the median enrollee, a one S.D. increase in any of the non-price factors evaluated (density of branches, telephone access, alternative medicine, and other optional supplemental benefits) is offset by a one-eighth S.D. decrease in premiums, or €1 per month. In other words, even when service quality and non-essential benefits play a role in the decision to choose health plans, enrollees are willing to trade them against lower premiums at a rather small rate. However, we find that up to 70% of all enrollees do not value these non-price attributes at all. On the other hand, we find that heterogeneity in consumer valuation of non-price health plan attributes is very large and a minority of enrollees may value service quality a great deal.

These findings hold in an institutional setting where differences in service quality and optional benefits should, in principle, be more salient than in other markets—due to a heavy federal regulation including standardization of the benefit package and cost-sharing parameters. Yet, the minor relevance of service quality and additional benefits could also be a result of heavy federal regulations and standardized benefits. On the other hand, online portals—that cover the entire German market and provide information on standardized non-price attributes—facilitates the comparison and switching of plans. Our empirical approach is based on exactly this standardized supply-side information that consumers use to make their health plan choice.

According to standard economic theory, absent adverse and risk selection, allowing health plans to diversify more would unambiguously increase consumer choice and welfare. Our

empirical findings, which suggest heterogeneity in consumer valuation of non-price attributes underscores this notion. However, recent research in behavioral economics has challenged the rationality assumption and provides evidence for phenomena such as decision overload, an occurrence that stems from complex multidimensional choice sets. With regard to Germany's heavily regulated market, allowing insurers to diversify their product to a greater extent could imply (i) selective contracting and the formation of provider networks, (ii) more leeway to vary cost-sharing amounts, or (iii) more leeway to exclude benefits from the very generous federally mandated benefit package.

As in every empirical study, the strict interpretation of our results is limited to the specific setting, in our case the German market. However, we believe that the findings are of broader relevance, particularly since ours is one of the first studies exploiting and disentangling the role of the two factors "non-essential benefits" and "service quality" relative to prices. One promising opportunity for future research could address the ability of consumers to cognitively process the information provided to them and transmit the information into behavioral action.

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