

A Physician Fee that Applies to Acute but not to Preventive Care: Evidence from a German Deductible Program

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Abstract

A certain German sickness fund offers € 240 per year to its clients if they pay the first € 300 of their health care bills, except for physician visits, for which a flat rate of € 20 applies. This paper studies the effects of this deductible scheme on health care demand by comparing about 5,000 participants with a control sample, using the sickness fund's claims data covering in-patient care, prescription drugs, and ambulatory care. The data extend to three years: the year of the start of the program and the two years preceding this. We apply a parametric approach that models the choice of the deductible program, the probability of positive expenses as well as the demand for health care services, conditional on demand being positive. Instruments for the participation decision are used, and the results are compared with those of an exogenous specification of the program choice. The physician fee appears to significantly decrease the number of visits as well as the expenses for curative care. By contrast, prevention activities, not subject to the co-payment, remain constant.

Zusammenfassung

Eine deutsche gesetzliche Krankenkasse offerierte im Rahmen eines Modellvorhabens ihren freiwillig Versicherten für 240 € einen Selbstbehaltvertrag, wenn diese sich bereit erklärten, für Arztbesuche 20 € zu zahlen und die Kosten der Innanspruchnahme anderer Leistungen bis zu einem Gesamtbetrag von 300 € selbst zu übernehmen. In der vorliegenden Arbeit untersuchen wir die Auswirkungen dieses Anreizvertrages, indem wir die Leistungsanspruchnahme von rund 5.000 Teilnehmern des Modellvorhabens mit jener einer Kontrollgruppe vergleichen. Die Individualdaten decken zwei

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Jahre vor und das Jahr nach der Einführung des Modellvorhabens ab. Wir wählen einen parametrischen Ansatz zur Untersuchung der Determinanten, sich für den Selbstbehaltvertrag zu entscheiden, mindestens einmal zum Arzt zu gehen sowie der Höhe der Leistungsinanspruchnahme. Wir vergleichen einen Instrumentvariablenansatz mit einer exogenen Spezifikation der Selbstbehaltwahl. Die Resultate zeigen, dass der Selbstbehalt die Zahl der Arztbesuche und die Höhe von kurativen medizinischen Leistungen signifikant reduziert. Umgekehrt bleibt die Inanspruchnahme von präventiven Leistungen, die nicht dem Selbstbehalt unterliegen, konstant.

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1. Introduction

Germany is undertaking small steps toward significant co-payments in social health insurance. The 2004 health care reform introduced a one-off user charge of € 10 per quarter for physician visits, creating an incentive to avoid unnecessary visits. The latest reform, which became effective on April 1, 2007, endows sickness funds with the right to offer their clients co-payments and preferred provider policies. These and further incentive schemes have been designed to curb the ever-rising health care expenditure in German social health insurance.

This paper studies a single co-payment contract of the sickness fund ‘Techniker Krankenkasse’ that was approved by the German Federal Ministry of Health as a test program starting in 2003. This contract offers a payment of € 240 at the beginning of a year if an individual accepts a deductible of € 300. Expenditures for in-patient care and drugs are fully charged to the deductible while, with respect to ambulatory care, an enrollee pays a € 20 flat fee per physician visit. The program is restricted to so-called voluntary statutory health insurees, i.e. individuals who remain under social health insurance despite their right to opt out and choose private insurance. In the first year, about 10,000 persons enrolled in the deductible program.

We compare health care demand between the program and a control group in the years 2001, 2002, and 2003. We use the demand in the two years preceding the program start to explain the enrollees’ contract choice in 2003 and apply a parametric model that uses instruments to identify the effect of the deductible on health care demand. The focus is on the € 20 co-payment for physician visits. After analyzing the decision to participate in the deductible program, we study the decision to visit a physician at least once in 2003. Finally, we analyze the number of physician visits in 2003, conditional on being positive.

Special emphasis is given to the difference between acute care and preventive care as the latter is exempted from the deductible. While we expect the

deductible to decrease curative health care demand, we might see an increase in the demand for preventive care due to its exemption and given that physicians have some discretionary choice on the set of services provided to a patient seeking preventive care.

This paper is organized as follows: Section 2 describes the deductible program in more detail. Section 3 presents the data and section 4 describes the econometric methodology, while section 5 discusses the choice of instruments. Section 6 gives the results, and section 7 concludes.

2. The Deductible Program TK 240

Germany has a two-tiered health insurance system. Social health insurance applies to employees and their families whose earnings do not exceed € 3,450 a month. Employees with higher earnings as well as self-employed and civil servants have the right to opt out of social health insurance and to buy private insurance instead. Moreover, students, retirees, and welfare recipients may also have the status of a voluntary insuree according to social insurance regulations and for other reasons. About 90 percent of the German population is covered by social health insurance.

A substantial number of the 11 million persons eligible for private insurance remain within social health insurance and contribute an average 14.3 percent of their earnings to this end. Abstracting from altruistic motives, there are two main reasons for not choosing private insurance. First, both a spouse not participating in the work force and children receive free coverage under the breadwinner's social health insurance policy. Second, a possibly unfavorable health status could lead to higher payments since private health insurance charges risk-equivalent premiums. Currently, there are about 5.1 million employees voluntarily socially health insured. Private health insurers and sickness funds compete fiercely for the coverage of these so-called voluntarily paying members since they and their families on average incur relatively low health care expenditures. Sickness funds are particularly interested in covering these families because a risk-adjustment scheme which controls for age and gender (only) and provides funds to compensate for the lack of contributions of family members applies in social health insurance. Thus, sickness funds expect a profit with members who are voluntarily insured within this system. Under these circumstances, a deductible program is a profitable option for the sickness fund both to make it attractive for members to remain within social health insurance and to gain new clients. From a social health insurance perspective, a deductible is beneficial if it reduces the extent of moral hazard.

The deductible program of the Techniker Krankenkasse (TK) is called 'TK 240 program' as it pays out a bonus of € 240 at the beginning of a year. The deductible is € 300 and applies to all medical services. With respect to

ambulatory care, a flat € 20 fee per visit applies irrespective of the actual cost of treatment. Physician services are reimbursed by the physicians' association, which receives its financial endowment from the sickness funds based on capitation. This payment system implies that the sickness fund has no precise information about an enrollee's actual demand for ambulatory care, which explains why the TK 240 program depends on the participants' declaration of the number and purpose of their physician visits during the year. Enrollees are likely to reveal this information truthfully since the sickness fund can deduce the patients' number of consultations from their physicians' drug prescriptions.

The deductible program TK 240 differentiates according to the age of the co-insured family members. The health expenditure of the spouse and of grown-up children is charged to the deductible while the expenditure of children below the age of 18 is not. A second differentiation regards the nature of the care, curative or preventive, that is sought by the enrollee. If an enrollee visits the physician for preventive care, such as a screening service or a consultation for contraception, the co-payment does not apply.

The German Ministry of Health approved of the program in 2002 and, based on a report by the sickness fund on the progress of the program, renews its approval every year. With the new health reform law in place since April 1, 2007, approval is no longer necessary and the sickness fund is free to offer the program to all its clients.

3. Data

10,155 persons with an additional 5525 family members enrolled in the first year of the program. The control group consists of 12,891 persons (plus an additional 14,586 family members) randomly drawn from the sickness fund's approximately 1 million voluntary insurees. An exception to this concerned the regional distribution, which was adjusted to that of the program participants sample. The focus of this paper is on the 4744 persons for whom we have comprehensive information on the demand for health care (for the results regarding the insurees with incomplete demand data, see Felder and Werblow, 2006). Individual drug and in-patient demand data come from the sickness fund, while data on ambulatory care (number of physician visits, expenses for curative and preventive care) stem from three physicians' associations (North-Rhine, Hamburg and Schleswig-Holstein).

Table 1 shows the descriptive statistics of the main socioeconomic variables and expenses in the year 2003 of the program participants as well as of the control group. Health care utilization is much lower among participants. They see a physician less often and spend less on drugs and in-patient care. While they spend one third on curative care, their spending on preventive care is half as high as that of non-participants. In the two years preceding the program,

expenses on health care were also much lower for those that enrolled for the 2003 start, indicating a rational choice. The participants are on average six years younger than members of the control group and the program's share of women is 3 percentage points lower than the control group's.

Table 1
Descriptive statistics

	Non-participants		Participants	
	Mean	Std. dev.	Mean	Std. dev.
Expenses (€)				
Total 2001	674	(1979)	241	(952)
Total 2002	786	(2459)	173	(479)
Total 2003	871		191	
Curative care in 2003	261	(714)	88	(162)
Preventive care in 2003	15	(44)	8.80	(30)
Hospital care in 2003	376	(2593)	75	(642)
Drug use in 2003	219	(902)	19	(70)
Consultations in 2003	11	(14)	4	(6)
Age	48	(10)	42	(7)
Share of women (%)	18	—	15	—
Share of self-employed (%)	18	—	14	—
Monthly earnings (€)	2641	(2100)	3245	(1855)
Severity status (%)	4	—	1	—
Score to go to private health insurance	19	(25)	37	(30)
Share of not urban (%)	17	—	14	—
N of months insured	12	(0)	9	(4)
N of individuals	2737		2037	
Breadwinner's family				
Share with co-insured (%)	54	—	28	—
Share with adult co-insured (> 18) (%)	40	—	11	—
N of adults ^{a)}	1.27	(0.56)	1.11	(0.38)
Age of adults ^{a)}	41	(13)	35	(11)
Share of men ^{a)} (%)	15	—	16	—
Share with children (< 18) (%)	34	—	24	(43)
N of children ^{a)}	1.89	(0.82)	1.87	(0.73)
Age of children ^{a)}	10	(4)	10	(4)
Share of boys ^{a)} (%)	48	—	49	—

^{a)} conditional on being positive.

The individuals' earnings are measured using their earnings that are liable for health insurance contributions. This figure is provided by the employer if a person is employed (the average earnings of an employee in the sample are € 4141). The self-employed declare their own earnings, which can vary be-

tween zero and the threshold for compulsory insurance (€ 3,450). As these declarations may not be accurate, we set the earnings of the self-employed to zero, as we do for other groups of non-employed persons (students, welfare recipients, retirees, and others). Altogether, the resulting average earnings of the voluntary insurees are below the earnings threshold for compulsory insurance. The participants on average earned an additional € 600 per month.

There were four times more non-participants with a severity status than participants. The variable “Score to go to private health insurance” is an index calculated by the sickness fund based on a client’s characteristics. The score is much higher for participants, which points to the fact that the deductible program is tailored to clients with a high probability of leaving social insurance. The inclusion of ‘number of months insured’ is necessary since insurees could enroll in the program anytime during the first year.

With respect to the characteristics of the breadwinner’s family, the most striking difference between participants and non-participants is the much lower share of adult co-insured family members among participants. This can be explained by the design of the deductible: the expenses of a breadwinner’s adult co-insured family member are charged to the deductible. Compared to a single person, a breadwinner with an adult co-insured faces a much higher risk of losing the € 300 rebate and is less likely to participate in the program.

4. Methods

The empirical analysis of the deductible program TK 240 faces two challenges, the endogeneity of the program choice and the distribution characteristics of the dependent variables (physician visits and medical demand).

Regarding endogeneity, consider the linear demand equation for the year 2003 with y as the medical expenses, D as a dummy for an individual’s choice of the deductible ($D = 1$ if yes, $D = 0$ if otherwise) and X as a vector of additional explanatory variables,

$$(1) \quad y = X'\beta + \delta \cdot D + \varepsilon,$$

where the error term ε captures the influence of non-observable factors. The coefficient δ will not correctly measure the effect of the deductible if D is correlated with the error term ε . In this case, the decision to participate in the program is endogenous. In order to identify the selection effect, we first use a probit model to estimate the probability that an individual chooses the deductible, $\Pr(D = 1)$. Socioeconomic factors, household composition, and demand for health care in the years 2001 and 2002 feature as explanatory variables Z . In addition, we take into account unobservable components and measurement errors in ϖ :

$$\begin{aligned}
 (2) \quad \Pr(D = 1) &= \Pr(Z'\alpha + \varpi > 0) \\
 &= \Pr(\varpi < Z'\alpha) \\
 &= \Phi(Z'\alpha),
 \end{aligned}$$

where Φ is the cumulative distribution function of the standard normal distribution. In general, one assumes a bivariate normal distribution for the two error terms with covariance matrix $\Sigma = \begin{bmatrix} 1 & \rho \\ \rho & \sigma_\varepsilon \end{bmatrix}$, where ρ is the correlation between the two error terms, the variance of ϖ is normalized to unity, and σ_ε is the variance of the error term ε . The demand for medical care can consistently be estimated under certain assumptions regarding the error terms in equations (1) and (2), provided Z contains at least one explanatory variable not included in the vector X . Hence, for the purpose of identification, we need at least one variable explaining the choice of the contract which has no direct effect on health care demand.

With this identifying assumption, we can write the expected difference in the demand variables, taking into account the participation choice as a so-called control function approach in the following way:

$$\begin{aligned}
 (3) \quad E(y | D = 1, X, Z) - E(y | D = 0, X, Z) \\
 &= \delta + [E(\varepsilon | D = 1, X, Z) - E(\varepsilon | D = 0, X, Z)] \\
 &= \delta + \rho \cdot \sigma_\varepsilon \cdot \left[\frac{\phi(Z'\alpha)}{\Phi(Z'\alpha) \cdot (1 - \Phi(Z'\alpha))} \right],
 \end{aligned}$$

where ϕ is the density function of the standard normal distribution. This equation can be estimated by simple OLS, where the covariance matrix has to be adjusted according to the estimation of the participation choice (see, for example, Wooldridge, 2002).

An alternative is the instrumental variable estimator.¹ Define $\tilde{X} = [X, D]$, the corresponding parameter vector $\tilde{\beta} = [\beta, \delta]$, and the matrix of instruments ZI (with all exogenous variables in X and the exogenous instruments not included in X). Then, we can write the estimator as follows:

$$(4) \quad \tilde{\beta} = (\tilde{X}' P_{ZI} \tilde{X})^{-1} \tilde{X}' P_{ZI} y,$$

with the projection matrix $P_{ZI} = ZI(ZI'ZI)^{-1}ZI'$.

In the main section, we use an instrumental variable estimator with the predicted program choice \hat{D} from equation (2) as its first instrument. In the sensitivity analysis, we compare the results of this estimator with those of the estimator which follows from (3).

¹ See Vella / Verbeek (1999) for a comparison of the two different estimators.

In the empirical implementation, one also has to take into account the distribution characteristics of health care expenses. First a substantial number of insurees has zero expenses in a given year. Second positive health expenses are typically log-normally distributed. It is common, then, to use a two-part model. The first part estimates the probability of a positive demand, while the second explains the amount of the log of expenses, given the demand is positive. A Heckit approach would be an alternative, where the first and the second part of the model are connected through the correlation of the error terms. See Jones (2000) for an authoritative overview of the pros and cons of the two approaches.

We test for the endogeneity of the program choice by comparing the estimation results under an exogenous choice with those under an endogenous choice. Under the latter assumption, we use a GMM estimator for the first part of the two-part model (Probit-GMM), estimating a non-linear function with instruments (see Greene, 2003, or Hayashi, 2000, for an introduction into GMM). For the construction of the estimator, we use the orthogonality conditions $E[ZI'(y - \Phi(\tilde{X}'\tilde{\beta}))] = 0$, where, again, ZI includes all exogenous variables (inclusive instruments), \tilde{X} is the vector of included exogenous and endogenous variables in the model, $\tilde{\beta}$ is the parameter vector to be estimated, and Φ is the cumulative density function of the standard normal distribution (see Hayashi, 2000). The second part is estimated by OLS with instruments (in the case of endogenous regressors) or simple OLS (if the program choice is exogenous).

We will also analyze the number of physician visits and apply a two part model – a so-called hurdle model – in this context. The hurdle model assumes that the participation decision and the positive count are generated by separate probability processes. Therefore, the two parts of the model can be estimated separately: with a binary process for the first part and the truncated-at-zero count model for the second part (Jones, 2007). We assume a Poisson distribution for the positive counts in these count data and apply a multiplicative specification of the error term to account for unobservable heterogeneity (see Winkelmann, 2000, or Wooldridge, 2002):

$$(5) \quad y = \exp(\tilde{X}'\tilde{\beta}) \cdot u.^2$$

For a consistent estimation with endogenous regressors, instruments have to be found that fulfill the condition $E(u - 1 | ZI) = 0$. The estimation with instruments can again be carried out using a GMM approach (GMM with count models have been applied by Windmeijer / Santos-Silva, 1997, and Schellhorn, 2001). With an exogenous specification of the program choice, a Poisson model can be implemented at the third step of the estimation.

² The normal Poisson model would be realized if $y = \exp(X'b) + \varepsilon$ with $E(y) = \exp(X'b) = \text{Var}(y)$.

The estimation procedure presented here, however, disregards the truncated nature of the second stage (only positive counts). For this reason, we first test for the bias of ‘misspecification’ by comparing a simple Poisson model with the truncated version (for more sophisticated models, see Pohlmeier/Ulrich, 1994 and Santos Silva/Windmeijer, 2001).

Table 2 provides an overview of the estimation procedures and the models used. Step I analyzes the program choice, providing possible instruments for the contract choice in the demand model. Steps II and III include the demand model with and without instruments. The possible endogeneity of the program choice is tested with the Hausman specification test (see Greene, 2003), which compares the model with and without an endogenous selection. Under the null hypothesis of an exogenous choice, both approaches produce consistent estimation results but the endogenous model is inefficient. Under the alternative hypothesis, the endogenous model only leads to consistent estimates. The intuitive test idea, then, is that H_0 , i.e. the exogeneity of the program choice, can only be rejected if the difference between the estimates of the two models is sufficiently large.

Table 2
A 3-step demand model

Step	Description	Model			
I	Program choice	Simple Probit			
II	Two-part model 1 st part: probability of positive demand	Endogeneity of the contract choice			
		No	Yes		
		Simple Probit	Probit with instruments (GMM)		
III	2 nd part: conditional (positive) demand	visits Poisson	demand OLS	visits Poisson (GMM)	demand OLS with instruments

5. Choice and Tests of Instruments

If endogeneity exists, the choice of appropriate instruments becomes crucial. Instruments should be correlated with the choice variable and be orthogonal to the error process. We test the first requirement with an auxiliary regression of the deductible on all exogenous variables and on a list of all sorts of instruments that come into question. The relevance of the instruments can be tested by an *F*-Test of the joint significance of the instruments in this auxiliary regression.

The validity of instruments requires that they are independent of the error process. In particular, instruments should not be correlated with the error term

of equation (1). If more instruments are available for the identification of the equation than required (overidentified equation), the second requirement for valid instruments can be checked by the overidentifying restrictions test. In the linear context, we use the residuals of the OLS regression with endogenous choice (2SLS) and regress them to all exogenous variables. If the explanatory power of this regression is low, the instruments appear to be uncorrelated with the residuals of the 2SLS-regression. The test statistic is N times the regression's R^2 and follows a chi-square distribution with K degrees of freedom, where K is the number of overidentifying restrictions (see Wooldridge, 2002). A test for overidentification in the (non-linear) GMM-context follows the same intuition (see Greene, 2003).

Possible endogeneity in steps II and III, which will be discussed below, is tackled by using the estimated choice of the program as the main instrument. This instrument fully fits the first requirement, given its explanatory power in the program choice estimation. The second requirement for our main instrument for the program choice can be tested only in relation to other instruments, so that an overidentifying restrictions test will apply.

Economic theory should guide the selection of candidates (see Newhouse / McClellan, 1998). The difficult task here is to find, from the set of variables, those instruments that fulfil the above requirements and, at the same time, appear to be economically sound. We used 'Score to go private insurance / other sickness fund', 'Income', 'member with co-insured spouse or children' as instruments. As all these instruments have pros and cons, which we address below, we present an alternative specification in the form of the control function estimator [equation (3)] in section 6.4.

6. Results

6.1 Program Choice

This section presents the results of step I of the estimation procedure. *Table 3* gives selected estimation results of the probit model with the dependent variable D taking the value of one if a person in 2003 opted for the deductible, and 0 otherwise. As independent variables, we use age, gender, income, characteristics of the co-insured spouse and children, and expenses in the two preceding years.

In general, the results are plausible: the probability of participating in the deductible program first increases with age and then decreases in higher age classes. Women show a much lower participation probability than men. The higher a person's likelihood of opting for private health insurance is, the more likely that person is to choose the deductible. Earnings are negatively correlated with the probability of participating in the program. Blue and white collar

Table 3: Choice of the program: Probit model (selected results)

Variable	Coeff.	Std. err.	Variable	Coeff.	Std. err.
Characteristics of the breadwinner ^{a)}					
Female	-0.887**	(0.283)	Other groups	-0.912**	(0.175)
Age	0.063*	(0.033)	Student	-1.153**	(0.219)
Age ²	-0.815**	(0.318)	Welfare recipient	-0.645**	(0.257)
Female * age	0.015**	(0.006)	Retired	-0.224	(0.253)
Not urban	-0.090	(0.057)	Self-employed	-0.436**	(0.158)
Severity status	-0.147	(0.182)	Monthly income	-0.374**	(0.137)
Score to go private	0.011**	(0.003)	(Monthly income / 1000) ²	0.077**	(0.028)
(Score to go private) ²	-0.044**	(0.014)	Contract duration	0.225	(0.197)
Characteristics of the spouse and children					
With adult co-insured	0.549**	(0.252)	Age of spouse	-0.015**	(0.005)
N of adult co-insured	-0.466**	(0.108)	Age of child	-0.007	(0.009)
With co-insured child	-0.100	(0.140)	Male child	0.046	(0.090)
N of co-insured children	-0.030	(0.053)	Male	-0.096	(0.160)
Expenses of the breadwinner					
Amb. care in 01	6.24E-04**	(1.16E-04)	Expenses of the adult co-insured		
Amb. care in 02	-2.97E-04**	(1.40E-04)	Amb. Care in 01	-5.05E-04**	(2.21E-04)
Positive demand for preventive care	0.234**	(0.051)	Amb. Care in 02	-2.98E-04*	(1.75E-04)
Positive demand in all quarters of 01 and 02	-0.328**	(0.085)	Hospital care (average 01 and 02)	-3.82E-05	(4.74E-05)
N of consultations (average 01 and 02)	-0.048**	(0.007)			
Hospital care (average 01 and 02)	-7.79E-05**	(3.19E-05)			
Pseudo R ^{2b)}	0.587				
N	4774				
Log L	-2527				
Log L ₀	-3258				
			$LR = 2 \cdot (\log L - \log L_0) = 1461^{**} (DF = 39)$		

^{a)} Other factors are: score to choose a different sickness fund (linear and quadratic), earnings, contract duration, lives at the border (implying that some of the health care demand occurred outside the domain of the local physician association).
^{b)} We use the formula suggested by Zavoina and McKelvey for Probit models (see Greene, 2003, 684).
** level of significance 1 %.
* level of significance 5 %.

workers show the highest participation preference compared to students, welfare recipients, retirees, self-employed, and other persons. Since we have no precise information on the earnings of students, welfare recipients, retirees, and self-employed, we also run the estimation without these groups of persons. This estimation, with a residual sample of $N = 3,338$, shows the same qualitative results as the original estimation.

A bigger family is less likely to participate in the program. The same result holds with an increase in the age of a co-insured adult family member. In contrast, the number, the gender, and the age of children are not significant in explaining participation. These results are in line with the incentives given by the program, i.e. while adult co-insured family members are subject to the deductible, the children are not.

Finally, past health care expenses of the breadwinner and his/her spouse show the expected sign. More demand for ambulatory care (total expenses as well as the number of consultations) in 2002 decreases the probability that a person chooses a deductible in 2003. The coefficient for average hospital care expenses in 2001 and 2002 points in the same direction. By comparison, higher expenses for preventive care increase the participation rate. The coefficient of the breadwinner's ambulatory care expenses in 2001 is positive. However, for all three years, the average medical expenses are higher in the control group. The positive sign in 2001 is due to outliers among the program participants. The negative coefficient of "positive demand in all quarters" points in the same direction: clients that saw a physician in all eight quarters preceding the start of the program had a significantly lower probability of participating in the deductible program than low users.

Despite the small sample size, the explanation power of the estimation model is remarkably high. A likelihood ratio test reveals that the joint hypothesis of zero coefficients can be rejected. In all, the results show a high rationality of the clients' choice: factors that increase expected future expenses decrease the probability that they opt for the deductible program.

6.2 Physician Visits

Table 4 presents the results on the probability of consulting a physician at least once in 2003, i.e. step II of the estimation procedure. The GMM test of overidentification confirms the choice of instruments. The six additional instruments are earnings, earnings squared, the share with co-insured adults, the share with children, the score to opt for private insurance, and the score to choose a different sickness fund. These instruments make sense as they will affect the choice of contract. Of special interest is the fact that the model picks earnings as an instrument, implying that one's earnings do not have a strong effect on the decision to see a physician at all. This result contradicts findings

from, for example, the ECuity project (see, among others, Doorslaer et al., 2004, and Jones / Wildman, 2005). It may, however, be explained with the specific insurance status of the study group, as we only include voluntary insured persons who either earn a high income or for whom income plays a minor role. Besides the coefficients age and sex, coefficients of those variables that are closely related to the health status are highly significant. The model reveals that not only clients with a high demand for health care in the last two years but also clients with regular past visits to the doctor (at least one physician visit in every 8 quarters of 2001 and 2002) have a high probability of visiting the physician this year.

Table 4
Estimation of the probability of consultations
Simple Probit and GMM Probit

Dependent variable	If number of consultation is at least 1, then 1			
	$D = 0$		$D = 1$	
<i>N</i> of observations	2737		2037	
Mean	0.858		0.713	
	Simple Probit		GMM Probit	
Independent variable	Coeff.	Std. err.	Coeff.	Std. err.
Intercept	-0.017	(0.173)	-0.203	0.344
Age	0.008**	(0.003)	0.009**	0.003
Female	0.282**	(0.077)	0.255**	0.069
Severity case	-0.104	(0.181)	-0.081	0.177
<i>N</i> of months insured	-0.027**	(0.009)	-0.018	0.015
At least one physician visit in all 8 quarters of 01 and 02	1.049**	(0.210)	1.088**	0.221
<i>N</i> of specialist visits in 01 and 02	0.056	(0.059)	0.051	0.059
Log of total cost in 01	0.107**	(0.012)	0.108**	0.012
Log of total cost in 02	0.172**	(0.013)	0.174**	0.013
Lives at the border	-0.048	(0.048)	-0.055	0.044
Deductible	-0.330**	(0.053)	-0.202	0.180
R^2	0.54	Number of instruments		7
		GMM test		8.67
		Hausman		0.99

** level of significance 1 %.

* level of significance 5 %.

The Hausman test strongly indicates that the exogenous specification of the model cannot be rejected. Therefore, the coefficient of the deductible is an unbiased measure of the effect of the deductible on the probability of seeing a physician. The difference between participants and non-participants is 0.07. The marginal effect in a Probit for a dummy-variable is $\Pr | 1 - \Pr | 0$. As the

observable difference is 0.145 – see head of *table 4* –, the selection effect is 50 percent of the total effect.

The difference is significant in the exogenous model only. It is interesting to differentiate between GP visits and specialist visits (results are not reported). The results indicate that the probability of consulting a GP is unaffected by the deductible. By comparison, program participants show a significantly smaller probability of consulting a specialist than do non-participants.

For the positive consultations, we first checked for the influence of a misspecification of the assumed distribution for the count data. Since the distribution contains all non-negative integers (including zero), we expect that the (truncated) positive values do not strictly follow a Poisson distribution. A comparison between a simple Poisson and the truncated model yields similar results, indicating that the misspecified distribution of the count data is not decisive. For this reason, we used the multiplicative Poisson model in the estimation.

Table 5 shows the estimation results for the number of conditionally positive consultations (step III, count data). Again, exogeneity of the program choice cannot be rejected. The result is weaker as only two additional instruments, the share of clients with co-insured adults and the score to opt for private insurance, can be detected. These show a high correlation with the contract choice and a small correlation with the demand.

The effect of the deductible on conditional consultations is –30 percent ($\approx \exp(-0.363) - 1$), which corresponds to 3.93 consultations. If the selection effect did not exist, we would expect 9.17 consultations ($= 13.1 - 3.93$) among the program participants. From the observed difference of 7.2 consultations between the two groups, 55 percent can then be attributed to moral hazard while the remaining 45 percent is due to selection.

6.3 Curative vs. Preventive Health Care Demand

The deductible in ambulatory care applies to physician visits. The patient pays € 20 per visit. This gives the patient an incentive to reduce the number of consultations and to extend the demand for services per consultation. We checked for the latter possibility and found the co-payment had no effect on this: no difference exists in the expenses per visit between participants and non-participants for both the breadwinner and their co-insured family members (results not presented).

If the consultation mainly aims at preventive care or concerns contraception, the co-payment does not apply. We might, therefore, detect differences in the demand between curative and preventive care. Again, a two-part model was estimated with and without endogenous contract choice. As expected, the probability of incurring a positive demand for curative care was the same as the probability of seeing a physician at least once (results not presented).

Table 5

**Estimation of consultations, conditional on being positive
Poisson multiplicative and GMM Poisson multiplicative**

Dependent variable	Number of consultations > 0			
	$D = 0$		$D = 1$	
<i>N</i> of observations	2348		1458	
Mean	13.1		5.9	
	Poisson multiplicative		GMM multiplicative	
Independent variable	Coeff.	Std. err.	Coeff.	Std. err.
Intercept	1.086**	(0.311)	1.148**	(0.530)
Age	0.001	(0.011)	-0.001	(0.012)
Age ²	0.143	(0.106)	0.155	(0.111)
Female	0.608**	(0.165)	0.601**	(0.174)
Female * age	-0.009**	(0.003)	-0.009*	(0.004)
Severity status	0.220**	(0.075)	0.211	(0.080)
<i>N</i> of months insured	-0.013	(0.008)	-0.014	(0.018)
At least one physician visit in all 8 quarters of 01 and 02	0.471**	(0.039)	0.465**	(0.048)
<i>N</i> specialist visits in 01 and 02	0.047	(0.050)	0.051	(0.049)
Log of total cost in 01	0.057**	(0.010)	0.058**	(0.010)
Log of total cost in 02	0.125**	(0.013)	0.126**	(0.014)
Lives in at the border	0.041	(0.034)	0.042	(0.035)
Deductible	-0.363**	(0.050)	-0.381*	(0.194)
		Number of instruments		3
		GMM test		3.34
		Hausman		0.15

** level of significance %

* level of significance 5 %.

Table 6 presents the results of the conditional demand for curative care (step III, continuous data). The applied instruments were (additional to the predicted choice of the deductible) the share of co-insured adults, the share of co-insured children, the score of opting for private insurance, and the two earnings variables. First, the variables that are closely related to the health states show a high significance (at least one physician visit in all 8 quarters of the years 2001 and 2002, log of the total cost of the two previous years). The estimation results do not depend on whether we include the instruments for the programme choice or not. In particular, the coefficients for the choice remain unchanged and significant. Tests point in the same direction, as they confirm the choice of instruments and do not reject the exogeneity of the program choice either.

The coefficient for the deductible is significantly negative. Hence, the deductible reduces the conditional demand for curative care. For the entire two-

part model, we derive the following result: a difference of € 172.60 in curative ambulatory care expenses between participants and non-participants is observed. € 71.70 or 42 percent of this difference can be explained by selection, i.e. 58 percent of the difference is attributable to moral hazard.

Table 6
Estimation of conditional curative expenses
OLS and OLS with instruments

Dependent variable	Log (curative expenses > 0)			
	<i>D</i> = 0		<i>D</i> = 1	
<i>N</i> of observations	2348		1458	
Mean expenses	303		123	
	OLS		OLS with instruments	
Independent variable	Coeff.	Std. err.	Coeff.	Std. err.
Intercept	3.634**	(0.339)	3.610**	(0.512)
Age	−0.008	(0.013)	−0.008	(0.015)
Age ²	0.216	(0.120)	0.215	(0.145)
Female	0.157	(0.192)	0.159	(0.212)
Female * age	−0.001	(0.004)	−0.001	(0.005)
Severity status	0.176	(0.109)	0.178	(0.116)
<i>N</i> of months insured	−0.014	(0.008)	−0.013	(0.016)
At least one physician visit in all 8 quarters of 01 and 02	0.506**	(0.050)	0.507**	(0.057)
<i>N</i> specialist visits in 01 and 02	−0.037	(0.049)	−0.037	(0.050)
Log of total cost in 01	0.082**	(0.011)	0.082**	(0.011)
Log of total cost in 02	0.162**	(0.012)	0.163**	(0.012)
Lives at the border	0.074*	(0.035)	0.074*	(0.038)
Deductible	−0.404**	(0.048)	−0.393*	(0.172)
<i>R</i> ²	0.275	Number of Instruments		6
			GMM test	1.35
			Hausman	0.014

** level of significance 1 %.

* level of significance 5 %.

Let us finally consider preventive health care demand. The model explaining the probability of incurring a positive demand is not sensitive to the selection of the instruments (step II of the estimation procedure, see *Table 7*). Again, the hypothesis of an exogenous choice of the contract is not rejected. The probability of a positive preventive demand increases with the amount of expenses in the preceding years. Individuals with a severe status have a lower probability of a positive demand for preventive care, while women have a higher probability. The deductible (which does not apply to preventive care) is not significant.

This is somewhat surprising as one might expect a positive cross-price effect of the co-payment for curative care on the demand for preventive care. Substitution between preventive and curative care is possible as physicians can provide curative care in a consultation that mainly has a preventive character.

Table 7
Estimation of probability of preventive care
Simple Probit and GMM Probit

Dependent variable	If expenses for preventive care >0, then 1			
	$D = 0$		$D = 1$	
<i>N</i> of observations	3402		1372	
Mean	0.328		0.232	
	Probit		GMM Probit	
Independent variable	Coeff.	Std. err.	Coeff.	Std. err.
Intercept	-4.680**	(0.460)	-4.714**	(0.937)
Age	0.113**	(0.018)	0.109**	(0.026)
Age ²	-0.767**	(0.173)	-0.722**	(0.241)
Female	3.025**	(0.240)	3.033**	(0.309)
Female * age	-0.042**	(0.005)	-0.042**	(0.006)
Severity status	-0.236*	(0.142)	-0.190	(0.175)
<i>N</i> of months insured	-0.024**	(0.009)	-0.018	(0.026)
At least one physician visit in all 8 quarters of 01 and 02	0.136**	(0.064)	0.136*	(0.081)
<i>N</i> of specialist visits in 01 and 02	0.025	(0.058)	0.031	(0.082)
Log of total costs in 01	0.091**	(0.013)	0.093**	(0.019)
Log of total cost sin 02	0.057**	(0.013)	0.061**	(0.021)
Lives at the border	0.031	(0.073)	0.044	(0.098)
Deductible	-0.001	(0.054)	0.074	(0.287)
R^2	0.49	Number of instruments		5
			GMM test	3.51
			Hausman	0.58

** level of significance 1 %.

* level of significance 5 %.

Table 8 shows the results for the conditional demand for preventive care. Only the gender variable is significant: women incur higher expenses. Higher demand on the part of women can be explained by the fact that social health insurance covers more preventive services for women than for men. The deductible is, again, not significant. Thus, a cross-price effect is once again absent.

The explaining power of the model is weak. R^2 is low and an F -test does not reject the joint hypothesis of zero coefficients. It appears that factors other than those included in the model might explain the demand for preventive care.

Table 8
Estimation of expenses for preventive care being positive
OLS and OLS with instruments

Dependent variable	Log (preventive expenses > 0)			
	D = 0		D = 1	
N of observations	899		473	
Mean	48.7		37.9	
	OLS		OLS with instruments	
Independent variables	Coeff.	Std. err.	Coeff.	Std. err.
Intercept	2.522**	(0.458)	2.986**	(0.630)
Age	0.026	(0.017)	0.022	(0.018)
Age ²	−0.151	(0.159)	−0.118	(0.162)
Female	0.620**	(0.209)	0.609**	(0.209)
Female * age	−0.013**	(0.004)	−0.012**	(0.004)
Severity status	−0.116	(0.116)	−0.154	(0.122)
N of months insured	0.003	(0.009)	−0.016	(0.020)
At least one physician visit in all 8 quarters of 01 and 02	0.023	(0.052)	0.000	(0.056)
N specialist visits in 01 and 02	−0.042	(0.058)	−0.036	(0.058)
Log of total costs in 01	0.014	(0.013)	0.012	(0.013)
Log of total costs in 02	−0.010	(0.014)	−0.017	(0.015)
Lives in at the border	−0.046	(0.069)	−0.062	(0.071)
Deductible	−0.048	(0.052)	−0.266	(0.209)
R ²	0.011	Number of instruments		5
			GMM test	3.06
			Hausman	1.15

** level of significance 1%
* level of significance 5%

6.4 Sensitivity Analysis of the Model Specification

In the sensitivity analysis, we study more closely the consequences of three possible misspecifications for the results: (1) the endogeneity of the program choice in the applied parametric estimations, (2) the effects of different samples and sets of explanatory variables, and (3) the inflexibility of parametric estimation methods.

(1) Endogeneity

In all the models, we used the predicted program choice of the clients, based on a probability model, as the main instrument (see section 6.1). Alternatively, we employed additional variables that accurately explained the program choice. These, by implication, are highly correlated with the main instrument.

For instance, the correlation between \hat{D} and the instrument ‘score to opt for private insurance’ is 0.55. Multicollinearity between instruments can lead to biased estimation results. To exclude this, we estimated all models with endogenous choice using two further specifications: First the use of the control function approach according to (3), and second the use of all other instruments except \hat{D} in the instrumental variable setting as in (4). In both cases, we find a high congruence of the results with those of the original specification, i.e. the exogeneity of the program choice cannot be rejected and we again find a significantly negative coefficient for the deductible.

(2) Different samples and different variables

In the original estimations at steps I and II of the model (demand), we simply used the same set of explanatory variables guided by the result at step II (probability of positive demand). This led to an unsatisfactory adaptation at step III. Further estimations show that the insignificant coefficient for age is due to multicollinearity. If we model age with a linear term only, it becomes significantly positive. On the other hand, using squared terms for past demand in the explanation of demand for curative care significantly improves the estimation. Both modifications, however, do not qualitatively change the result regarding the deductible: it significantly reduces demand for health care.

A further problem concerns the incorporation of persons with voluntary insured status who differ from the average high-earnings person. Therefore, we excluded students, welfare recipients, and retired persons from the sample, and repeated the estimations. There is no significant difference in the results.

(3) Parametric estimations

The results of the parametric estimations depend on the assumed distribution (normal, Poisson). Furthermore, all estimations reveal that we can capture the endogeneity of the program choice with observable factors. This allows for the use of non-parametric methods, which do not require the distributions of the explanatory variables and the error terms to be parameterized, and are strictly restricted to the observability of all the factors that cause the endogeneity of the program choice.

For this reason, we also conducted a matching method based on the propensity score estimated from the first stage [see (2)]. We employed the nearest-neighbourhood method, which, for each program participant, searched for a sibling in the control group. This approach led to a matching group which was almost identical to the group of participants with respect to the important control variables. The effect of the deductible could then simply be derived from the difference in the means of the two groups.

The results of the two approaches are similar. However, the effects of the deductible are substantially smaller in the matching approach. This follows

from the exclusion of ‘outliers’ in the group of participants, i.e. persons that could not be sufficiently matched with persons in the control group.

7. Conclusion

Patients’ co-payments in health care are widely debated in European countries. Switzerland combines a mandatory deductible for medical services with a co-payment rate of 10 percent exceeding the deductible and offers a choice of higher deductibles and lower premiums. Two years ago, Germany introduced a one-off fee of € 10 per quarter for physician visits. Critics of such incentive measures that are designed to curb rising health care costs argue that health demand is absolutely price-insensitive and thus point to their adverse distributional effects. Of course, if copayments do not work, introducing them in social health insurance would be ill-advised.

The effect of the Swiss deductibles on health care has been intensively studied. Gardiol et al. (2003) and Werblow/Felder (2003) find a significant price-elastic demand for health care. Schellhorn (2001), analysing physician visits, disagrees as he finds no association between the number of GP visits and the deductible. In a more recent study (Gerfin/Schellhorn, 2006), applying non-parametric methods to minimize the necessary distribution assumption for the identification of the incentive effect, he concludes that a higher deductible leads to a decrease in the number of GP visits. Winkelmann (2004) studies the effect of the German health care reform of 1997 on the number of physician visits. He finds that the increase in patients’ co-payments of drug expenses leads to a 10 percent decline in physician visits, with differences between high and low users of health care services. In a more recent quantile regression study (Winkelmann, 2006), he more closely investigates the distributional effects and finds that the increase in co-payments has a much larger effect in the first quantile than in the upper side of the distribution.

While the reform in 1997 was a natural experiment as it was extended to all insurees in German social health insurance, the present paper addresses an option offered by a certain sickness fund to a subset of its clients, the voluntarily insured in social health insurance. As its cornerstone, it includes a € 20 co-payment for physician visits, which, however, does not apply if a consultation has mainly preventive character (for instance, seeing a physician for a screening service or for contraception). If a contract is optional, an accurate modelling of the decision choice is crucial. If the regressors are endogenous, estimators become inconsistent. We use enrollee characteristics, including expenses for health care in the two years preceding the start of the deductible program, to model the program choice. This information is so rich that, according to the econometric results, the decision to participate in the deductible program can be regarded as exogenous. Thus, Van Vliet (2004), also using a rich data set

including diagnostic information, was probably right to assume the exogeneity of the contract choice.

The estimation results confirm the economic intuition. The € 20 user charge has a significant effect on the number of physician visits. The probability of seeing a physician at least once in 2003 declines by 7 percentage points. A closer inspection reveals that the doctor fee has no effect on the probability of visiting a general practitioner while it significantly decreases the probability of seeing a specialist. These non-existent or, rather, weak effects on the first visit confirm the findings by Augurzky et al (2006), which in a difference-in-differences approach found no significant effect of the one-off € 10 fee introduced in 2004.

In contrast, we find that the € 20 visit fee has a strong effect on the number of consultations, conditional on being positive. This drops by 30 percent (which corresponds to 4 consultations). Parallel with the decline in physician visits, we observe a decrease in the expenses for curative health care. Interestingly enough, we see no increase in the expenses per consultation as a result of the decline in the number of visits.

By comparison, any demand for preventive care, exempted from co-payment, is not affected by the deductible. This holds true for both the probability of incurring positive preventive demand and the conditional demand. The political agenda gives high priority to prevention. The present paper lends support to a view that prevention activities can, in fact, be shielded by a physician fee that covers curative care but not preventive care.

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