Does physicians' compensation affect the probability of their vetoing generic substitution?

David Granlund*
Department of Economics, Umeå University
SE-901 87 Umeå, Sweden; and
The Swedish Retail Institute (HUI)
SE-103 29 Stockholm, Sweden
david.granlund@econ.umu.se

March 2008

Abstract

Physicians' decisions whether or not to veto generic substitution were analyzed using a sample of 350,000 pharmaceutical prescriptions. Point estimates show that - compared to county-employed physicians on salary - physicians working at private practices were 50-80% more likely to veto substitution. The results indicate that this difference is explained by the difference in direct cost associated with substitution, rather than by private physicians' possibly stronger incentives to please their patients. Also, the probability of a veto was found to increase as patients' copayments decreased. This might indicate moral hazard in insurance, though other explanations are plausible.

Keywords: doctors; salary; fee for service; moral hazard; prescriptions; drugs

JEL classification: D86; I11; L33

*I am grateful to Kurt Brännäs, Kerstin Gustafsson, Karl-Gustaf Löfgren, Carl Lönnbark, Niklas Rudholm and participants at seminars at Umeå and Gävle Universities for helpful comments and suggestions, and to the Swedish Competition Authority for a research grant that supported the work. I also thank Västerbotten County Council and IMS Sweden for providing the datasets and Eva Lindberg for contributing the information about the physicians.

1 Introduction

Imperfect information often results in "Principal-agent problems" when one person, the agent, is employed to act on behalf on another, the principal (Ross, 1973). In the healthcare sector the problem is complex since efficiency requires physicians to act not only as agents for their patients, but also for third-party payers, insurers (Blomqvist, 1991, and Shortell, 1998). This study analyzes how economic incentives affect physicians' decisions whether or not to veto generic substitution, and also whether their decisions suggest that they internalize differently the costs occurring to their two principals.

Since October 2002, pharmacists in Sweden have been required to substitute the prescribed pharmaceutical product to the cheapest available generic when neither the prescribing physician nor the patient oppose it. Patients who oppose substitution have to pay the difference in price themselves, but if the physician vetoes it for medical reasons, they are subject only to the normal copay requirement under Swedish pharmaceutical insurance.

Although similar reforms have been introduced in many European countries and American states, what determines whether physicians' veto substitution has, to my knowledge, not been studied previously.[1] This is an important issue because physicians' decisions not only directly affect patients' and insurers' costs for pharmaceuticals, but also indirectly since more bans against substitution likely reduces price-competition between pharmaceutical firms.[2] In the sample used for this study, brand-name products for which substitution was vetoed by physicians were on average 218% more expensive than the cheapest generic alternative; whereas the corresponding figure for other brand-name products was only 15%. This correlation might indicate that physicians' decisions whether or not to veto generic substitution have an important effect on price-competition among pharmaceutical firms, but this warrants further research.

The primary purpose of this study was to analyze whether privately employed physicians were more or less inclined to oppose substitution, compared

[1] One explanation could be lack of data: Hellerstein (1998) noted that the US NAMCS-data unfortunately lack information about whether substitution was vetoed, while Mossialoa, Walley and Rudisill (2005) noted on a general scarcity of good prescription data for several European countries.

[2] Granlund and Rudholm (2007) analyzed the effect on price-competition and pharmaceutical prices of the Swedish substitution reform.

to county-employed physicians. Private physicians have a stronger incentive to please their patients in order to keep them, since their income depends on the number of patient-visits, whereas county physicians work on salary. Opposing substitution, if the patient suggests that, might be a costless way of doing this. Allowing substitution might also be time-consuming for the physician if it worries the patient. Hence, each consultation could take longer, resulting in fewer of them, and again less income. Private physicians might also have stronger brand-name loyalty since, for example, they are less restricted, compared to county-employed physicians, from participating in education organized and paid for by pharmaceutical companies. The hypothesis to be tested is thus that private physicians were more likely than county physicians to veto substitution.

Another purpose was to analyze the effect of patients' copayments on physicians' decisions. The Swedish pharmaceutical insurance is non-linear, with patient-copayments decreasing as total expenditure increases. This provided an opportunity to study whether physicians internalized patients' costs more than costs to the insurer, indicating moral hazard in insurance (Pauly, 1968).[3]

The analyses were done using a sample of 350,000 observations drawn from a micro-dataset covering all prescriptions dispensed in the county of Västerbotten, Sweden - or dispensed elsewhere in Sweden to inhabitants of Västerbotten - during 43 month after the substitution reform. The dataset includes information about the patients, prescribers, prices, copayments, pharmaceuticals prescribed and dispensed, and about whether the physician or patient opposed substitution.

Since the values were observed at micro-level, the risk of estimators being biased towards zero was reduced; this is otherwise a common problem when aggregated data are used as proxies for mirco-variables.[4] Using register-data also eliminated recall-bias, as well as selection-bias, which can be a problem if for example not everyone participates in an experiment or answers a questionnaire. The size of the dataset also substantially reduced the risk of accepting a false null-hypothesis which is otherwise a common problem when studying questions,

[3]Moral hazard in insurance is also called ex post moral hazard, to distinguish it from moral hazard referring to changed risk behavior.

[4]Proxy variables can be seen as measurements with errors of the micro-variables. Greene (2003, Chapter 5) describes how measurement errors lead to bias towards zero, so called attenuation.

such as here, where a large part of the variation, for various reasons, cannot be explained by the observables.

Gosden, Pedersen and Torgerson (1999) reviewed the literature on the effects of salary payments on physicians' behavior. They reported some evidence that payments of salary was associated with fewer referrals and tests compared both with fee-for-service (FFS) payments and capitations. Compared with FFS payments, salary payment also correlated with fewer procedures per patient, fewer patients per physician, longer consultations, more preventive care, and different patterns of consultation. Nassiri and Rochaix (2006) found that primary-care physicians in Quebec reacted both to temporary removal of expenditure caps and to changes in the relative price of consultations by changing their treatment pattern. Dusheiko et al. (2006), studying the effect of financial incentives on general medical practices in England, found that abolishing foundholding increased elective surgery by 3-5%.

The theoretical analysis here follows Hellerstein (1998) and Lundin (2000) who both - as well as Leibowitz, Manning and Newhouse (1985) - studied the choice between prescribing brand-name or generic pharmaceuticals. Using U.S. data, the first and third study found that the choice was not a function of the insurance plan.[5] On the other hand, using a Swedish dataset covering seven pharmaceuticals, Lundin found evidence of moral hazard: Patients with low copayments were more likely to receive brand-name pharmaceuticals. Crown et al. (2004) found no statistically significant effect of insurance plans' mean copayment-rates on patients' treatment patterns for asthma. However, Rudholm (2005) found significant effects of individual patients' copayment-rates on both quantities dispensed and price. Rudholm also included a variable indicating for privately employed physicians in his regressions but, except in one subsample, found no statistically significant effects of this variable.

Empirical results presented in this paper show that physicians were more likely to oppose substitution if they were privately employed and the lower the patients' copayment-rates were. However, the likelihood of private physicians vetoing substitution was not found to increase faster than that for county-employed physicians when patients' copayment-rates decreased. According to the theoretical model, this implies that private physicians' higher likeliness to

[5]However, Leibowitz, Manning and Newhouse found that individuals with more generous insurance plans bought more prescription pharmaceuticals.

veto substitution was caused by them having higher direct costs associated with substitution rather than stronger incentives to please their patients.

The next section describes the compensation of private and county-employed physicians in the county of Västerbotten, as well as the Swedish pharmaceutical insurance system. Section 3 presents the theoretical model. The data are discussed in section 4.1, the empirical specification in section 4.2, while section 4.3 contains the results. Finally, in section 5 the paper's conclusions are presented.

2 Rules and incentives

2.1 Physicians and their compensation

There were nearly 1000 physicians working in the county of Västerbotten during the study-period. Most of them were county-employed, paid on salary, but nearly 40 physicians worked at small private practices, indicated here as *Private*.[6]

Twenty of the private physicians worked at practices that were nevertheless paid by the county council, according to three different types of contract, while the remainder were paid privately, either directly by patients, or possibly under contract to private health insurance companies.[7] The most common type of county-council contract stipulated fee-for-service reimbursement, according to the so called national rate (Nationella taxan). The second type, individually negotiated contracts, also stipulated fixed compensation per procedure. All practices paid according to these two types of contracts were single practices. Finally, four physicians worked at two so called "house-doctor practices", which were paid fee-for-service plus a capitation per patient registered at their practice.

All three types of contract stipulated that compensation increased with the number of patient-visits. The compensation-schemes were nonlinear however: Compensation per procedure was reduced if the practice reached certain breakpoints. All contracts also stipulated that, for practices to receive compensation from the county, they were not allowed to charge higher copayments for

[6] The National Board on Health and Welfare estimated the number at 37 in 2005. In addition five physicians worked at private occupational health services, excluded from this analysis.

[7] In the data it is possible to identify whether a prescription was written by a private physician, but not whether or not they had a contract with the county.

patient-visits than did the public healthcare providers. Hence, for all these publicly-financed physicians, price was essentially fixed, and their only competition variable was quality.

There were also physicians working in other organizational forms, including ten working for the private company Carema, which ran Dragonen's health center during the last five months of the study-period. Carema received a lumpsum payment for the first 12 months, after which compensation would depend on the number of registered patients at the health center. The incentives for physicians working for Carema probably differed from those for other private physicians in two ways: The incentives for the company differed from those for private practices and then there were probably internal principal-agent problems.[8]

2.2 Patients' copayments and the substitution reform

In the Swedish pharmaceutical insurance system patients pay costs up to 900 Swedish crowns per 12-month period; 50% of the cost from 900 to 1700 SEK; 25% from 1700 to 3300 SEK; and 10% from 3300 to 4300 SEK; after which all costs during the period are paid by the insurance (specifically, by Swedish county councils). However, there are some exceptions: Some pharmaceuticals are always free of charge for the patient, and others are not covered by insurance at all.

Another exception is because of the substitution reform that came into effect October 1, 2002. The reform (Lag 2002:160) requires that pharmacists inform

[8] Another organizational form is the so-called community company, which ran Jörn's health centre since October 2003 with two part-time physicians. The company received its compensation in fixed lumpsum payments. The incentives for this health center therefore seem similar to those of the health centers managed by the county council. In addition, there were two personnel-managed health centers with greater autonomy from the county council, which were also compensated by lumpsum payments. Granlund, Rudholm and Wikström (2006) found no clear effect of increased autonomy on the prescription-behavior of physicians working at these health centers. Therefore, these centers, and Jörn's health center were treated in the empirical analysis here as ordinary county health centers. I tried including indicator-variables for prescriptions written at Jörn's health centre and at the personnel managed health centers, but the odds-ratios for these variables were not significantly different from unity, and the qualitative results were not affected by including them. By qualitative results I mean that the odds-ratios for *Private* and for the four copayment variables are significantly above unity, and that the point estimates for the copayment-variables monotonically increased as the copayments decreased.

patients when there are substitutes available, and that the cheapest available generic product considered to be a perfect substitute by the Swedish Medical Products Agency would be provided within the Swedish pharmaceutical insurance system.[9] Patients need not accept substitution, but the entire extra cost will then be charged to them.[10]

Physicians can veto substitution for medical reasons, in which case the extra cost is covered by the pharmaceutical insurance system. Thus, patients who would otherwise refuse substitution could save money if their physician opposed it instead, given that their pharmaceutical cost for that 12-month period was more than 900 SEK. As total pharmaceutical cost goes up, patients could save more if their physician vetoed substitution; above 4300 SEK total cost, they would pay nothing if their physician opposed substitution, versus paying the entire difference themselves.

Of course, patients who would not refuse substitution could have lower co-payments if their physician allowed substitution. If there is no medical reason against substitution, physicians will generally agree to it for such patients, given that their direct cost for this (discussed below) is not too high, since there is then no conflict between the rules physicians should follow and the patient's interest.

3 Theoretical model

The physician chooses which pharmaceutical to prescribe and whether or not to veto substitution. If the physician does not veto, the patient then decides whether or not to refuse substitution. The equations below aim to describe how patients' and physicians' utilities are affected by a veto against substitution.[11]

[9] All pharmaceuticals in Sweden are sold through a nationwide government-owned monopoly.

[10] Some employees were covered by supplemental medical insurance for prescription drugs, provided by their employer. (According to Lundin, 2000, 10% of the employees were covered by such insurances in 2000.) However, even if the entire out-of-pocket cost were covered by such extra insurance, the cost was not reduced to zero for the patient, since such fringe benefits were subject to taxation. Also, many patients were retired (45% in the dataset used here) and thus not covered by extra insurance.

[11] The model is inspired by Hellerstein (1998) and Lundin (2000).

The patient's utility function is written

$$U_p = H + (Y - cost^p), \tag{1}$$

where H is monetized health; Y is income; and $cost^p$ is the patient's monetary cost for the pharmaceutical consumed. Thus $(Y - cost^p)$ represents the utility that the patient obtains from consumption of other goods. Let ΔP be the price-difference between the prescribed pharmaceutical and the cheapest generic, and $E[\Delta H]$ be the patient's expectation of the resulting difference in health outcomes. The patient will then refuse substitution if $E[\Delta H] > \Delta P$.[12]

The physician's utility function is written as

$$U_{ph} = \gamma_1 H - \gamma_2 cost^p - \gamma_3 cost^{in} - \gamma_4 c, \tag{2}$$

where γ_1, γ_2, γ_3 and γ_4 are the weighs the physician puts on the health of the patient, the patient's monetary cost, the insurer's monetary cost ($cost^{in}$) and the physician's own direct cost (c). The physician might internalize the consequences for the patient because of altruistic considerations, or because of pecuniary incentives. Pecuniary incentives could arise, for example, since patients can change physician if they are not satisfied. The last term, c, is a direct cost that the physician might experience from allowing substitution, for example if it raises questions from the patient about the difference between the prescribed and dispensed pharmaceutical. Answering such questions might be time-consuming, reducing the number of possible patient-visits per day. This term could also express the strength of brand loyalty, reflected in a non-pecuniary cost for allowing substitution away from a brand-name pharmaceutical, if for example the physician feels a moral responsibility to support the firm that has invested in research to develop the pharmaceutical.

A principal-agent problem between patients and physicians arises because physicians have private information about patients' health (Blomqvist, 1991). For notational simplicity, I illustrate this asymmetry in information by assuming that the physician knows with certainty the difference in health outcome, ΔH, caused by the prescribed pharmaceutical versus the cheapest generic.[13]

[12] Remember that the patient had to pay the whole price-difference if they rejected substitution.

[13] The difference in health effect may arise, for example, if a patient was sensitive to inert ingredients, or simply because a substitution to a pharmaceutical with another color or form might cause some patients to confuse their pharmaceuticals.

Defining θ as the patient's copayment-rate and inserting ΔP and ΔH in equation (2) yields that the physician, if he/she knows that the patient would allow substitution, would veto it if

$$\gamma_1 \Delta H - \gamma_2 \Delta P \theta - \gamma_3 \Delta P(1-\theta) + \gamma_4 c > 0. \tag{3}$$

If $\theta = 1$, $\gamma_1 = \gamma_2$ and $c = 0$, equation (3) simplifies to $\Delta H > \Delta P$. That is, if insurer cost is not affected, the physician internalizes the patient's health and monetary cost equally and has no direct cost; then the physician would act as a perfect agent for the patient.[14] If $\theta \neq 1$ and $c \neq 0$, the physician would only take societally optimal decisions if $\gamma_1 = \gamma_2 = \gamma_3 = \gamma_4$. If $\gamma_2 > \gamma_3$, then the physician is more likely to oppose substitution the lower the patient's copayment-rate is, that is, there is moral hazard in insurance.

If the physician knows that the patient would oppose substitution, the physician would only oppose it if

$$\gamma_2 \Delta P(1-\theta) - \gamma_3 \Delta P(1-\theta) + \gamma_4 c > 0. \tag{4}$$

If $c = 0$, equation (4) simplifies to $\gamma_2 > \gamma_3$. Hence, the model shows how the presence of moral hazard can be tested for by analyzing whether physicians were more likely to oppose substitution when the patients' copayment-rates were low.

The model also illustrates two set of reasons why private physicians might be more inclined than county physicians to veto substitution, which can be tested for. First, private physicians might have higher direct costs from allowing substitution. If, as noted earlier, doing so requires more time per patient and hence result in fewer patient-visits per day, this will be more costly for private physicians since it will affect their income. They might also have stronger brand-name loyalty since they, as opposed to county-employed physicians, are not restrained by their employer from participating in education organized and paid for by pharmaceutical companies.[15] If private physicians' decisions are affected by either of these mechanisms, we would expect them to be more likely to veto substitution irrespective of the patients' copayments.

Second, private physicians might have stronger pecuniary incentives to please their patients in order to keep them and/or attract new patients, which could

[14] γ_1 could differ from γ_2 because of paternalism, for example, or fear that neglecting to do what is best for the patient's health might result in an official complaint.

[15] Andréa Mannberg and Mikael Lindberg brought this possibility to my attention at a seminar at Umeå University.

result in a greater difference between γ_2 and γ_3 for them compared to county-employed physicians. As illustrated in the model above, this would have a larger impact on the physician's decision the lower the patients' copayments are. Thus, if this mechanism is operative, we would expect private physicians' likelihood of vetoing substitution to increase faster than that of county-employed physicians, as patients' copayments fall.

4 The empirical analysis

4.1 Data

The prescription dataset used in this study was provided by the county council of Västerbotten, Sweden. It contains all prescriptions sold in the county, or sold in other parts of Sweden to residents of the county, from January 2003 through October 2006, except for November and December, 2003, and September, 2004. Data for these three months are not available since the county council's data files for these months were damaged. Prescriptions issued before the substitution reform of October, 2002 and prescriptions of pharmaceuticals packed in patient-doses, were excluded since in these cases physicians were not asked if they opposed substitution.[16] Non-pharmaceutical prescriptions as well as prescriptions issued by others than physicians (e.g. dentist and nurses) were also excluded. Finally, after excluding nearly 270,000 observations originating from other workplaces than health centers, clinics or private practices in Västerbotten (e.g. emergencies, labs, occupational health services, or workplaces in other counties), or unknown workplaces, and 630,000 that lack data on ATC-group[17] or did not belong to any ATC-group, 5.1 million observations of pharmaceutical prescriptions remain.[18]

In 1.7 and 2.8% of the observations the physician and patient, respectively, opposed substitution. All these observations were used plus a random sample

[16] Patients with stable medication, who might have some problem keeping track of how much they should take, often receive their prescriptions in "patient-doses".

[17] In the World Health Organization's Anatomical Therapeutic Chemical (ATC) classification system, pharmaceuticals are divided into groups according to the organ or system on which they act, and their pharmacological, therapeutic, and chemical properties. In the ATC-groups used here, pharmaceuticals with the same active ingredients are grouped together.

[18] Using also observations lacking ATC-group data (including those that did not belong to any ATC-group) did not change the qualitative results.

of 2.5% of the remaining observations, resulting in a final sample of 350,180 observations. A sample had to be drawn because of limited computer-capacity for running iterative estimation procedures. Because of the low percentage of physicians opposing substitution, all those observations were used in order to reduce the variance in the logistic regressions, compared to using a random sample from the whole population, resulting in the same number of observations.[19] All observations when the patients refused substitution were used in order to minimize the effect of individual measurement-errors of the copayment variables that may exist for these observations.[20]

Some descriptive statistics are presented in Table I, where the sample is grouped based on whether the physician vetoed substitution ($V = 1$) or not ($V = 0$). The observations are weighted according to the inverse of their probability of being sampled. For the indicator-variables the percentage of observations in each category are presented. For the continuous variable Age, means and standard deviations are presented instead. The variables $Private$ and $County$ take the value one if the prescribing physician worked at a small private practice or was employed by the county, respectively, while the next three variables indicate which healthcare district ($Umeå$, $Skellefteå$ or South of $Lapland$) their workplace was located in.

The dataset includes information about the total cost of the prescription as well as the patient's copayment, from which the copayment-rate the patient had prior to paying for the current prescription was calculated (calculations are available from the author upon request). The indicator-variables $Copay100$, $Copay50$, $Copay25$, $Copay10$ and $Copay0$ show these predetermined copayment-rates. Some prescriptions are always free of charge ($Free$) for the patient and others are excluded from the insurance system ($Unsub$) irrespective of the patient's copayment bracket. The last two variables refer to the gender and age of the patient.

[19] Following Boyes, Hoffman and Low (1989) and Greene (1992) - who also over-sampled observations where the dependent variable took the value one, because of the low share of such observations in the population - I used sampling-weights in the estimations. Greene (2003, Chapter 21) describes why sampling-weights should be used to avoid bias that otherwise could arise because of choice-based sampling. In this study the same qualitative results were obtained when sampling weights were not used, with the exception of the results for $Copay10$.

[20] These measurement errors can arise since ΔP cannot be perfectly observed. Excluding these observations did not change the qualitative results.

Table I. Descriptive statistics

Variable	Sample	$V = 1$	$V = 0$
V	1.68	100	0
$Private$	7.38	9.05	7.36
$County$	92.62	90.95	92.64
$Umeå$	54.69	64.49	54.52
$Skellefteå$	28.63	22.46	28.73
$Lapland$	16.69	13.04	16.75
$Copay100$	36.05	25.17	36.24
$Copay50$	13.87	12.51	13.89
$Copay25$	11.04	11.06	11.04
$Copay10$	5.33	6.01	5.32
$Copay0$	29.59	43.60	29.35
$Unsub$	3.26	1.26	3.29
$Free$	0.85	0.39	0.86
$Women$	59.45	58.80	59.46
Age	59.17±20.19	61.93±17.58	59.12±20.23
Population size	5,112,236	85,678	5,026,558
Sample size	350,180	85,678	264,502

In addition, the dataset includes information about the prescribed pharmaceutical's ATC-code, the patient's municipality of residence, and the date when the prescription was written. Of the 883 seven-digit ATC-groups present in the sample, 276 have less than 10 observations; 334 have 10 to 100 observations; 206 have 100 to 1000 observations; and 67 have more than 1 000 observations. 36% of the prescriptions were written to inhabitants of Umeå, the county's largest municipality; 28% to inhabitants of Skellefteå; 1-5% to inhabitants of each of the county's other municipalities; and 3% to individuals not living in the county. 1-3% of the observations were issued in each of the 49 months from the substitution reform, effective in October 2002, through October 2006.

The descriptive statistics provide some support for the hypotheses tested here. First, private physicians are over-represented among the subsample where substitution was vetoed. Second, the same is true for patients' with low copayments whereas the opposite is true for those with high copayments.

The dataset described above was linked with another dataset, provided by the company IMS Sweden, that classified 50% of the prescribed pharmaceuticals as originals (here called brands).

4.2 Empirical specification

The baseline empirical specification (specification 1) is

$$\begin{aligned}\Pr(V_i &= 1) = F(a + \beta_1 Private_i + \sum_{c=1}^{4} \delta_c Copay_{ci} + \beta_2 Unsub_i + \beta_3 Free_i \\ &+ \beta_4 Women_i + \sum_{a=1}^{20} \eta_a Age_{ai} + \sum_{g=1}^{882} \kappa_g ATC_{gi} \\ &+ \sum_{m=1}^{15} \lambda_m Mun_{mi} + \sum_{d=1}^{2} \mu_d District_{di} + \sum_{q=1}^{16} \tau_q Quarter_{qi} + \epsilon_i).\end{aligned}$$

In all estimations, a maximum-likelihood logit estimator which adjusts for sampling-probability was used.[21] In addition, the error terms (ϵ_i) were allowed to be heteroskedastic and correlated within workplace units.[22,23]

Private was included to test the main hypothesis in this study, that private physicians were more inclined to veto substitution.[24] The copayment-indicators were included to test the hypothesis that moral hazard in insurance exists. What really influences physicians' decisions is probably their expectation of their patients' copayments at the end of the insurance period, since this determines the share of the cost of a veto borne by the patient. This is not observable, but those with a predetermined copayment-rate of 0% will also have a zero-rate at the end of the insurance period. The other copayment-variables are only proxies, since for example those with a predetermined copayment-rate of 25% will have a rate of 25% or lower at the end of the insurance period.

[21] The same qualitative results were obtained when a probit estimator was used instead.

[22] For county-employed physicians, the workplace unit is the health centre or clinic where they work. Private physicians are grouped together in the data to one workplace unit per healthcare district. Allowing for this correlation is important since Hellerstein (1998), Coscelli (2000) and Lundin (2000), among others, found persistence in physicians' prescription behavior.

[23] A Huber-White sandwich-estimator was used to calculate robust standard errors.

[24] In the baseline specification, prescriptions written at Dragonen's health centre after it became private were excluded since the incentives for those writing these probably differed from those of both county-employed physicians and those working at small private practices.

Predetermined copayment-rates were used in order to avoid endogeneity caused by the value of the dependent variable for observation i affecting the value of independent variables for that observation. Nevertheless, persistence can cause endogeneity. For example, a physician who previously vetoed substitution of a particular pharmaceutical for a particular patient, might be more inclined to veto substitution again the next time for the same patient and pharmaceutical. At the same time, the past decision might affect the patient's predetermined copayment-rate. To study whether this possibility affects the results, the baseline specification was also estimated on a subsample of only antibacterial drugs (ATC-group J01), since these are very seldom prescribed repeatedly to a patient. Another problem that has to be kept in mind when interpreting the results is that the copayment-variables are correlated with previous pharmaceutical expenditures.

$Unsub$ and $Free$ also reflects copayments and were therefore included. However, pharmaceuticals that are always free of charge, or always excluded from the insurance, belong to a small number of ATC-groups, with which these variables are highly correlated, so high that some ATC-indicators were excluded from the estimations due to multicollinearity. Therefore, the coefficients for $Unsub$ and $Free$ probably captured other effects besides those relating to moral hazard.

$Women$, indicator-variables for 5-year age-groups, and the ATC-indicators were used as proxies for differences in health outcome in which the prescribed pharmaceutical and the cheapest generic might result.[25] The ATC-indicators also controlled for the fact that in some ATC-groups there are no generics, so that the physicians' willingness to allow substitution had no effect.[26] Finally, the ATC-indicators controlled for heterogeneity among ATC-groups with respect to price-differences between the prescribed pharmaceutical and the cheapest available generic, ΔP. I did not directly control for ΔP since physicians might

[25] I also tested using the continues variables age and age-squared, as well as using larger age-groups, and including date of prescription as a continues variable. The specification presented was chosen over these alternatives since it had a better (lower) value on the AIC information-criterion. The qualitative results are the same regardless, including not controlling for ATC-groups.

[26] Observations from these ATC-groups were used since descriptive statistics indicated that physicians had imperfect information about which pharmaceuticals had substitutes; in this sample they vetoed substitution in 21,000 cases where no substitutes existed. However, descriptive statistics indicate that a veto against substitution was twice as likely when substitutes existed. Excluding prescriptions whit no substitutes did not change the qualitative results.

decide simultaneously which pharmaceutical to prescribe and whether or not to veto substitution. Endogeneity could therefore arise if ΔP were included in the specification.[27]

The municipality-indicators, including one variable indicating whether or not the patient lived in the county of Västerbotten, were included, together with the demographic variables, to capture socioeconomic differences among the municipalities. Differences among municipalities might be important to control for, since disproportionately many private physicians are located in the two biggest municipalities, Umeå and Skellefteå. Also, I controlled for which healthcare district the prescribing physician belonged to, and in which of the 17 quarters the prescription was written. The estimation results from the baseline specification are presented in the first column of Table II (next section).

The theoretical model suggests that the effect of a patient's copayment-rate on the probability of a veto will be stronger the higher the difference between γ_2 and γ_3, i.e., the greater the difference between how much the physician weight the patient's and insurer's costs. If the difference was higher for private than for county physicians, then interaction-terms between *Private* and the six variables reflecting patients' copayment-rates should be included. The estimation results obtained when these interaction-terms were included (specification 2) are presented in the second column of Table II.

In the first two specifications, private physicians are compared to county-employed physicians, irrespective of whether they worked at health centers or clinics. Estimation results for county-employed physicians alone indicate that those working at clinics (primarily specialists) were more inclined to veto substitution than were those working at health centers (primarily general practitioners, GPs). Nearly half of the private physicians, but less than 20% of the county-employed physicians, were GPs. It is therefore quite possible that, among private physicians, a higher share of prescriptions was written by GPs, compared to those written by county-employed physicians. But since the data does not indicate whether each individual prescription was written by a GP or not, it is not possible to compare private and county-employed GPs separately, and private and county-employed specialists separately.

If being a GP makes a physician less inclined to oppose substitution, then

[27] With the exception of $Copay10$, the qualitative results were not changed by including ΔP separately or interacted with the copayment-variables.

the odds-ratios for *Private* in the first two specifications will be underestimated, and can then be understood as lower bounds. In specification 3 (and specification 4), upper bounds were estimated by including an indicator-variable that takes the value one for prescriptions written at clinics, so that only physicians working at health centers were used as a control group for the private physicians.

The fourth column in Table II presents the results obtained by comparing physicians working at Dragonen's health center with physicians at other health centers. This specification includes an indicator-variable for all prescriptions written at Dragonen's health center and another for prescriptions written there after it became private ($Dragonen\text{'}s^{post}$). The latter was included to help test the hypothesis that private physicians were more likely to veto substitution.

4.3 Results

The estimation results in Table II are presented in terms of odds-ratios.[28] Since physicians vetoed substitution in less than 2% of the cases, the odds-ratio is approximately equal to the relative probability evaluated at the mean value of each independent variable.[29]

The point estimates from the first two specifications indicate that on average private physicians were approximately 50% more likely to oppose substitution, compared to county-employed physicians, ceteris paribus. As noted that is a lower bound. The corresponding figure for the upper bound, obtained from the last two specifications, is about 80%. The different estimates regarding *Private* are not significantly different from each other, but all are significantly different from unity at the 5% level, and thus provide clear support for the main hypothesis in this paper.

A second purpose of the study was to analyze moral hazard in insurance. The odds-ratios increase as the patients' copayment decreases. This can indicate moral hazard but, as noted, can also have other explanations. The same pattern was observed when restricting the sample to only antibacterial drugs, except

[28] The odds-ratio for an independent variable X is $[\Pr(V=1)|(X=1)/\Pr(V=0)|(X=1)]/\ [\Pr(V=1)|(X=0)/\Pr(V=0)|(X=0)]$. Hence, an odds-ratio of one means that the variable X does not affect the probability of a veto.

[29] Formally, as $\Pr(V=0)$ approaches 1, the odds-ratio for a variable X approaches the relative probability; that is, $\Pr(V=1)|(X=1)/[\Pr(V=1)|(X=0)$.

Table II. Estimation results, odds-ratio

	1	2	3	4
Private	1.50**	1.54***	1.79***	1.75***
	(0.24)	(0.23)	(0.25)	(0.25)
*Copay*50	1.18***	1.18***	1.16***	1.16***
	(0.03)	(0.03)	(0.03)	(0.03)
*Copay*25	1.31***	1.32***	1.28***	1.28***
	(0.04)	(0.04)	(0.03)	(0.04)
*Copay*10	1.39***	1.38***	1.35***	1.35***
	(0.04)	(0.04)	(0.04)	(0.04)
*Copay*0	1.97***	1.99***	1.85***	1.85***
	(0.14)	(0.15)	(0.09)	(0.09)
Unsub	1.35***	1.32***	1.34***	1.33***
	(0.13)	(0.12)	(0.13)	(0.13)
Free	2.46	0.89	0.82	0.83
	(1.90)	(0.52)	(0.44)	(0.44)
Women	1.10**	1.10**	1.13***	1.13**
	(0.05)	(0.05)	(0.04)	(0.04)
Skellefteå	0.68***	0.68***	0.77**	0.76**
	(0.09)	(0.09)	(0.09)	(0.09)
Lapland	0.53***	0.53***	0.66***	0.64***
	(0.07)	(0.07)	(0.06)	(0.06)
$Private * Copay50$		1.05		
		(0.05)		
$Private * Copay25$		0.96		
		(0.08)		
$Private * Copay10$		1.07		
		(0.05)		
$Private * Copay0$		0.91		
		(0.09)		
$Private * Unsub$		1.31*		
		(0.19)		
$Private * Free$		1.32		
		(0.73)		
Clinic			1.63***	1.58***
			(0.23)	(0.22)
Dragonen's				0.74***
				(0.04)
$Dragonen's^{post}$				1.51***
				(0.11)
AIC	50,891	50,889	50,668	51,077
Pseudo R^2	0.1434	0.1434	0.1471	0.1479
Sample size	346,381	346,384	346,384	349,073

Notes: The asterisks ***, ** and * denote significance at the 1%, 5% and 10% levels.
Robust standard errors are shown in parentheses.
Estimation results for age-, ATC-groups, municipalities and quarter of prescription are suppressed in order to save space, but are available from the author upon request.

that the odds-ratio for $Copay25$ was below unity, though not significantly so (odds-ratio 0.93; std. err. 0.22). Because of increased standard error, the only copayment estimate that was found to be significantly different from unity was that for $Copay0$ (odds-ratio 2.39; std. err. 0.51).[30] Thus at least the results regarding copayments do not seem to be driven solely by persistence in the physicians' prescription decisions. However, one cannot conclude whether the results are driven by previous pharmaceutical expenditures and/or moral hazard. An argument for the former is that high pharmaceutical expenditures are probably positively correlated with the number of different pharmaceuticals a patient consumes, and that a high number of pharmaceuticals can be a valid reason for a physician to veto substitution, e.g., due to the risk that the patient otherwise confuse the drugs. On the other hand, the nearly linear relationship with the patients' copayments that the estimates for $Copay50$, $Copay25$ and $Copay10$ show suggest that the results might be driven by moral hazard. That this pattern is broken by the high point estimates for $Copay0$ can be explained by less measurement-error for that variable, and hence less attenuation.

As mentioned, the variables $Unsub$ and $Free$ are highly collinear with several ATC-groups, and the estimates for these variables should therefore be interpreted with caution. The results suggest that physicians were more inclined to veto substitution for pharmaceuticals which were always unsubsidized, compared to other pharmaceuticals where the patients' copayments were 100%. The odds-ratios for pharmaceuticals that were always free of charge were unstable and not significantly different from unity. This is probably due to the collinearity problem.

That the patient was a woman was found to increase the probability of a veto. Physicians in the healthcare districts of Skellefteå and South of Lapland were less inclined to veto substitution than those in the omitted healthcare district (Umeå). Estimation results for age-, ATC-groups, municipalities and quarter of prescription are not reported in order to save space, but are available from the author upon request. A Wald test (not reported) shows that these groups of variables had significant effects.

Among the interaction-variables included in the second specification, only the interaction with $Unsub$ was significantly different from unity.[31] Thus the

[30] The estimates for Copay50 and Copay10 are 1.16 (0.21) and 1.31 (0.31). Full results from this sample are available from the author upon request.

[31] The differentials of the odds ratios with respect to the variable *Private* and the six

results do not indicate any differences between private and county-employed physicians regarding the degree to which they internalize patients' costs relative to the insurer's costs. Hence, the fact that private physicians were more likely to veto substitution does not seem to depend on them wanting to please their patients. Rather, the results suggest that the difference between the two physician groups can be explained by differences in direct costs associated with substitution. Estimation results (not reported) show that the difference between private and county-employed physicians' likeliness of vetoing substitution was approximately five times higher when brand-name pharmaceuticals were prescribed, compared to non-brand name ones. This indicates that a large part of the difference between the two physician groups might be explained by private physicians having stronger brand-name loyalty.

The results from the third and fourth specifications clearly show that substitution was more likely to be vetoed if the prescription was written at a clinic instead of a health center. As mentioned, the point estimates for $Private$ became larger when controlling for $Clinic$. The estimated odds-ratios for the healthcare districts became closer to unity, which makes sense since a disproportionably high share of the prescriptions originating from the omitted healthcare district ($Umeå$) were written at clinics. Controlling for $Clinic$ also resulted in slightly lower odds-ratios for the copayment variables.

The hypothesis that private physicians were more inclined to veto substitution was given further support by the results regarding Dragonen's health center, reported at the bottom of the last column. Physicians there became approximately 50% more likely to veto substitution when the center became private.[32] However, even though nearly 3000 prescriptions in the sample originated from this health center after it became private, it is still only one health center. Thus the pattern found is only the result from one case study. An advantage of studying this center though is that the data includes observations from

variables reflecting patients' copayments, are $Private$: 1.52 (0.24); $Copay50$: 1.18 (0.03); $Copay25$: 1.31 (0.04); $Copay10$: 1.39 (0.04); $Copay0$: 1.97 (0.14); $Unsub$: 1.35 (0.12); $Free$: 0.91 (0.52).

[32] Since physicians working at this health centre knew before June 2006 that it would be privatized, it is possible that they started to adjust to the reform before that date. Therefore specification 4 was estimated excluding observations written at Dragonen's health center between February 2006 - when the contract regarding privatization of the health centre was signed - and June 2006. This did not change the qualitative results, but the estimate for $Dragonen's^{post}$ became slightly larger, 1.57 (0.10).

both before and after it was privatized. Therefore time invariant heterogeneity regarding the health center can be controlled for.

5 Discussion

The importance of the form of compensation that physicians receive and the presence of moral hazard in insurance were analyzed by studying the determinants for whether physicians vetoed substitution or not.

The primary purpose was to test if physicians working at private practices were more likely to oppose substitution than county-employed physicians working on salary. It was found that private physicians were indeed more likely to veto substitution. Depending of how the control group was specified private physicians were estimated to be 50-80% more likely to veto substitution. Also the results show that physicians working at Dragonen's health center became approximately 50% more likely to veto substitution when the center became private.

The difference in the likeliness of private and county-employed physicians vetoing substitution was not significantly affected by patients' copayment-rates. This suggests that the observed difference between the two physician groups was caused by differences in direct costs associated with substitution, rather than private physicians being more inclined to please their patients in order to secure a high number of patient-visits. There could be such a difference if, as seems possible, private physicians have stronger brand loyalty. Allowing substitution might also be time-consuming for the physician, if it worries the patient. Hence, it could reduce the number of patient-visits per day, which would be more costly for private physicians, since their income depends on that number.

Since a physician can choose whether or not to work privately, it cannot be ruled out that the pattern found was caused by selection: The physicians who chose to work privately might have differed systematically from those that did not, for example, they might have had stronger brand-name loyalty already before becoming private physicians. Similarly, patients that chose to visit private physicians might have had systematic unobserved differences from those that did not.

A second purpose was to analyze if moral hazard in insurance affected the physicians' decisions, that is, if physicians internalized costs borne by their patients more than costs borne by the insurance. The results are consistent with that moral hazard affected the physicians' decisions, and the point estimates imply that physicians were nearly twice as likely to oppose substitution if all costs were borne by the insurance rather than by the patient. Thus physicians appeared to act more as agents for their patients than for the insurer. The patients' copayment-rates are a function of their previous pharmaceutical expenditures, however, so it cannot be ruled out that the results were caused, for example, by physicians being more likely to veto substitution the more pharmaceuticals a patient was using.

A veto against substitution not only leads to higher cost for the current prescription but also risks reducing price-competition between pharmaceutical firms. Therefore these results are important to consider when designing physicians' contracts, and perhaps also when designing pharmaceutical insurance. However, more research is needed, especially regarding moral hazard in insurance, preferably based on data where patients can be followed over time so that persistence in pharmaceutical consumption can be studied and the number of pharmaceuticals a patient consumes can be controlled for. Further research about physicians' compensation should preferably be based on data where the share of private physicians is largely affected by policy changes, so that selection effects can be separated from treatment effects.

References

Blomqvist, Å. 1991. The doctor as double agent: Information asymmetry, health insurance, and medical care. *Journal of Health Economics* **10**: 411-432.

Boyes, W., D. Hoffman and S. Low 1989. An econometric analysis of the bank credit scoring problem. *Journal of Econometrics* **40**: 3-14.

Coscelli, A. 2000. The importance of doctors' and patients' preferences in prescription drug markets. *Journal of Industrial Economics* **48**: 349-369.

Crown, W.H., E.R. Berndt, O. Baser, S.N. Finkelstein, W.P. Witt, J. Maguire and K.E. Haver 2004. Benefit plan design and prescription drug utilization among asthmatics: Do patient copayments matter? *Frontiers In Health Policy Research* **7**: 95-127.

Dusheiko, M., H. Gravelle, R. Jacobs and P. Smith 2006. The effect of financial incentives on gatekeeping doctors: Evidence from a natural experiment. *Journal of Health Economics* **25**: 447-478.

Gosden, T., L. Pedersen and D. Torgerson 1999. How should we pay the doctors? A systematic review of salary payments and their effect on doctor behaviour. *Quarterly Journal of Medicine* **92**: 47-55.

Granlund, D. and N. Rudholm 2007. Consumer information and pharmaceutical prices: Theory and evidence. Umeå Economic Studies 709, Umeå University.

Granlund, D., N. Rudholm and M. Wikström 2006. Fixed budgets as a cost containment measure for pharmaceuticals. *The European Journal of Health Economics* **7**: 37-45.

Greene, W.H. 2003. A statistical model for credit scoring. Working Paper EC-92-29, New York University.

Greene, W.H. 2003. *Econometric analysis*, fifth edition. Prentice Hall: New York.

Hellerstein, J.K. 1998. The importance of the physician in the generic versus trade-name prescription decision. *RAND Journal of Economics* **29**: 108-136.

Lag 2002:160 om läkemedelsförmåner m.m. [Act 2002:160 on Pharmaceutical benefits, etc.] (in Swedish).

Leibowitz, A., W.G. Manning and J.P. Newhouse 1985. The demand for prescription drugs as a function of cost-sharing. *Social Science and Medicine* **21**: 1063-1069.

Lundin, D. 2000. Moral hazard in physician prescription behavior. *Journal of Health Economics* **19**: 639-662.

Nassiri, A. and L. Rochaix 2006. Revisiting physicians' financial incentives in Quebec: A panel system approach. *Health Economics* **15**: 49-64.

Pauly, M.V. 1968. The economics of moral hazard: comment. *The American Economic Review* **58**: 531-537.

Ross, S.A. 1973. The economic theory of agency: The principal's problem. *The American Economic Review* **63**: 134-139

Rudholm, N. 2005. Pharmaceutical insurance and the demand for prescription pharmaceuticals in Västerbotten, Sweden. *Scandinavian Journal of Public Health* **33**: 50-56

Shortell, S.M. 1998. Physicians as double agents: Maintaining trust in an era of multiple accountabilities. *Journal of the American Medical Association* **280**: 1102-1108.