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Measuring annual price elasticities in Dutch health insurance: A new method

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Abstract in English

This paper proposes a new method for estimating annual price elasticities from market share data of health insurers. In contrast to traditional methods the elasticity is derived from bilateral price elasticities which relate the net share of switchers between two health insurers not only to their premium difference but also to the market share and premium of the higher priced health insurer. Our new method explains the annual variation in the Dutch market share data better than the traditional methods. We find in the Dutch social health insurance for the period 1996-2005 rather low negative annual price elasticities ranging between −1 and 0. In that period stickiness of insurer choices was high and less than 5% of the population switched annually from health insurer. This result, however, was in sharp contrast with an exceptional high price elasticity of −7 for the year 2006, where after a major health care reform about 18% of the population switched mostly to lower priced health insurers. Besides large media coverage, one important difference with previous years was that many consumers holding an individual contract could switch to a lower priced group contract.

Key words: health plan choice, premium elasticities, switching costs
JEL code: D12, I11, I18, L11

Abstract in Dutch

Deze studie presenteert een nieuwe methode voor het schatten van jaarlijkse prijselasticiteiten in de zorgverzekeringsmarkt. Bij de nieuwe methode worden bilaterale prijselasticiteiten geschat waarbij het aantal overstappers tussen twee zorgverzekeraars niet alleen wordt gerelateerd aan hun premieverschil, maar ook aan het marktaandeel en de premie van de duurdere zorgverzekeraar. De nieuwe methode kan de variatie in de marktaandelendata beter verklaren dan de traditionele methoden. In de ziekenfondsmarkt vinden we voor de periode 1996-2005 lage jaarlijkse negatieve korte-termijnprijselasticiteiten van tussen de −1 en 0. Veel verzekerden bleven trouw aan hun zorgverzekeraar en minder dan 5% van de populatie wisselde jaarlijks van zorgverzekeraar. Voor het jaar 2006 vinden we daarentegen een uitzonderlijke hoge prijselasticiteit van −7. Tijdens dit jaar van de stelselherziening switchte ongeveer 18% van de bevolking naar veelal goedkopere zorgverzekeraars. Naast sterke media-aandacht, speelde ook een belangrijke rol dat veel verzekerden met een individueel contract konden kiezen voor een goedkoper collectief contract.

Steekwoorden: Zorgverzekeringen, prijselasticiteit, zoekkosten

Contents

Summary

In many countries, freedom of choice and premium competition have been introduced in health insurance. Consumers may choose their favourite health insurer on the basis of various characteristics such as price and quality. The idea is that this pressure by consumers will provide incentives to health insurers to control costs and to increase quality. The success of competition will depend for an important part on the willingness of consumers to switch from health insurer in response to a change of their premium or quality. Indicators that measure the propensity of consumers to switch are called elasticities. In our context, a premium or price elasticity measures the effect of differences in health insurers nominal premiums on health insurer choice for the basic benefit package. A quality elasticity measures the effect of differences in quality on this choice.

In this study we present a new method to measure elasticities. In contrast to traditional methods we derive the price elasticity from bilateral price elasticities which relate the net share of switchers between two health insurers not only to their premium difference but also to the market share and premium of the higher priced health insurer. Our new method explains the annual variation in the Dutch market share data better than the more traditional estimation methods

In general, in health insurance many consumers are indifferent, or find search costs too high, to actually switch from health insurer. This general notion seemed also present in the Dutch social health insurance. While during 1996-2002 the population stayed for the basic benefit package on average at lower priced health insurers, in 2003, after a sudden change of premium setting by health insurers, the population stayed on average at higher priced health insurers. Dutch consumers did not seem to respond very strongly to this sudden change in premium setting by health insurers. Indeed, during the period 1996-2005 stickiness of insurer choices was rather high and less than 5% of the population switched annually from health insurer. This corresponds with the rather low negative annual price elasticities, ranging between -1 and 0, that we find in the Dutch social health insurance for the period 1995-2005.

Although, generally, consumers may not be very price sensitive, one important message of our study is that the price sensitivity of consumers can be increased. We show that the Dutch population was very price sensitive after a major health care reform and found an exceptionally high premium elasticity of -7 for the year 2006 and -2 for the year 2007, while comparable figures for the years 1997-2005, in the Dutch social health insurance without large reforms, were around $-0{,}6$. These high elasticities correspond with the large switching gains that we observed. The total population gained 130 million euros in 2007 and 32 million euros in 2007, while the total annual switching gains in the social health insurance did never rise above 7 million euros in previous years.

Through the reforms, combined with massive media coverage of premium differences, the population seemed to have become much more aware about their possibilities to switch health

plans. One important difference with previous years was that many consumers with an individual contract could opt for a lower priced group contract (sometimes with the same insurer). Many consumers, probably those that were most price sensitive, used this opportunity and the share of group contracts in the Dutch population increased from 38% in 2005 to 53% in 2006 and 56% in 2007. Also, before the reform bad risks were often locked in to an insurance contract, but after the reform were suddenly free to choose their own health insurer. The decline of the price elasticity to -2 in 2007 suggests a return to lower future price-elasticity values and suggests that the high price elasticity in 2006 was a once-only event.

As from 2006 consumers could also take “quality” information about insurers service levels into account. The estimation results indicate an insignificant insurers’ service level elasticity for 2006 , but a significant service elasticity of around 1 for 2007. This may suggest that it requires some time before consumers become accustomed to new information.

Our analysis contains various limitations. Due to a lack of data, we could not consider that more than 90% of the population bought also some form of supplementary insurance. Also, measuring the price sensitivity in 2006 and 2007 more precise would require data about the possible premium offers that groups receive from health insurers.

Data on the quality of health care provision is still lacking in the Dutch health insurance. However, in the last few years the government and also players in the health care market are undertaking more and more efforts to increase transparency with respect to these variables. If, finally, quality information will become available for consumers, than again, this may inspire consumers to reconsider their choice of health insurer and this may again result in a large number of switchers.

1 Introduction

In many countries, freedom of choice and premium competition has been introduced in health insurance. Consumers may choose their favourite health insurers on the basis of various characteristics such as price and quality. The idea is that this pressure by consumers will provide incentives to health insurers to control costs and to increase quality. The success of competition will depend largely on the willingness of consumers to switch from health insurer in response to a change of their premium or quality. Indicators that measure consumers' switching propensity are called elasticities. In our context, a premium or price elasticity measures the effect of differences in health insurers nominal premiums on health insurer choice for the basic benefit package. A quality elasticity measures the effect of differences in quality on this choice.

From the health insurance literature follows that the price (and quality) elasticity depends on switching costs (Buchmueller, 2006, 2006a). Consumers do not tend to switch very often from health insurer because of high transaction costs or the uncertainty about the alternative health insurers. In a managed care environment, where health insurers and health care providers are vertically integrated, switching costs may be even higher since switching from health insurer may require switching from health care provider as well.

There is now a fast expanding literature on measuring price elasticities in health insurance. In the U.S. the literature consists mainly of case studies. Some examples are Cutler and Reber (1998), who study a health insurance reform carried out at the Harvard University and report a price elasticity of ─2. A strong effect of price on switching was also found by Buchmueller and Feldstein (1997), during a reform at the University of California, and by Royalty and Solomon (1999) during a reform at Stanford University. This latter study reports price elasticities that ranged from ─1 to ─6. Although these three "university" experiments are limited to certain geographical areas and to a certain part of the U.S. population, they typically find a high responsiveness of consumers after a reform.

There are reasons to expect that the response of consumers may be different in a more regular setting when consumers have to react annually to changes in premiums. Studies that provide estimates based on national population data and a large number of health insurers report lower price elasticities. Most of these studies regress health insurers' market shares on out-of-pocket premiums. For example, Dowd, Feldman and Coulam (2003) study the response of Medicare beneficiaries in the United States in 1999. They report low price elasticities but argue that market share losses associated with small changes in a health insurers' premium, relative to its competitors, may be sufficient to discipline premiums in a competitive market. More time periods are considered in the Dutch social health insurance. Schut and Hassink (2002), Schut, Gress and Wasem (2003) and Van Dijk et al. (2006) estimate price elasticities over a somewhat longer period, without major reforms, and find also lower price elasticities, that ranged from ─0.1 to ─0.4 for compulsory coverage. Schut, Gress and Wasem (2003) and Tamm et. al.

(2007) provide substantially higher price elasticities for the German health insurance market, but these price elasticities are lower than the above reported price elasticities after a reform. A health insurance reform is likely to activate people to switch quicker.

There is also a smaller but growing literature on consumer responses to health insurer quality information. Most studies seem to find that quality has an effect on choice. For example, Wedig and Tai-Seale (2002), Beaulieu (2002) and Jin and Sorensen (2006) find that consumers respond significantly to quality differences when quality measures were introduced. However this evidence is not supported by Abraham et al. (2006), who do not find a link between quality information and switching behaviour. They find that switching is influenced by changes in premiums and whether an individual has an existing relationship with a health care provider.

Most studies in the health plan choice literature choose the multinomial discrete choice logit model as starting point for estimation (McFadden, 1974). The discrete choice model follows from an equilibrium framework were all consumers have full information on all observable choice factors of health insurers, such as price and quality. The consumer now deliberately chooses each year that health insurer that maximizes his/her expected utility. Although these models have some convenient properties there are also some limitations (Dowd and Feldman, 2006 and Train, 2003). The underlying assumption in these models that all consumers deliberately choose their health insurer can be challenged. While this is probably true for switchers the main question is whether non-switchers stayed deliberately at their health insurer. Most non-switchers probably not. This possibility is consistent with the common fact that very few consumers annually switch from health insurer. Evidence of status quo bias is presented by Strombom et. al. (2001) and Schut et. al. (2003) who show that premium elasticities for new entrants are higher than for incumbent enrolees. New entrants often must choose a health insurer and stickiness of insurer choice plays only a minor role. Samuelson and Zeckhauser (1988) find similar evidence and challenge the presumption that discrete choice models always provide a valid descriptive model in these kind of markets.

In this paper we also find that stickiness of consumer choice is large. This finding has some important implications for the econometric analysis. In the case of stickiness we find that it is more appropriate to estimate a price elasticity by a dynamic model that explains *changes* in market shares by the level of premiums than a static model that explains *levels* in market shares by the level of premiums. These result are also found by Tamm et. al. (2007).

A new feature of our method is that we explain changes in market shares by differences in premiums (or quality) on a bilateral basis, in which we assume that generally consumers will switch from higher to lower priced health insurers. This implies asymmetry, since it matters for the price elasticity whether we observe hundred price-switchers from a large to a small insurer than from a small to large insurer. In the first case we consider the potential number of switchers to be larger which has a downward effect on the price elasticity.

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The paper is organised as follows. In section two, we present our new method for computing price and quality elasticities. In section three, we apply our method on data from the Dutch social health insurance market over the years 1997-2005. In section four, we study the effect of the Dutch health care reform on switching behaviour and compute price and quality elasticities for the year 2006 and 2007. Section five concludes.

2 A new method for deriving an annual price elasticity

In the health insurance market each year a lot of changes may take place that may have an impact on the price elasticity. At the health insurance market, new health insurers may enter, and existing health insurers may leave or merge. Important are also changes in the rules of the game that are imposed by the government, for example in the Dutch health social health insurance the government increased annually the ratio of the premium contribution paid by the employer and the consumer, or undertook policies that changed annually the population eligible for insurance. Moreover, in 2006 the Dutch government implemented huge market reforms in the health insurance market to increase efficiency. All these changes in the market may have an impact on the price elasticity. This may be important for policymakers that want to know whether the concept of managed competition works and whether policy should be adjusted.

A common approach to study health plan choice that use aggregate or market share data are cross sectional models that are derived from the conditional logit model of McFadden (1974). An important assumption behind these models is that all consumers choose their health insurer deliberately. In the health insurance market however the overwhelming majority of the population does not switch from health insurer and stays at their current insurer. An important question is whether these non-switchers stayed at their health insurer “deliberately”, as can be explained by available price or quality information or that other “unobservable factors” play a role, such as status quo bias, high transaction costs, past historical, cultural reasons or mergers. In the Dutch social health insurance the latter reasons seem to dominate given the fact that very few consumers even considered switching from health insurer (Laske-Aldershof and Schut, 2005).

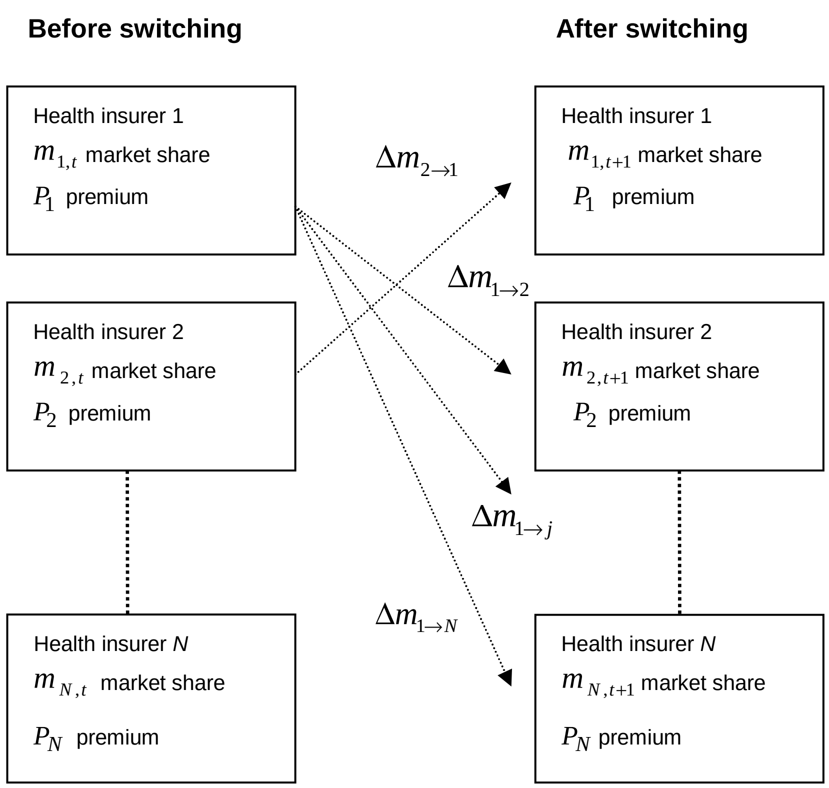
Another questionable feature of most traditional models is that the absolute change of an insurers’ market share is independent of the fact whether an insurer raises or lowers its premium by the same amount. As far as we know, it is not clear whether this property is empirically valid. A health insurer lowering its premium may attract new enrolees from other health insurers but this potential number of enrolees may be quite different compared to an health insurer raising its premium, since raising the premium will only affect its own number of enrolees .

In the next subsection we present our new method for computing *annual* price elasticities from annual data on market shares, premiums and quality indicators of all health insurers in the market. Next we will compare our method with the more traditional estimation methods. We will show that our methods yields a better explanation of the variation in the market share data than the traditional methods. Finally in the last subsection we show how our method can be extend if the researcher has the availability of additional explanatory variables.

2.1 Single variable case

In this section, we first present our method for the single variable case; i.e. consumers base their decision to switch from health insurer on only one characteristic that is the price. Since health insurance is in most countries mandatory and the possibility for switching occurs annually, we assume that a consumer stays at a certain health insurer and receives each year a new offer from his or her current health insurer. The consumer now may, or may not, compare this offer with all other offers in the market and eventually may decide to switch to another health plan. A graphical illustration of the state and flow variables of our model is presented in figure 2.1.

Figure 2.1 A graphical illustration of the switching model



The model contains the following variables ($i, j = 1,...N$):

$m_{i,t}$, market share of health insurer i at time t (before switching).

$m_{i,t+1}$, market share of health insurer i at time t+1 (after switching).

$\Delta m_{i\to j}$, net share (related to total market) of switchers from health insurer i to j ($i \neq j$).

P_i , $i = 1,...,N$: premium or price of health insurer i .

Since we allow in this single variable case only for differences in price, we can model for each of the $N(N-1)$ flows of net shares $\Delta m_{i\to j}$ in figure 2.1 a bilateral price elasticity. A consistent model should satisfy the property $\Delta m_{i\to j} = -\Delta m_{j\to i}$ and to model a price elasticity we need to know which consumers decide to switch on basis of what price differences. Therefore we assume that consumers flow from a higher priced health insurer to a lower priced health insurer in the following way:

$$\Delta m_{i\to j} = \begin{cases} \varepsilon_{i,j} m_{i,t} \dfrac{P_j - P_i}{P_i}, & \text{for } P_j < P_i,\ i,j = 1,\dots,N,\ i \neq j \\[2ex] \varepsilon_{i,j} m_{j,t} \dfrac{P_j - P_i}{P_j} & \text{for } P_j > P_i,\ i,j = 1,\dots,N,\ i \neq j \end{cases} \tag{1}$$

In this equation $\varepsilon_{i,j}$ can be directly interpreted as a bilateral premium elasticity. The first equation in (1) states that if $P_j < P_i$ then the number of switchers moving out of health plan *i* will depend on the size of the higher priced health plan. If on the other hand, $P_j > P_i$ then the amount of switchers will depend on the size of the higher priced health plan *j*. Note that our specification satisfies the property $\Delta m_{i\to j} = -\Delta m_{j\to i}$.

Note that we assume that there exists no equilibrium of market shares. Only the changes in market shares can be explained by the model but not their levels. So market shares are not necessarily stable and are determined by price differences in the past. A dynamic model is probably a better way of modelling price elasticities than the traditional static models.[1]

There are two practical limitations with this approach. First, estimating the different price elasticities in equation would require data of all bilateral consumer flows among all health insurers. This data is often not available. Second, our main interest is not to obtain bilateral price elasticities but an average price elasticity of the total market. We therefore impose that $\varepsilon_{i,j} = \varepsilon$ for *i*, *j* =1,...*N* and define, ε, as the *average bilateral price elasticity*. Since, in general $\varepsilon_{i,j} \neq \varepsilon$ we have to estimate ε, therefore we combine the $N(N-1)$ equations in (1) to obtain a set of equations that can be expressed in market share data:

$$\sum_{j=1}^{n} \Delta m_{i\to j} = m_{i,t+1} - m_{i,t} = \varepsilon\ S_{i,t}^{P} + e_{i,t+1}, \quad i=1,\dots N, \tag{2}$$

$$\text{where } S_{i,t}^{P} = \sum_{j,P_j<P_i}^{N} m_{it} \frac{(P_j - P_i)}{P_i} + \sum_{j,P_j>P_i}^{N} m_{jt} \frac{(P_j - P_i)}{P_j}$$

In (2) we have *N* equations, explaining the difference in market share of each health insurer *i*, and relate this to $S_{i,t}^{P}$. The set of equations in (2) can be estimated using standard estimation procedures. Since

[1] Tamm et al. (2007) found strong indications that market shares in Germany follow almost a unit-root process. In the appendix, we show that this result also seems to hold for the Netherlands.

$$\sum_{i=1}^{N} S^{P}_{i,t} = 0 \quad \text{for all } t \tag{3}$$

we define $S^P_{i,t}$ as the *relative* net share of premium switchers between two health insurers in the market. After estimating the *average bilateral price elasticity* $\hat{\varepsilon}$ we can compute for each health insurer *i*: $\hat{\varepsilon}\ S^P_{i,t}$,the *estimated* net share of switchers between two health insurers. Note that it is fairly easy to construct $S^P_{i,t}$ in practice since this requires only data on market shares and premiums of all health insurers.

From the *average bilateral price elasticity* we derive the average *price elasticity of the total health insurance market*, denoted by η. In the literature the premium elasticity of health insurer choice is defined as the percentage change in the *j*th health insurers' market share associated with a one percent change in its own nominal or out-of-pocket premium (see e.g. Dowd, Feldman and Coulam, 2003). In our new method the bilateral price elasticity depends not only on the change in premiums but also on the ranking position of the insurers' premium and the market shares of the higher priced health insurers. This complicates the definition for an average price elasticity of the total insurance market. Therefore, we derive the *price elasticity of the total health insurance market* from an equilibrium situation in which all premiums and market shares are equal and calculate the effect of a percentage change of the market share for an average insurer (with market share $1/N$) associated with a 1% change in its own premium. In that case the effect on the market share (in absolute terms) is the same whether one raises or lowers the premium by one percent.[2] A health insurer with market share $1/N$ that raises or lowers its premium by 1%, keeping everything else constant, would loose or attract from all its competitors $\hat{\varepsilon}$ $(1-1/N)$% = $\hat{\varepsilon}$ $(N-1)1/N$ % consumers. Therefore we define the average *price elasticity of the total health insurance market* by $\eta = \hat{\varepsilon}\,(N-1)$. We use this elasticity as an indicator for the price elasticity of the total market. It reflects the percentage change in market share of an average insurer, namely η%, associated with a 1% change in premiums.

In the next sections we will use the above method to estimate the *annual price elasticity of the total health insurance market.* First, however we will show that, at least for the Netherlands, our estimator explains the variation in the data better than price elasticities that are calculated with the standard approaches used in the literature.

2.2 A comparison with more traditional estimators

Whether our new method explains the annual variation in the market share data better than the more traditional methods can only be tested by comparing our model in equation (2) with the

[2] Even in that case it is not strictly true. It would be strictly true if the nominator in the first summation in equation (2) would be P_j in stead of P_i. However, in practice these differences are marginal. More fundamental, however, is the conceptual problem whether even in the equilibrium situation raising the premium should have a similar effect on the market shares of health insurers (in absolute terms) than lowering the premium.

more traditional methods that are used in the literature. We consider the following three traditional cross sectional models (i=1,...,N):

Model 1: Regressing first differences of market shares on premium levels:

$$m_{i,t+1} - m_{i,t} = \alpha_d + \delta_d P_i + \varepsilon_i \,,$$

Model 2: Regressing growth in market shares on premium levels:

$$\frac{m_{i,t+1} - m_{i,t}}{m_{i,t}} = \alpha_g + \delta_g P_i + \varepsilon_i \,,$$

Model 3: Regressing levels of market shares on premium levels

$$m_{i,t+1} = \alpha_l + \delta_l P_i + \varepsilon_i \,,$$

All three cross sectional models explain the growth, change or level in market share by their own premium (in relation to the average premium in the market), and therefore these models do not take into account the possible asymmetries we included in our model. Using our data, we can now estimate and compare all four models.[3] A comparison of the models can be done by using standard model-selection criteria such as R^2 or loglikelihood.[4] Although we do not go into detail here, in Appendix A we show that our model of equation (2) outperforms in terms of standard selection criteria all other three traditional models. We will present here a summary of our results and refer to Appendix A for a more extensive report.

We find that the traditional model 2 and 3 perform bad, while model 1 sometimes comes close to our model in terms of goodness-of-fit criteria.

The problem with a growth specification as in model 2 is that it may create outliers in the dependent variable. Especially small health insurers with a large change in their market share may have large growth rates. Annual price elasticities that follow from growth models seem not very robust and may exhibit very large differences between two consecutive years. Differences which one would, a priori, not expect such.

Researchers often regressing market share levels on premiums, as in traditional model 3, since this model follows from a conditional logit specification (see e.g. Dowd and Feldman, 2006). However, for our dataset traditional model 3 performs very bad with annual R^2's of lower than 0,1 while our model generates R^2's of above 0,99! The likely reason is that an important assumption behind the traditional model 3, that all consumers choose deliberately their health insurer, is not correct. This suggest that stickiness of health insurer choice is very large and that a form of stickiness should be included in the model when estimating a price elasticity. A way to capture stickiness of health insurer choice is to use fixed insurer effects.

[3] We obtained the premium and market share data of the Dutch social health insurance for the years 1995-2005 from the Dutch Health insurance board (CVZ). For the years 2006-2007 the data premium and market share data is obtained from the Dutch healthcare authority that processed raw survey data from VEKTIS (a Dutch information company on insurer data).

[4] Since each model has the same number of explanatory variables, goodness-of-fit measures as Akaike's or Schwartz Bayesian information criterion do not provide additional information.

However, in appendix A we show that including only fixed effects to the model is not enough and that it is essential to include the lagged market share variable to the model.[5]

A form of stickiness is included in the traditional model 1 by including lagged market shares. The result in Appendix A show that this improves the results considerably. The estimated price elasticities are relatively stable and close to the elasticities we derived with our new method. However, the goodness-of-fit criteria of the traditional model 1 are always, and sometimes substantially, smaller than our new model. Thus, we find that including $S_{i,t}^{P}$ as explanatory variable yields better results than including the premiums level $P_{i,t}$. In Appendix A we show that this result still holds, albeit the improvement is smaller, in fixed effect panel data models.

Summarizing, we conclude that our model as specified in equation (2) is preferred over other models and that asymmetry in the estimation process may yield better results. Therefore we will only use in the remaining part of this study the estimation result of equation (2) .

2.3 Multi variable case

Besides price, there may be various reasons for consumers to switch from health insurer. In this section we extend our estimation method to the multi variable case. The extension is straightforward. For each additional variable that one wants to include in the estimation process one first should ask the question: How do we expect that consumers will react on information about this new variable? In the case of price, everything else equal, we expect that consumers will move from a higher priced health plan to a lower priced health plan. In case of quality, everything else equal, we expect that consumers will move from a lower quality health plan to a higher quality health plan. In the latter case we first construct a new variable, just as we did for price in equation (2), but now by using the assumption that consumers move from a lower quality to a higher quality health insurer:

$$S_i^Q = \sum_{j,Q_j<Q_i}^{N} m_{it} \frac{(Q_i - Q_j)}{Q_j} + \sum_{j,Q_j>Q_i}^{N} m_{jt} \frac{(Q_i - Q_j)}{Q_i} \qquad (4)$$

We will use the above quality variable in section four.

Remains the case of the inclusion of a variable where a priori the researcher has no idea how consumers will react on information about this variable. In that case, we would not recommend our procedure since a prerequisite of our new method is pre-knowledge of the expected reaction of consumers.

[5] Tamm et. al. (2007) find the same result for the German health insurance market.

3 Application: Dutch Social Health insurance 1996-2005

Until 2006, the Dutch health insurance for basic cure services consisted of a two-pillar system. One pillar was the social health insurance system for people in the lower income brackets, and the other pillar was the voluntary private health insurance system for people with higher incomes. In this section we consider the social health insurance system that was administered by sickness funds, or not-for profit health insurers. This public scheme regulates insurance for those with labour income below a certain threshold (about 32 600 euro in 2005). Insurance is obligatory for those who are eligible, and covers about two-thirds (10 million people) of the Dutch population. Enrolees face equal basic benefit packages, as designed by the government and face two types of premiums: a basic income-related premium that is uniform across health insurers and a out-of-pocket premium. The basic premium is collected by the government and reimbursed to sickness funds after applying risk adjustment. The out-of-pocket premium accounted during the years 1996-2005 for about 10%-20% of the total premium.

In order to stimulate insurers to become more efficient, the Dutch government introduced in 1996 elements of regulated competition by allowing health insurers to set different out-of pocket premiums for basic health insurance. The out-of pocket premium is equal for all enrolees at the same health insurer (community rating). Enrolees may change yearly from health insurer, and acceptance is obligatory. Community rating may create predictable losses for health insurers on enrolees with predictably high medical expenditures. Therefore the Dutch government started in 1991 to develop a risk adjustment system, which has subsequently been further improved (Douven, 2004). A crucial precondition for a model of regulated competition is that consumers are sensitive to price.

3.1 Descriptive statistics of the basic social health insurance

In Table 3.1, we present some basic features of the Dutch social health insurance for the years 1996-2005. The first row presents the population size and the second row the number of premium payers. Since children under the age of eighteen are not required to pay out-of-pocket premiums we base our price sensitivity calculations on the premium payers. The population size was relatively stable with a peak in the year 2000 when an additional 0,4 million people entered the social health insurance market. This number is almost completely related to a policy of the Dutch government to bring lower-income self-employed persons (and dependents), which were previously privately insured, under the mandatory social health insurance. Note that the argument of ‘status quo bias’ does not hold for this group since the self-employed were forced to choose a new health insurer.

Over the period 1996-2005 the number of health insurers decreased, particularly during after the new Millennium. Almost all insurers leaving during the sample period merged with other insurers. Mergers do not necessarily alter the price sensitivity of the market since the default

Table 3.1 Insurers characteristics of the Dutch social health insurance 1996-2005

	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
Population size of total market (millions)	9.8	9.9	9.9	9.9	10.3	10.3	10.2	10.1	10.2	10.1
Population of premium payers (millions)	7.7	7.9	8.0	8.0	8.3	8.3	8.2	8.2	8.2	8.2
Total number of health insurers	27	29	29	29	26	24	21	21	21	21
Number of insurers leaving the market	0	0	2	0	3	2	3	0	0	0
Number of insurers entering the market	1	2	2	0	0	0	0	0	0	0
Insurers with a population > 100 000	21	21	19	19	19	18	15	15	15	15

option still applies, that is enrolees can accept the offer of the merged company. On the other hand, new entrants may have an impact on the price sensitivity of enrolees since they may exert efforts, e.g. media exposure, to gain quickly a large population. During our sample five new health insurers entered the market, where four of them entered with a very small population. One health insurer entered the market in 1998 with a larger population after concluding a contract with a large company.

Since our data contains only insurers' market shares there is no exact data on the number of switchers. A survey indicates that the annual percentage of 'unforced' switchers are between 2-3% for the years 2001-2004, and around 4% in 2005 (Laske-Aldershof and Schut, 2005). For the earlier years, 1996-2000, the percentages are likely to be even lower since the general impression is that in those years most consumers were not even aware of the possibility to switch. This impression is confirmed by the same survey that indicated that the percentage of the population that did not even consider switching declined from 83% in 2001 to 77% in 2005.

In Table 3.2 we present information on basic health insurance premiums. In the first row we present for each year the unweighted mean over all health insurers. There will be a tendency towards this mean in case switchers do not consider premium as an important variable when choosing their health insurer.

The second row represents a weighted mean premium, according to the market shares of premium payers before switching. This represents the mean premium if all enrolees would opt for the default option, that is all enrolees would accept the annual offer of their current health insurer. An interesting turn around occurred after 2002. While before 2002, the weighted mean was always lower than the unweighted (random switching) mean, this changed drastically after 2002. Health insurers incurred substantial losses in 2001 and 2002, resulting in decreasing financial reserves.[6] As of 2003 all health insurers substantially raised their premiums with as result that former, before 2003, relatively low priced health insurers became relatively high priced health insurers as of 2003.

[6] For a more thorough description of this period, see Douven and Schut (2006). Note also that the Spearman's rank correlation for premiums of two consecutive years dropped to below 0.5 for the years 2002 and 2003, while the correlation was relatively stable and above 0,8 until 2002.

Table 3.2 Premiums and switching gains in the basic health insurance, 1996-2005 (in euro)

	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
Unweighted (random switching)	155.7	98.2	97.7	179.0	190.0	163.6	181.5	344.7	304.6	378.1
Weighted (before switching)	155.4	97.7	97.7	178.4	187.3	156.4	180.8	354.5	306.5	383.4
Weighted (after switching)	155.4	97.7	97.5	178.3	186.9	156.2	180.6	354.0	305.9	382.3
Standard deviation	3.7	7.8	8.4	12.6	16.2	25.6	25.7	32.8	30.9	41.0
Total switching gains (million euros)	– 0.1	0.1	0.2	0.9	1.5	1.4	1.2	4.4	3.8	7.0

The third row represents the mean premium in the market after switching. This mean is almost always lower, albeit sometimes very little, than the means in the second row of Table 3.2, which indicates that on average every year switchers opted for a lower premium. Remarkable, however, is that for the years 2003 till 2005 the weighted premium after switching is still much higher than the unweighted (random switching) mean. This suggests that consumers hardly reacted to the huge changes in premium setting.[7] This observation, combined with the fact that more than 77% of the population did not even consider to switch, may suggest high transaction costs or a large status quo bias. Thus, under the assumption that the total population chose rationally or deliberately, we may as well find a *positive* premium elasticity for the years 2003 till 2005, since compared to an annual random switching model the population paid a higher price.

The fourth row in table 3.2 represents the standard deviation in premiums. While in the year 1996 the variation in premiums was low, the variation gradually increased. This reflects for a certain part the increase in price competition in the market.[8] In general one can argue that the larger the variation in premiums the more people might switch from health insurer and the larger the total switching gains will be. This is shown in the last row of the table, where the total switching gains are reported. On average the total money gained by switchers was very low in early years of the sample but this figure increased to about 7,0 million euros in 2005. Assuming 5% switchers in the market, this means that in 2005 a switcher moved to a on average 35 euro lower priced health insurer.

[7] In a discrete choice model consumers may have stayed with the same insurer if other factors, such as supplementary insurance, quality or service, changed as well. Quality and service information was hardly available for consumers while information on supplementary insurance could have played a role, but adequate information is unfortunately unavailable.

[8] An increase in price competition may also lead to lower price variation. However, since competition was almost absent before 1996 we interpret the increase in premium variation, partly, as an increase in competition.

3.2 Short-term price elasticities in the basic social health insurance

In this section, we present and discuss the estimation results of equation (2) in section 2.1.

$$m_{i,t} - m_{i,t-1} = \alpha + \eta \frac{S^P_{i,t}}{N_t - 1} + e_{i,t}, \quad i=1,...N, \tag{5}$$

Note that we added $N_t - 1$ in the nominator of the explanatory variable $S^P_{i,t}$, which implies that η can be interpreted as the *annual price elasticity of the total health insurance market*. We run the regression in equation (5) now for each single year. Table 3.3 presents a summary of our estimation result.

The first row in table 3.3. shows the annual ordinary least square estimator for the price elasticity $\hat{\eta}$ in a given year and the second row shows the standard errors. In every year we find negative annual short term price elasticities. This is in accordance with the reported positive total switching gains in table 3.2. Although the switching gains show an upward trend, with the largest gain of 7 million euros in 2005, this trend is not fully reflected in the estimated price elasticities because the annual premium variation shows also an upward trend (see table 3.2). The table shows insignificant estimators for the first three years of the sample which corresponds with the low total switching gains and little premium variation during these years. After the year 1998 we find that the estimated price elasticities are all significant at a 5% level.

An interesting event occurred in the year 2000. In 2000 a large group of self-employed (0,38 million), which were previously privately insured, were "forced" by new legislation to enter the social health insurance market.

Table 3.3 Estimated price elasticity $\hat{\eta}$ and other characteristics

	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
Estimated price elasticity $\hat{\eta}$	–0.04	–0.08	–0.06	–0.16*	–0.25*	–0.14*	–0.06*	–0.16*	–0.20*	–0.22*
Standard error	(0.16)	(0.09)	(0.04)	(0.04)	(0.10)	(0.04)	(0.03)	(0.04)	(0.05)	(0.04)
R^2	0.00	0.03	0.06	0.42	0.22	0.42	0.18	0.46	0.49	0.66
Number of observations	27	29	29	29	26	24	21	21	21	21
$\hat{\eta} * 1000 / \overline{P}$	– 0.2	– 0.6	– 0.5	– 0.7	– 1.1	– 0.7	– 0.3	– 0.4	– 0.6	– 0.6
–1%* 400000 * $\hat{\eta} * 1000 / \overline{P}$ effect after a unilateral 1% decline in premium (number of switchers)	1000	3000	2000	3000	4000	3000	1000	2000	3000	2000

The numbers in this table are rounded off. A * indicates that the estimated price elasticities are significant at a 5% level. The numbers in the last two rows are adjusted for consumer price inflation (see footnote 9).

The fact that they were "forced" to choose a health insurer raised the price elasticity substantially. This may explain the relatively high price elasticity of −0,25 that we find for the year 2000. If we correct our price elasticity for these self-employed entering the market we find an elasticity of −0,07.[9] We also tested the effect of a relatively large health insurer that entered the market on the price elasticity. Excluding this entrant from the sample decreased the price elasticity, albeit the effect proved to be small.

Comparing price elasticities for different years (or for different countries) is not straightforward because of different base levels of premiums. If consumers would mainly choose their health insurer on the basis of percentage premium differences, and not on the basis of absolute premium differences, the price elasticities η reported in row 1 of table 3.3 would be sufficient. However, if consumers mainly choose on the basis of absolute premium differences, a comparison of these price elasticities seems less useful. In that case a comparison of the statistics in row 4 seems more appropriate, where we scaled the mean premium level $\overline{P}$ as we reported in table 3.2 for each year to 1000 euro.[10] The results imply for 2005 that a health insurer that increases its premium by 1% (10 euro), while all other health insurers held their prices constant, would loose about −0,6% of its premium payers, which corresponds with 2000 premium payers (see last row in table 3.3). The results in both last two rows seem to imply that the premium sensitivity of consumers was relatively stable over the years in the social health insurance.

Our analysis is limited in the sense that we only consider the premium for the basic health insurance. Although consumers are likely to put most emphasis on this variable, many choice variables of consumers are unobserved. If these omitted variables are correlated with our explanatory variable $S_{i,t}^{P}$ the price elasticity may be biased. Adding an insurer specific fixed effect may capture (some of) these unobserved characteristics. Therefore we constructed an unbalanced panel and tested also the following fixed effect model[11]:

$$m_{i,t} - m_{i,t-1} = \alpha_{0i} + \alpha_{0t} + \eta_t \, \frac{S_{i,t}^{P}}{N_t - 1} + \varepsilon_{i,t}\,, \qquad i=1,...,N\,,\ t=1996,...,2005 \tag{5}$$

[9] Excluding the 0,38 million self-employed from the sample, and using a price elasticity for the group of self-employed of −4, which we took from a study of Schut, Gress and Wasem (2003), would have decreased our price elasticity η in year 2000 substantially to −0,07 (=(−0,25*8,3−(−4*0,38))/(8,3-0,38)).

[10] We adjusted for consumer price inflation by using premiums that correspond with a premium of 1000 euro in the year 2006. This makes a comparison possible with the premium elasticities in 2006 that are reported in the next chapter. Consumer price inflation was about annually 2% with higher rates of around 4% in 2000 and 2001.

[11] See also appendix A.

Table 3.4	**Price elasticity estimates using a fixed effect panel data model**									
	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
Estimated price elasticity $\hat{\eta}_t$	–0.04	–0.07	–0.05	–0.06	–0.16*	–0.09*	–0.00	–0.10*	–0.14*	–0.16*
Standard error	(0.17)	(0.05)	(0.05)	(0.05)	(0.05)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)

A * indicates that the estimated price elasticities are significant at a 5% level.

The estimated fixed effect panel data model contained 248 observations and yield an R^2 of 0.75. Table 3.4 shows that the estimated price elasticities are somewhat lower than the annual estimations in the previous table 3.3. This is a standard result which indicates that the fixed effects eliminate some variation in the data that otherwise would be picked up by the explanatory variable $S_{i,t}^{P}$.

The most prominent variable that we excluded from our estimation analysis seems to be the premium for supplementary insurance, since more than 90% of the population bought some form of supplementary insurance. To control for this variable, however, we would need detailed data on the various supplementary benefit packages that health insurers offered to their enrolees. This data is unavailable. Data on health care quality and the service of the health insurer was for a long time unavailable. However in the last few years the government is undertaking more and more efforts to increase transparency with respect to these variables. In the next section we will therefore consider, next to premiums for basic health insurance, also service aspects of a health insurer.

4 Application: Dutch Health Insurance 2006-2007

For many years, Dutch health insurance for basic cure services consisted of a two-pillar system. One pillar consisted of the social health insurance system for people in the lower income brackets (see previous chapter), and the other pillar was the voluntary private health insurance system for people with higher incomes. In 2006, the Dutch government implemented radical health insurance reforms and the two pillars fused into one mandatory national health insurance system executed by private insurers. The key idea of the market reforms is to increase efficiency by promoting more competition on the health insurance market as well as on the health care provider market.

4.1 The Dutch health care reform in 2006

Introducing competition into the health care market is not without risks, as it may threaten solidarity. To preserve solidarity the government followed a setup along the lines of the Dutch social health insurance and introduced a basic benefit package that is mandatory for all Dutch citizens, community rating and a risk adjustment system. Insurers have to accept all applicants at the community rated premium. All citizens, except children under 18, must pay an income dependent contribution, levied by the tax collector, to the Health Insurance Fund (HIF). The HIF also receives contributions from the government (for example for expenditure on children under eighteen). All consumers pay also a nominal premium directly to their health insurer. While this financing system was also present in the social health insurance system before the reforms, the size of the nominal premium was in the former social health insurance system much smaller than the size of the income dependent contribution. Under the new health care system the law requires that 50% of all expenditures must be paid by income dependent contributions and 50% by nominal premiums. This implied a significant rise of the nominal premium for people in the lower income brackets from about 380 euros in 2005 (see table 3.2) to about 1050 euros in 2006. A high nominal premium should make people more aware of the high health care costs. The increase in nominal premiums would have resulted in a loss of spending power of the lower income groups without additional measures and therefore the government now compensates more than 5 million citizens with monthly income-dependent subsidies.

Health insurers obtained more tools to attract consumers than in the former social health insurance. With respect to the basic benefit package they are allowed to offer premium discounts for group contracts (the discount is capped at 10% of a similar individual contract) and a voluntary deductible that may vary between 100 and 500 euro per year. These deductibles are on top of a mandatory no-claim rebate of 255 euro per year for the entire population. Health insurers can also compete with different supplementary insurance packages, service levels and different types of preferred provider networks.

Table 4.1	**Insurers characteristics of the Dutch health insurance 2005-2007**		
	2005	2006	2007
Population size of total market (millions)	16.3	16.3	16.4
Population size of premium payers (millions)	12.5	12.5	12.6
Total number of health insurers	28	29	30
Total number of insurers that offer an individual contract	27	28	29
Total number of insurers that offer an individual contract, individual population > 100 000	17	17	15
Total number of insurers that offer a group contract	28	29	30
Total number of insurers that offer a group contract, group population > 100 000	16	18	17
Total number of insurers that price the group premium lower than the individual premium	n.a.	21	23

The first remarkable result of the reform is that health insurers started in 2006 and also 2007 a premium war. The threat that many customers would change from health insurer had a profound impact on their premium setting. In particular, premiums of group contracts were offered below the break-even price. It is estimated that in 2006 health insurers lost 565 million euros on the provision of the basic benefit package (DNB, 2007). These losses can still be managed by most health insurers since they (especially the larger ones) have substantial financial reserves.

A second unexpected result was that in 2006 about 18% of the Dutch population switched from health insurer. Such a high degree of switching was never seen before in the Dutch health insurance. Through the reforms, combined with massive media coverage of premium differences, the population seemed to have become much more aware about their possibilities to switch health plans. Many people switched from a higher priced individual contract in 2005 to a lower priced group contract in 2006, also within their own health insurer. In 2007 the percentage of the population that switched from health insurer declined to about 4,4%, but this figure was still slightly higher than we observed in the previous 10 years in the Dutch social health insurance.

Some characteristics of the health insurance market are presented in table 4.1. Although the insurance was mandatory for the total population, about 1,5% of the population remained uninsured. The table presents the number of health insurers that we used in our computations.[12]

4.2 Short-term price and quality elasticities in the basic health insurance

The calculation of price (and quality) elasticities after the reforms is more complicated than in the social health insurance since two different insurance systems merged into one system. While the social health insurance offered mainly individual contracts, in the private health insurance the majority of the population held a group contract. As of 2006 most health insurers offered

[12] This data uses survey information gathered at the end of 2005, the end of 2006 and in the beginning of 2007. The actual number of health insurers is slightly larger than reported here since sometimes large insurance concerns hold more than one label and data of two small health insurers were missing. NZA (2006) reports at the beginning of 2006, 14 insurance concerns and 33 health insurers.

both contracts and many people could opt for some group contracts. These group contracts were not only employment-based but were also offered to other groups, sometimes with a large number of potential insured such as the major labour unions, national sport federations and clients of a large cooperative bank. Group contracts were even offered to interest associations for the elderly and several groups of chronic patients (e.g. diabetes and rheumatoid arthritis). These latter contracts are feasible because health insurers are compensated for predictable expenditures by the risk adjustment system.

Although the premiums differ for an individual and a group contract, in general all enrolees receive the same quality of basic health care services. The main difference between individual and group contracts is that certain group contracts may not be not accessible for parts of the population and that group premiums may differ per group. Moreover, group premiums are often not publicly announced, so that an insurer can tune the premium to the characteristics of a group.

Since health insurers can offer different premiums for its individual and group contract, we have to modify our method for measuring premium elasticities in section 2. We therefore split in our computations each health insurer up into two health insurers. One health insurer holding all individuals contracts and the other one holding all group contracts. This division is feasible since for each single insurer we have data on individual and group market shares separately, before and after switching. The premium of the insurer holding the individual contract is the individual premium that the health insurer publicly announces. The premium of the same insurer holding the group contracts is the (weighted) average premium of all his group contracts, as we collected them afterwards. The implicit, and strong, assumption is that all consumers can switch between all different types of contracts, and thus also switch from an individual contract to a group contract within the same health insurer (or vice versa).[13]

Some statistics with respect to premiums are presented in column two and three of table 4.2. We find that in 2006 the premium offer of 1035,5 euro before switching was below the mean unweighted premium. Thus, on average the population was already assigned to lower priced health insurers. The 18% switchers moved to a on average 60 euro lower priced health insurers and the population gained 130,0 million euros through switching.[14] Compared to the Dutch health insurance in the previous section this is a incredible increase in the number of switchers and switching gains. The switching gains decreased substantially to 31,8 million euros in 2007, but the 4.4% switchers moved in 2007 also to a on average 60 euro lower price health insurer.[15] Table 4.2 shows that the weighted and unweighted premium were in both years substantially lower than their unweighted (random switching) mean. The reason is that at the start of 2007 more than half of the population had already a group contract.

[13] The assumption is strong since 70% of the group contracts are employment-based which indicates that this group is only attainable for the employees. Moreover, it is not true that a premium for some group would be attainable for other groups as well. More data and research would be needed to control for these issues.

[14] The average gain of a switcher was about 130/(18%*12,5)=60 euros in 2006.

[15] The average gain of a switcher was about 31,8/(4,4%*12,5)=60 euros in 2007.

Table 4.2	Premiums, quality and switching gains in the Dutch health insurance, 2006-2007 (in euro)			
	Premiums		Quality	
	2006	2007	2006	2007
Unweighted (random switching)	1036.8	1114.9	7572	7651
Weighted (before switching)	1035.5	1105.4	7508	7656
Weighted (after switching)	1025.3	1102.9	7516	7659
Standard deviation	45.5	48.1	0.30	0.23
Total switching gains (in million euros)	130.0	31.8	– 0.10	– 0.04

In 2006 and 2007, some quality information of insurers' service aspects were available. At a website, under responsibility of the government, each health insurer receives a mark for its service aspects.[16] These marks ranged in 2006 from 7,1 to 8,4 with a standard deviation of 0,30 and in 2007 from 7,3 to 8,1 with a standard deviation of 0,24. Note that there is no distinction made between quality for an individual and a group contract within the same health insurer. We should note that if consumers switch to a lower priced health insurer then quality may decline because premiums and quality may be positively correlated. For example, if 18% of the population in 2006 switches only for price reasons then we would observe with respect to quality a switching to the mean. In that case we would find a mark of about 7,519 which is even higher than the weighted average after switching of 7,516.[17] To capture these correlation effects we run the annual regressions with price and quality as explanatory variable:

$$m_{i,t+1} - m_{i,t} = \eta \frac{S^P_{i,t}}{N_t} + \nu \frac{S^Q_{i,t}}{N_t} + e_{i,t+1}\ , \quad i=1,...N, \tag{6}$$

In this equation S^Q is the transformed quality variable, which takes into account, besides quality differences, also the quality and market size of the *lower* quality health insurer (see equation (4)). The results of the regressions for 2006 and 2007 are reported in table 4.3.[18]

We found an exceptionally high significant premium elasticity of −7,0 in 2006 which declined to a (significant) premium elasticity of −2,0 in 2007.[19] This high elasticity indicates that if an insurer with 400000 enrolees, with an average premium, raises or lowers its premium by 10 euro that his population would decrease or increase by 27000 insured. There are many explanations possible for these premium elasticity differences. First of all, 2006 was an exceptional year with a lot of media exposure and the switching period was longer, from November 2005 until May 2006, than usual. Also the number of forced switchers may have been large. An example of the latter is that families, for which one of the partners had an

[16] The marks for service levels are obtained from the website www.kiesbeter.nl. Note that there is no information availably on the quality of health care delivery.

[17] If 18% switched randomly then this would result in a mark of about 0,82*7,508+0,18*7,572=7,519.

[18] The number of observations are lower than the number of health insurers since there were no quality marks available for four health insurers in 2006 and one health insurer in 2007.

[19] We performed some sensitivity analysis but the high premium elasticity proved to be quite robust. For example running regressions without quality, thereby increasing the number of observations, yielded an even higher price elasticity of –7,8 in 2006 and a similar price elasticity of–1,9 in 2007.

Table 4.3	**Estimated price elasticities for the year 2006 and 2007**			
	2006		2007	
	Premium	Quality	Premium	Quality
Estimated price elasticity $\hat{\eta}$	– 7.0*	0.1	– 2.0*	1.0*
Standard error	(1.1)	(1.7)	(0.4)	(0.5)
R^2		0.47		0.38
Number of observations		49		53
$\hat{\eta}*1000/\overline{P}$	– 6.8		– 1.8	
effect after a unilateral 1% decline in premium or quality (number of switchers)	27000	0	7000	4000

The numbers in the table are rounded off. A * indicates that the estimated price elasticities are significant at a 5% level. The effect in the last row are for premiums computed as 1%*400000* η_p *1000/ $\overline{P}$ and for quality as 1%*400000* η_q / $\overline{Q}$.

income above the income threshold and the other one below the threshold, before 2006 were required by law to hold separately a social and private insurance contract, sometimes offered by different health insurers, whereas after 2006 they could opt for one contract offered by a single health insurer. Another explanation is that before the reform bad risks were often locked in to an insurance contract, but after the reform were suddenly free to choose their own health insurer. It is also likely that people do not switch every year so that especially the first year after a reform more people than normally will reconsider their choice. Also survey information suggests that in the first year of the reform many people considered the possibility to switch from health insurer. For example, in the month January 2006, 68% of the population, stated in a survey that they considered to switch from health insurer in 2006. One year later, in the month December 2006 for the year 2007, this percentage already declined to 15%, while 70% of the population did not consider to switch from health insurer at all in 2007 (Vektis, 2007). This also suggests that "status quo bias" played only a minor role during the first year of the reform.

Our estimation results in table 4.3 do not provide evidence that quality was an important indicator for switchers in 2006. For 2007 the effect of quality on switching behaviour might be more prominent since we find a significant quality elasticity of 1. This indicates that a 0,1 change in quality mark, for an insurer with 0,4 million enrolees with an average quality mark of 7,5, is associated with a change of about 4000 consumers.

4.3 Price sensitivity of the individual and group market

The underlying assumption of our previous calculations was that all bilateral elasticities, between individual and individual, individual and group, and group and group contracts, are equal. However, premium elasticities for group contracts are likely to be higher, in absolute size, than individual contracts since a group may have lower search costs than a "group" of single individuals.

Table 4.4	Characteristics of individual and group market					
	2005		2006		2007	
Type of market	Individual	Group	Individual	Group	Individual	Group
Population share	61.6%	38.4%	47.2%	52.8%	44.2%	55.8%
Premiums (euros)						
Unweighted (random witching)	n.a.	n.a.	1069	1007	1152	1077
Weighted (before switching)	n.a.	n.a.	1057	1001	1144	1071
Weighted (after switching)	n.a.	n.a.	1056	998	1144	1070

Some characteristics of the individual and group market are presented in table 4.4. First of all the table shows that the largest expansion of the group market occurred in the first year of the reforms when 14.4% of the population switched from higher priced individual contracts to lower priced group contracts. The 130 million euros total switching gains in 2006 (see table 4.2) are for the largest part determined by people that switched from an individual to a group contract. Indeed, using the numbers from table 4.3 one can show that these switchers gained 107 million euros[20]. The remaining 23 million euros is gained by individuals that switched within the individual market and groups that switched within the group market. If we perform the same calculations for 2007, then we have that 3.0% of the population switched from an individual to a group contract which amounted to gains of about 28 million euros. This comes also close to the reported total switching gains of 31,8 million euros in table 4.2.[21] This also implies that the large switching gains can be mainly attributed to the individuals that switched to lower priced group contracts.

As long as group contracts are lower priced than individual contracts the switch from an individual to a group contract may be a one-off switch, since once enrolled in a group contract one may as well stay in this market.[22] Another consequence is that the price elasticity of the individual market may diminish over the years since the most price sensitive individuals are likely to move to the group market. Given the data at hand it is complicated to measure different elasticities for the group and the individual market. First, it is difficult to define both markets in terms of consumers since in principle all consumers have the possibility to choose a group contract, and also some consumers moved from a group contract in 2006 to an individual contract in 2007. Second, although each consumer is able to choose at least one group contract, the number of group contracts to choose from may be limited. Third, our method assumes that groups compare their premium with an initial offer, but this initial offer was unknown for the group market.

[20] 14.4% of 12,6 million premium payers times the difference of 1057 euro (mean individual premium before switching) and 998 euro (mean group premium after switching) yields about 107 million euros.

[21] 3.0% of 12,6 million premium payers times the difference of 1144 euro (mean individual premium before switching) and 1070 euro (mean group premium after switching) yields about 28 million euros.

[22] A psychological argument may play a role here. Consumers may believe that the only important choice is that you should be enrolled in a (lower priced) group contract.

5 Conclusions

In this study, we develop a new and easy method to compute *short-term* annual price and quality elasticities using annual data on market shares, premiums and quality indicators of *all* health insurers in the market. Our method has three important properties.

A first property is that it can handle stickiness of health insure choice. In the traditional health insurance choice literature often discrete choice models are used to measure premium elasticities from cross-sectional data. One prominent assumption behind these models is that deliberate consumer choice implies that insurers market shares adjust instantaneously to changes in observed insurer characteristics. However, this assumption is often violated since in general stickiness of insurer choice is high. In general, stickiness of consumer choice may be captured by using panel-data models that include fixed effects or lagged market shares as explanatory variables. Similar to Tamm et. al (2007), we find that adding a lagged market share variable is essential to capture these stickiness effects. Since our methods explains changes in market shares it adequately can handle stickiness in insurer choice.

A new feature of our method is that we estimate bilateral price elasticities which relate the net share of switchers between two health insurers not only to their premium difference but also to the market share and premium of the higher priced health insurer. In this method it matters for the price elasticity whether we observe hundred price-switchers from a large to a small insurer than vice versa. In the first case, the potential number of switchers is larger which has a downward effect on the price elasticity.

A general problem with measuring price elasticities is that the researcher does often not observe the reason for non-switching. An insurance market with no switchers at all may have a price elasticity between zero and infinity. Therefore switchers are needed to generate variation in the data and it allows the researcher to test his model. A third property is that our new method can explain the variation in the Dutch market share data well. A model comparison exercise showed that our new method outperformed, in terms of model selection criteria, more traditional estimation methods.

An important lesson of our study is that a health care reform can raise the price elasticity of the total population substantially. Before the health care reform many politicians, and also health economists, were concerned that the low price elasticities in the Dutch social health insurance would persist, also after the health care reform. However, our analysis shows that the Dutch population was very price sensitive after the health care reform and we found a premium elasticity of -7 for the year 2006 and -2 for 2007, while comparable figures, for the years 1996-2005 in the Dutch social health insurance without large reforms, yielded annual price elasticities of only between -1 and 0. Through the reforms, combined with massive media coverage of premium differences, the population seemed to have become much more aware about their possibilities to switch health plans. An important difference with previous years was that many consumers with an individual contract could opt for a lower priced group contract (sometimes within the same insurer). Also, before the reform bad risks were often locked in to an insurance contract, but after the reform were suddenly free to choose their own health insurer. The decline of the price elasticity to −2 in 2007 suggests a return to lower future price-elasticity values than the exceptionally high elasticity found in 2006. The price elasticity in the year 2006 may have been a once-only event.

As from 2006 consumers could also take "quality" information about insurers service levels into account. The estimation results indicate an insignificant "quality" elasticity for 2006 , but a significant service level elasticity of around 1 for 2007. This may suggest that it requires some time before consumers become accustomed to new information.

Our analysis contains various limitations. First, there are data limitations. Insurer characteristics that are important to consumers, such as the provision of supplementary health insurance, are unavailable which leaves the possibility of omitted variables bias. Another limitation of our analysis is that measuring price sensitivity of groups in 2006 and 2007 in principle requires information about all possible premium offers that groups receive from health insurers. This information is unavailable. There is also still a conceptual problem of how to estimate individual and group price elasticities with market share data where most, but not all, consumers can choose between all individual and some group contracts. This problem needs further research.

Data on the quality of health care provision is still lacking in the Dutch health insurance. However, in the last few years the government and also players in the health care market are undertaking more and more efforts to increase transparency with respect to these variables. If, finally, quality information will become available for consumers, than again, this may inspire consumers to reconsider their choice of health insurer and this may again result in a large number of switchers.

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Appendix A: A comparison of different estimators

In this Appendix, we compare our method for estimating a price elasticity with three other type of methods that are used in the literature. The first traditional method that we consider is obtaining price elasticities by regressing first differences in market shares on premium levels. In the second traditional method price elasticities are obtained by regressing growth in market shares on premium levels, and in the third method by regressing levels of market shares on premium levels.

Since we have the availability of panel data, we discuss and compare in the last section also the possibility of computing average (over time) annual price elasticities that follow from (dynamic) fixed effect panel data models.

Regressing first differences of market shares on premium levels

In this section we compare our method with price elasticities that follow from a standard difference model:

$$m_{i,t+1} - m_{i,t} = \alpha_0 + \eta_d P_i \frac{N}{\overline{P}} + \varepsilon_i, \qquad i=1,...,N, \tag{a1}$$

where $m_{i,t+1}, m_{i,t}$ are the market shares of health insurers and P_i the premium. We multiplied P_i with a constant $N/\overline{P}$ so that η_d can be directly interpreted as an average price elasticity. We compare in Table A1 the outcomes with our model specified in equation (2) of the paper:

$$m_{i,t+1} - m_{i,t} = \alpha_0 + \eta S_i^P \frac{1}{N-1} + \varepsilon_i, \qquad i=1,...,N, \tag{a2}$$

In this equation η can be directly interpreted as an average price elasticity.[23] In table A.1 we show that the price elasticities η in model (a2) are significant at a 5% level for the years 2000 until 2007. In those years the R^2 and LogLikelihood are also substantially higher for model (a2) than for model (a1), indicating that model (a2) explains the annual variation in the market share much better than model (a1). Therefore we conclude that size and premium of the highest price health insurer is important and should taken into account when measuring price elasticities.

Although the price elasticities of model (a1) and model (a2) are often in the same range this may not be a general phenomenon. This is shown in the example in Table A.2. In model 1 the largest health insurer (indicated by subscript 1 with 40020 enrolees) sets at the end of year *t* a higher premium (200 euro) than the much smaller health insurer 2 (indicated by subscript 2 with 200 enrolees and 150 euro premium). In model 2 we assume that the smaller health

[23] Note that we included in all equations a constant. Although the estimator for the constant is often not significant different from zero it eases the model comparison exercise. If all models have the same degrees of freedom then it is sufficient to use the R^2 and Loglikelihood as comparison criteria.

Table A.1		Price elasticities, R^2 and loglikelihood for the first difference models											
Elasticities		1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007
Model (a1)	η_d	0.08	– 0.03	– 0.05	–0.13*	–0.13*	–0.04*	–0.10*	–0.18*	–0.16*	–0.20*	–5.1*	–1.4*
Model (a2)	η	– 0.04	– 0.08	– 0.06	–0.16*	–0.25*	–0.14*	–0.06*	–0.16*	–0.20*	–0.22*	–7.0*	–1.9*
R^2													
Model (a1)	η_d	0.01	0.00	0.04	0.19	0.06	0.07	0.05	0.33	0.22	0.45	0.18	0.19
Model (a2)	η	0.00	0.03	0.06	0.42	0.22	0.42	0.18	0.46	0.49	0.66	0.47	0.34
LogLikelihood													
Model (a1)	η_d	156.8	151.8	170.8	172.3	129.7	131.1	119.1	112.7	108.7	112.8	155.1	247.0
Model (a2)	η	156.7	152.2	171.2	177.1	132.1	136.8	120.1	115.0	113.1	117.9	165.6	261.3

A * indicates that the estimated price elasticities are significant at a 5% level.

Insurer 2 sets a higher premium. Note the large differences in price elasticity in the various calculations. Our new estimator takes into account the "market size" effect and generates a much higher elasticity for model 2 than for model 1. The reason is that in both model their are 20 net switchers but in model 1 we have a potential of 40020 consumers that may decide to switch to the lower priced health insurer while in model 2 we have a potential of only 200 consumers that may decide to switch on price. The traditional difference estimator does not take into account the asymmetry in the model and finds a similar price elasticity for both models.

Table A.2	Price elasticities of two different models						
	$x_{1,t}$	$x_{1,t+1}$	$x_{2,t}$	$x_{2,t+1}$	P_1	P_2	η_d	$\hat{\eta}$
Model 1	40020	40000	200	220	200	150	– 0.007	– 0.002
Model 2	40000	40020	220	200	150	200	– 0.007	– 0.44

Regressing growth in market shares on premium levels

In this section, we follow a similar approach but now compare the following two models

$$\frac{m_{i,t+1} - m_{i,t}}{m_{i,t}} = \alpha_g + \eta_g \frac{P_i}{\bar{P}} + \varepsilon_i \,, \qquad \text{(a3)}$$

$$\frac{m_{i,t+1} - m_{i,t}}{m_{i,t}} = \alpha + \eta \frac{S_i^P}{m_{i,t}(N-1)} + \varepsilon_i \,, \qquad \text{(a4)}$$

Again both models care specified such that both η and η_g can be interpreted as annual average price elasticities. Note that model (a4) is similar to model (a2), but that it is rewritten to obtain the same dependent variable as in model (a3), which makes a comparison with model (a3) possible. In table A.3 we present the various statistics that follow from computing the annual

Table A.3		Price elasticities, R^2 and LogLikelihood for growth models											
Elasticities		1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007
Model (a3)	η_g	– 4.9	0.35	– 0.17	–4.3	–2.5	0.35	–0.11	–3.27*	–0.91*	–1.27*	–23.2	–1.8*
Model (a4)	η	–0.05*	–0.05*	–0.08*	–0.09*	–0.73*	–0.04	–0.00	–0.10*	–0.08*	–0.15*	-4.8*	0.15*
R^2													
Model (a3)	η_g	0.00	0.01	0.00	0.01	0.01	0.04	0.10	0.68	0.58	0.70	0.04	0.15
Model (a4)	η	0.29	0.55	0.58	0.99	0.86	0.10	0.10	0.93	0.82	0.98	0.77	0.18
LogLikelihood													
Model (a3)	η_g	– 71.1	– 11.2	– 6.6	– 79.4	– 52.1	18.4	35.4	3.3	24.3	21.2	– 160	17.0
Model (a4)	η	– 66.5	– 0.39	5.3	– 1.1	– 26.3	19.2	35.4	18.9	33.1	51.3	– 121	17.7

A * indicates that the estimated price elasticities are significant at a 5% level.

regressions of the equations in (a3) and (a4). The results are staggering. Again model (a4) outperforms model (a3) in all years and for most years we find huge differences in the R^2 and loglikelihood. Especially in the first years, the traditional growth model does not fit the data very well which is indicated by the very low R^2 . Another interesting fact is that the price elasticities that follow from traditional growth model seem much more volatile, with sometimes even a positive price elasticity, and thus the results do not seem very robust. One of the problem with growth models is that the computation of the elasticity is vulnerable for the presence of outliers. For example small health insurers may sometimes exhibit a large change in the market share which may "blow up" the elasticity.

Regressing the level of market share on premium levels

Finally, we regress market share levels on premium levels:

$$m_{i,t+1} = \alpha_l + \eta_l \, P_i \frac{1}{N\overline{P}} + \varepsilon_i \,, \qquad i=1,...,N \tag{a5}$$

$$m_{i,t+1} = \alpha \; + m_{i,t} + \eta S_i^P \, \frac{1}{N-1} + \varepsilon_i \,, \qquad i=1,...,N \,, \tag{a6}$$

Note that the growth specification of market shares is often the specification that is used by most researchers, since this specification follows from a conditional logit specification (see e.g. Dowd and Feldman, 2006).[24] Again both models are specified such that both η and η_l can be interpreted as annual average price elasticities. Note that model (a6) is similar to model (a2), thus all price elasticities and loglikelihoods will be the same as reported for model (a2) while the R^2 increases. Again the results are staggering. First of all, a level specification of market shares does not really seem to work. All reported R^2 for model (a5) are very small. This result

[24] Dowd and Feldman (2003) compare the price with a price of a reference plan. In our case consumers are not restricted in their choice and can choose, in principle, for each health plan in the market.

Table A.4		**Price elasticities, R^2 and LogLikelihood for level models**											
Elasticities		1996	1997	1998	1999	2000	2001	2002	2003	2004	2005	2006	2007
Model (a5)	η_d	– 3.1	– 0.90	– 0.16	– 0.40	– 1.54	– 1.7	– 0.06	3.5	0.1	1.4	– 5.1	– 6.0
Model (a6)	η	– 0.04	– 0.08	– 0.06	–0.16*	–0.25*	–0.14*	–0.06*	–0.16*	–0.20*	–0.22*	–7.0*	– 1.9*
R^2													
Model (a5)	η_d	0.01	0.01	0.00	0.00	0.02	0.10	0.00	0.12	0.01	0.03	0.06	0.04
Model (a6)	η	0.99	0.99	0.99	0.99	0.99	0.99	0.99	0.99	0.99	0.99	0.99	0.99
LogLikelihood													
Model (a5)	η_d	55.8	59.7	57.6	57.6	50.3	47.5	35.5	37.1	36.0	36.4	119.1	116.6
Model (a6)	η	156.7	152.2	171.2	177.1	132.1	136.8	120.1	115.0	113.1	117.9	165.6	261.3

A * indicates that the estimated price elasticities are significant at a 5% level.

suggest that most non-switchers stayed at their health insurer because of “status quo bias” or transaction costs. Furthermore, the sign of the price elasticity η_l is negative until the year 2002, positive for the years 2003 until 2005, and again negative for the years 2006 and 2007. This alternating sign represent the fact that in the beginning of the nineties all consumers were assigned to a regional health insurer on the basis of their place of residence. The year 1996 was the first year that the Dutch government allowed health insurer to differ their nominal premiums. The premium differences in 1996 were, however, extremely small and also the overwhelming part of the population did not know that there was a possibility to switch from health insurer. Therefore, the negative price elasticities η_l during 1996, and also during 1997-2002 are more likely to reflect the “coincidental” correlation between the historical market shares of health insurers and premiums, and may therefore not reflect the price sensitivity of consumers. Interesting is also that the price elasticity of –6,0 in the year 2007 in model (a5) remains high. This represents the fact that more than 50% of the population holds a lower priced group contract.

Fixed effect models

Many characteristics or variables of health insurers and consumers are unobserved. If these omitted variables are correlated with the observed price or quality variables then the price and quality elasticity may be biased. Adding an insurer specific fixed effect may capture (some of) these unobserved characteristics. A fixed effect model, however, can only be specified if the researcher has panel data. In the next section we will estimate two fixed effect models for the Dutch social health insurance for the period 195-2005. For the new health insurance system we consider the time span of two years, 2006-2007, too short to obtain reliable estimates from a a fixed effect model.

We now first show that adding lagged market shares to a traditional type of fixed effects models improves the results considerably. Consider the following fixed effect model:

$$m_{i,t} = \alpha_{0i} + \alpha_{0t} + \alpha_1 m_{i,t-1} + \eta_{d,t} P_{i,t} + \varepsilon_{i,t}, \qquad i=1,..., N_t, \quad t=1996,...,2005 \qquad (a7)$$

where $m_{i,t}, m_{i,t-1}$ are the market shares of health insurers, $P_{i,t}$ the premium and N_t the number of insurers in period *t*. Note that we also included fixed time effects to capture specific annual measures such as the annual changes in the nominal premium level, changes in the number of insurers, changes in basic benefit package etc. Estimating this fixed effect model yields an $\hat{\alpha}_1 = 0{,}988$ with a standard error of 0,023 indicating that the hypothesis $\hat{\alpha}_1 \neq 1$ cannot be rejected. This result shows that a lagged market share variable should be included in these type of specifications and that its coefficient is not only positive, but also very close to one.[25] This result has two important implications. Estimating a fixed effect panel data without adding a lagged market share data, as is done in many studies, may not be sufficient since fixed insurer effects cannot explain the variation that is explained by the lagged market share variable(s). Second, a model which explains changes in market shares by the level of premiums may be sufficient. This result is in line with a recent paper of Tamm et al. (2007) who find a similar result for German market share data. We also compared our new method with a traditional fixed effect model. Consider the following two fixed effects models:

$$m_{i,t} - m_{i,t-1} = \alpha_{0i} + \alpha_{0t} + \eta_{d,t} P_{i,t} \frac{N_t}{\overline{P}_t} + \varepsilon_{i,t}, \qquad i=1,...,N, \ t=1996,...,2005 \qquad (a8)$$

$$m_{i,t} - m_{i,t-1} = \alpha_{0i} + \alpha_{0t} + \eta_t \frac{S_{i,t}^P}{N_t - 1} + \varepsilon_{i,t}, \qquad i=1,...,N, \ t=1996,...,2005 \qquad (a9)$$

The specifications are similar to model (a1) and (a2) but now we included fixed period and insurer effects. The estimation results of both models are shown in table A.5[26]

Table A.5 A comparison of two fixed effects models

Model (a8)		1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
	$\hat{\eta}_{d,t}$	0.02	– 0.02	– 0.03	– 0.02	– 0.07	– 0.03	– 0.03	– 0.12*	– 0.12*	– 0.15*
	s.e.	(0.17)	(0.06)	(0.05)	(0.06)	(0.05)	(0.03)	(0.03)	(0.04)	(0.04)	(0.03)
Degrees of freedom:		190	R^2:	0.724	LogLikelihood:		1478.7				
Model (a9)		1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
	$\hat{\eta}_t$	– 0.04	– 0.07	– 0.05	– 0.06	– 0.16*	– 0.09*	– 0.00	– 0.10*	– 0.14*	– 0.16*
	s.e.	(0.17)	(0.05)	(0.05)	(0.05)	(0.05)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)
Degrees of freedom		190	R^2:	0.749	LogLikelihood:		1490.6				

A * indicates that the estimated price elasticities are significant at a 5% level.

[25] We performed many sensitivity tests, but this result proved to be robust.

[26] Both models are estimated by OLS. See Tamm et. al. (2007) for a more thorough discussion on estimation issues.

Before discussing the results in table A.5 we should note that without fixed insurer effects model (a8) yields an R^2=0,18 and model (a9) an R^2=0,33. This indicates that the fixed insurer effect can explain a lot of the variation in the data. Table A.5 shows that both models are close but model (a9) still performs slightly better than model (a8) in terms of goodness-of-fit criteria.