CPB Discussion Paper

No 56

February 2006

Consumer price sensitivity in health insurance

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ISBN 90-5833-257-8

Abstract in English

This CPB Discussion Paper presents new estimates for the price elasticity of the residual demand for health insurance. This elasticity measures the loss in market share of a health insurer as a consequence of a unilateral increase in price, assuming other firms keep their prices constant. The main findings are as follows: the price elasticity of residual demand for social health insurance by enrollees was very low during the period 1996-2002. We find small but significant effects of the price of basic insurance but no robust effect of the price of supplementary insurance. Young enrollees are more price sensitive than older enrollees. However, these findings are conditional on the limited variation in price observed in our data. At larger price differentials, the elasticity may well be higher. This Discussion Paper is based on joint work of Machiel van Dijk, Marc Pomp, Rudy Douven (all three at CPB), Trea Laske-Aldershof, Erik Schut (both Erasmus University), Willem de Boer and Anne de Boo (Vektis). We would like to thank Marieke Smit (Vektis) for her help in getting this project off the ground. We would also like to thank Katie Carman (Tilburg University) for her comments on a previous draft of this paper.

Key words: Health insurance, price elasticity

JEL code: D12, I11, I18, L11

Abstract in Dutch

Dit CPB Discussion Paper presenteert nieuwe schattingen van de prijselasticiteit van de residuele vraag naar zorgverzekeringen. Deze elasticiteit meet het verlies in marktaandeel van een verzekeraar na een prijsverhoging onder de aanname dat andere zorgverzekeraars hun prijs niet veranderen. De belangrijkste uitkomsten luiden als volgt: voor werknemers was de prijselasticiteit van de residuele vraag naar zorgverzekeringen erg laag in de periode 1996-2002. Jongere werknemers zijn iets prijsgevoeliger dan oudere werknemers. Een kanttekening bij deze resultaten is, dat de achterliggende gegevens vrij weinig variatie in prijzen tussen verschillende verzekeringsmaatschappijen laten zien. Bij grotere prijsverschillen zou de prijsgevoeligheid belangrijk kunnen toenemen. Dit Discussion Paper is het resultaat van gezamenlijk onderzoek van Machiel van Dijk, Marc Pomp, Rudy Douven (CPB), Trea Laske-Aldershof, Erik Schut (Erasmus Universiteit), Willem de Boer en Anne de Boo (Vektis). We danken Marieke Smit (Vektis) voor haar hulp bij de start van dit project. We danken Katie Carman (UvT) voor haar commentaar op een eerdere versie van dit paper.

Steekwoorden: Zorgverzekeringen, prijselasticiteit

Contents

1 Introduction

Competition between health insurance firms is a central pillar of the market-based reforms which are currently being introduced in the Dutch health care sector. Hence it is important to have a good idea of the current state of competition in the market for health insurance. Earlier research (summarised below) has indicated that competition among health insurance firms is rather weak. If this continues to be the case in more recent years, then additional measures to stimulate competition in this market may be called for.

In this paper we present new empirical estimates of an indicator of competition, the elasticity of residual demand for health insurance. This elasticity measures the loss in market share of a health insurer as a consequence of a unilateral increase in price, assuming other firms keep their prices constant. The elasticity of residual demand is an important determinant of the level of competition. If this elasticity is small, then insurers will be able to set prices substantially higher than marginal costs.

Our estimates are based on a dataset covering all Dutch citizens who obtained health insurance through one of the 20 sickness funds in 2002. We use this dataset to construct bilateral flows of insured during 1993-2002 between each pair of health insurance firms in our dataset. Regressing these bilateral flows on price differences between each pair of health insurance firms allows us to estimate price elasticities of residual demand. The advantage of using this dataset over previous work that relied on time series of market shares per firm, is that we are able to estimate differences in price elasticity for many different subgroups of the population. Another advantage is that looking at bilateral flows (rather than aggregate market shares) gives us sufficient degrees of freedom to correct (at least to some extent) for unobserved firm-specific effects.

The paper is structured as follows. The next section provides some background on the Dutch system of health insurance. Section 3 summarises existing estimates of the price elasticity of health insurance in the Netherlands and some other countries. Section 4 describes the data and the estimation method. Section 5 presents estimation results. Section 6 uses these results to calculate elasticities of residual demand. Section 7 concludes.

2 Health insurance in the Netherlands

2.1 The institutional setting before the reforms

Until 2006, the Dutch market for health insurance was split into two segments, distinguished mainly by income of the insured. The first segment, covering about 60 percent of the population, consisted of compulsory insurance for workers and their dependents with incomes below a certain threshold (2005: euro 32 600). This segment is served exclusively by so-called sickness funds: not-for-profit health insurers. Private insurers were not allowed to operate in this part of the market; these firms had to focus their activities exclusively on the part of the population with incomes above the income threshold of euro 32 600. Also health providers were heavily regulated, with respect to prices as well as with respect to entry.

The regulatory regimes differs in important respects between the two segments of the health insurance market. Sickness funds face a number of restrictions that do not apply to private for-profit insurers. In particular, sickness funds must:

- Offer a basic insurance policy, the coverage of which is determined by the government;
- Contract with every hospital;
- Pay administratively set prices to hospitals and general practitioners;
- Take part in the risk-equalisation scheme run by the government;
- Accept every citizen at the same nominal premium (community rating) for basic insurance, irrespective of expected health costs, during annual open enrolment periods.

As far as basic insurance is concerned, sickness funds are financed partly through income-related premiums set by the government, and partly by so-called nominal premiums set by the sickness funds themselves and paid by the insured. Currently the nominal premium accounts for about 15 - 20% of the total premium. By varying the nominal premium, sickness funds are able to compete on price in the market for basic insurance.

Sickness funds also sell supplementary insurance but via a separate legal entity, covering *inter alia* dental care, physical therapy and alternative medicine. None of the above restrictions apply to supplementary insurance. Sickness funds have to cover the costs of supplementary insurance entirely out of their premium income.

2.2 Basic insurance and supplementary insurance: tied sales?

In the Netherlands, the overwhelming majority of sickness funds enrolees (over 90%) also buy voluntary supplementary insurance for health costs not covered by basic insurance. As already pointed out, open enrolment does not apply to supplementary insurance. Moreover, almost all firms restrict supplementary insurance to those insured who purchase their basic policy from the same firm (Schut et al., 2004). This suggests that most enrolees will choose a health insurer on the basis of the price for the total package (basic insurance + supplementary insurance). However, because of the lack of transparency of the market for supplementary insurance due to the proliferation of different policies, enrolees may attach a greater weight to the price of basic insurance in their choice of sickness fund. For this reason, we will not *a priori* assume that consumers base their choice of sickness fund on the price of the total package (basic + supplementary insurance).

3 Literature review

In summarising the empirical literature on residual elasticities of demand, it is important to stress that most studies look at the elasticity of demand with respect to the *out of pocket premium* (the part of the premium directly paid by the insured). The out-of-pocket premium covers only a part of the insurance bill: the government and/or employers usually pay a substantial part of the premium. In the US, the out-of-pocket premium usually covers only 10 – 20 percent of the total premium, in Germany the share is about 50% while in the Netherlands, on average out of pocket premiums amount to 10-15% of the total medical expenses (Schut and Hassink, 2002). Because in the Netherlands, employers and employees both pay an income-related amount irrespective of the sickness fund chosen, consumers pay the full out-of-pocket price differential between health insurers. This is different from the situation in Germany where consumers pay only a percentage of the price differential (Schut et al., 2003). In the US, there is a shift among employers from a percentage contribution to a fixed subsidy based on the cheapest health plan on the menu. As a consequence, employees increasingly pay the full price differential among different plans.

Table 3.1 summarises recent estimates from the literature of the out-of-pocket elasticity of residual demand. Clearly, estimates differ widely, not only between countries but also within countries. Elasticity estimates for the Netherlands are low compared to Germany and the US. Indeed, Schut and Hassink (2002) show that their estimation results imply that raising the premium for basic insurance is a profitable strategy for an average sickness fund, at least in the short run (Schut and Hassink, 2002, p. 1023).

We should stress that a comparison of estimated price elasticities of health plan choice in different countries is not straightforward, because of different base levels of out-of-pocket premiums and market shares. For instance, in Germany out-of-pocket premiums are at least twice as high as the out-of premiums that typically paid by US employees with employment-based group insurance. Due to the higher level of out-of-pocket premiums in Germany the estimated price elasticities of plan choice in Germany are likely to be two to five times as large as in the US. Also, the German estimates refer to the choice of *type* of sickness fund, not to the choice of any specific sickness fund.

Table 3.1 Out-of-pocket elasticities of demand for health insurance: literature survey

Author	Period	Elasticity
A. The Netherlands		
Schut and Hassink (2002), basic insurance	1996-1998	– 0.3
Schut and Hassink (2002), supplementary insurance	1996-1998	– 0.8
Schut and Hassink (2002), basic + supplementary insurance	1996-1998	– 0.4
Schut et al. (2003)[a]	1996-2000	0.0 - – 0.4
B. Germany		
Schut et al. (2003)[b]	1996-2000	0.4 - – 5.3
C. US		
Strombom et al. (2002)[c]	1995	– 0.2 - – 1.7
Royalty and Solomon (1999)[d]	1994-1995	– 0.1 - – 1.5

[a] Most estimates are insignificant, the only exception being pensioners where a significant coefficient is found for supplementary insurance (the elasticity is– -0.36).

[b] Posiitive elasticity applies to pensioners; not significant

[c] Strombom et al. (2002) estimate total premium elasticities from which we have calculated out-of-pocket premium elasticities; the highest elasticities apply to young and to newly hired workers.

[d] Higher elasticities for younger, healthier workers; Royalty and Solomon (1999) also present much higher estimates based on a model including fixed effects. However, these are unrepresentative for the whole population since fixed effects logits can only be estimated on the part of the population that has actually switched.

4 Data and method

4.1 The dependent variable

4.1.1 Starting point: individual cross-section data

For the purposes of this research we have obtained access to the complete records of all 10 million Dutch citizens who were covered by one of the 20 sickness funds in 2002.[1] This dataset is maintained by Vektis, a private firm that is fully owned by the Dutch federation of health insurers (ZN). For estimation purposes we have exclude children who do not choose their own sickness fund, but are enrolled via their parents. This reduces the dataset to 8.0 million observations.

4.1.2 Transforming cross-section data into paneldata

Estimating a price elasticity from a single cross-section is problematic, the more so since we do not have much variation in our premium variable: every insured is faced with the same set of prices. As a result each insured faces the same price vectors of 20 prices each, one for basic insurance and several for supplementary insurance. This makes it impossible to correct for unobserved firm specific effects which could be correlated with prices, leading to a bias in the estimated coefficient for price.

However, we are able to transform a large part of our cross-section data into paneldata by exploiting the fact that until 1993, each sickness fund had a designated geographical area in which it was the sole provider of social health insurance. We make the following three assumptions (below we will return to the realism of these assumptions):

- All individuals in our dataset have been insured by a sickness funds uninterruptedly during the whole period 1993-2002 (not necessarily the same sickness funds).
- No individual in our dataset has moved between regions during the period 1993-2002.
- Individuals have switched at most once during the period 1993-2002.

[1] A small sickness fund, OZB, which worked exclusively for a large Dutch multinational company, is not included in the analysis. Furthermore, some firms merged during the period 1993-2002. We assume that the insured of the merged firms pay the premium of the largest of the merged firms. Using a weighted premium of the merged firms would have been better, but for supplementary insurance this required identifying a comparable supplementary policy for the merged firms. We do not have sufficient data to do this. For the firms that we observe in 2002 we take heterogeneity of supplementary insurance into account by including firm specific coefficients. However, this solution cannot be used to correct for within-firm heterogeneity. For basic insurance we checked whether including weighted premiums would alter our outcomes, and this turned out not to be the case. Weighted premiums where almost identical to the premiums of the largest of the merged firms, possibly because firms already coordinated their pricing behaviour prior to the formal merger. We also did a sensitivity analysis by excluding the Amsterdam region were a merged firm is the dominant insurer. This did not alter the estimation results either.

Given these assumptions, it follows that in 1993 all individuals who are in our dataset in 2002, were insured by the monopolist sickness fund in their region of residence. Therefore, individuals who were still insured by the former regional monopolist in 2002 can be classified as non-switchers. On the other hand, individuals who are not insured by the former regional monopolist must have switched from the former monopolist to their current insurer.

Dropping observations for which these assumptions are unlikely to hold: Unfortunately, we do not know whether someone who is insured by a sickness fund in 2002 had been insured by a sickness fund during the whole period. Indeed, for some groups in the population, it is likely that a substantial share switched from private insurance to social health insurance. For example, the self-employed were not covered by social health insurance until 2000. Starting in 2000, this group also became legally obliged to buy insurance from one of the sickness funds (Schut et al., 2003). Therefore, we omit the self-employed - about 300 thousand observations - from our empirical analysis.

Something similar applies to the elderly (Schut and Hassink, 2002). In 1997, about 90,000 privately insured elderly suddenly became entitled to social health insurance because of a substantial increase in the income threshold below which they were eligible for social health insurance. Therefore, we also omit the elderly (aged 65 and over in 2002) - about 1,5 million observations - from our empirical analysis. Other research (e.g. Buchmueller, 2000) indicates that price elasticities for the elderly are much lower than for workers. Note that since we will estimate age-specific elasticise, omitting one age group will not bias the results.

Among the non-working of working age, many individuals have been working during some years in the period 1993-2002, but lost their jobs, resulting in a fall in income. If this fall in income led to an income below the threshold for sickness funds, then these individuals became eligible for social health insurance. If these individuals chose another sickness funds than the former regional monopolist, then our procedure would erroneously classify this group as switchers. Thus, we leave out this groups as well.

Finally, most individuals below age 25 in 2002 were insured through their parents in 1993. If their parents were privately insured in 1993, and if these individuals chose another sickness funds than the former regional monopolist, then this group is erroneously classified as switchers. Therefore we also leave out this group. This leaves us with all observations on workers aged 25-64.

Since it is unlikely that workers have experienced a fall in income, it is likely that these workers were insured by a sickness fund in 1993.[2]

Moving between regions: If an individual moves from one region to another but stays with the same sickness fund, then our procedure results in erroneously classifying this individual as a switcher. Kalshoven (1999) reports annual movements between regions of 0.2% to 0.3% of all households enrolled in a sickness fund. Cumulated over 10 years this amounts to 2 to 3% of all households. This classification error is fairly small given that the percentage of switchers in our dataset is equal to roughly 20% of all households (see below).

Identifying regions: In order to determine which firm was the regional monopolist for a given individual, we used municipality codes. There are 496 municipalities in our dataset. Within each municipality we identify the former monopolist on the basis of market share. The largest firm is defined as the former monopolist. This always correctly identified which firm was the former regional monopolist *in most of the municipalities*. However, the designated regions did not always coincide exactly with municipality borders: in quite a few cases one geographical area within a municipality was served by one firm while some other geographical area within the same municipality was served by another firm. In these cases, we will incorrectly classify individuals who had been insured by the smaller of these sickness fund as switchers or as non-switchers. Therefore, we used a cut-off point of 60%: if in a given municipality the largest firm had a market share in 2002 of 60% or more, then we included the municipality in the analysis, otherwise we omitted this municipality. In this way we exclude most municipalities where in 1993 there was a second large sickness funds active.[3] Applying this rule leads us to exclude 65 municipalities and a drop in the number of cases of about one million. In section 7 we discuss sensitivity analysis using another cut-off point.

Resulting dataset: If we follow the procedure just described, we obtain the dataset presented in Table 4.1. The table shows the total number of individuals in our dataset along with the number of individuals who have ‘switched’ from one sickness fund to another sickness fund in the period 1993-2002, broken down by age (in 2002) and gender. Clearly the propensity to switch depends on age (the young switch more) and gender (men are more likely to switch). It

[2] There is one other group where similar problems might occur, but which we nevertheless include in our empirical analysis since we believe that for these groups the problem is fairly small. This group consists of (mostly) women who re-entered the labour force. Some of these women may have been insured via their husband’s private insurance policy before re-entering the labour force. After re-entering they may have become eligible for social health insurance. Of all working women between 1% and 2% are women who re-entered the labour force during the previous two years (calculated from data provided by Statistics Netherlands). This means that the flow of women the re-enters the labour force in each year is between 0,5 and 1% of the female labour force. Only a minority of these women had husbands who were privately insured. Assuming this share to be equal to the share in the population as a whole yields one-third. Thus, the annual number of women erroneously classified as switcher amounts to only 0.2- 0.3%. On the other hand, re-entrants are primarily found among the 25-45 year old. Given that the percentage of female switchers in these age groups in our data is about 20% of all households (see below), this classification error is fairly small.

[3] De Bekker and van den Brink (2002) report that as of 1999 the former regional monopolist still had a market share of 80% or more. Extrapolating to 2002 would lead to a market share of 70-75%. Using this ratio of 2002 markets shares to 1993 markets shares (i.e. ¾), a 2002 market share of 60% would correspond to a 1993 market share of 80%. This would imply that, if in1993 another firm was active in this municipality, its 1993 market cannot have exceeded 20% (100%-20%).

also emerges quite clearly that the former regional monopolist still has a very large share of this market: 81% (100%-19.2%) on average.

Table 4.1 Descriptive statistics: number of switchers 1993-2002, by age and gender (workers)

	Age				
	25-34	35-44	45-54	55-64	Row total
Women					
Number of obs	605863	458189	297572	169258	1530882
Number of switchers	140638	74711	40014	20400	275763
Percentage of switchers	23.2	16.3	13.4	12.1	16.3
Men					
Number of obs	585885	511814	367300	140393	1605392
Number of switchers	160247	114771	77702	24418	377138
Percentage of switchers	27.4	22.4	21.2	17.4	22.1
Total					
Number of obs	1191748	970003	664872	309651	3136274
Number of switchers	300885	189482	117716	44818	652901
Percentage of switchers	25.3	19.4	17.3	14.7	19.2

4.1.3 From individual data to bilateral flows

The paneldataset just created allows us to construct bilateral flows from former regional monopolists to the sickness funds that have entered the region since 1993. Of the 20 firms in our dataset in 2002, 15 had been a regional monopolist until 1993. Of the other five firms, four entered the market since 1993 while one firm was active in 1993 but not as a regional monopolist: it shared ‘ its’ region with another firm, so in that region there was a duopoly. We omitted this region from our dataset since in case of a duopoly our procedure for constructing paneldata breaks down.

With these numbers of firms, we can construct 15 x 19 = 285 bilateral flows: from each of 15 former regional monopolist to each of 19 other firms. Since we will calculate separate flows for men and women and for each of four age groups, the total dataset contains 2280 observations. We will use these flows from former regional monopolists to new entrants into the regional market as our dependent variable. Note that this procedure implies that new entrants can only gain customers. In order to adjust for differences in the size of the regional markets, we will divide each flow by the size of the total regional market in 2002.

Table 4.2 presents summary statistics for the bilateral flows that result from these calculations. Each entry in the table shows the average number of individuals who switched from a former regional monopolist to another firm, as a percentage of the total size of the regional market for the relevant age/gender group in 2002.[4] To arrive at the average total outflow from a regional

[4] Since we are restricting our dataset to insured who were (probably) continuously insured during 1993-2002, the total market in 2002 was the same size as the total market in 1993.

monopolist, each entry in the table must be multiplied by 19 (since each former monopolist faces potential competition in its former designated region from 19 other firms).[5]

Table 4.2 **Switching 1993-2002: bilateral flows by age and gender, % of total market**

	Age			
	25-34	35-44	45-54	55-64
Women				
Mean	1.3	0.9	0.8	0.7
SD	2.1	1.7	1.5	1.4
Men				
Mean	1.5	1.3	1.2	1.0
SD	2.4	2.1	2.1	1.8

Source: see text.

4.2 Explanatory variables

4.2.1 Prices

Our main interest is in the effect of *differences* in price between any two sickness funds on switching between these two firms. In order to estimate this effect, we calculate for each pair of firms the difference in premium between these firms. This is illustrated in Table 4.3. Thus, our hypothesis is that a larger positive price differential between the incumbent (the former monopolist) and a regional entrant (all other firms) will result in a larger flow from the incumbent to the entrant.

Until 1996 there was no substantial premium variation among sickness funds. The reason for this was that sickness funds were hardly at risk for the medical expenses of their enrollees. In 1996 the financial risk for sickness funds was raised from 3 to 15 percent of the difference between the expected and realised medical expenses of their enrollees (sickness funds received risk-adjusted capitation payments for compensating most of the expected medical expenses). Supplementary insurance played hardly a role until 1996. Starting in 1996 this changed when dental care and physiotherapy where transferred from basic insurance to supplementary insurance.

[5] Multiplying the figures in table 4.2. by 19 does not exactly reproduce the corresponding figures in table 4.1, since outflow rates differ between regions and firms.

Table 4.3 **Construction of premium variables**

From firm:	1	2	…	15
To firm:				
1	-	p1-p2	p1-p..	p1-p15
2	p2-p1	-	p2-p..	p2-p15
..	p..-p1	p..-p2	-	p..-p15
20	p20-p1	p20-p2	p20-p..	-

Since our dependent variable consists not of annual flows but flows of insured over a ten year period (1993 – 2002), we use the *average price difference* between each pair of firms during these years as an explanatory variable. We use average prices for the years 1996-2002 since (as argued above) prior to 1996 switching cannot have taken place in response to price. Moreover, prior to 1997 open enrolment applied two two-year periods and the open enrolment period was not synchronised across sickness funds. This suggests that between 1993 and 1996, the numbers of switchers will have been lower than in later years.

Using average prices may lead to a bias in estimated coefficients, since differences between average prices will tend to be smaller than differences in annual prices. This suggests that the bias is in the direction of overestimating the effect of prices: we are attributing observed switching behavior to smaller price differences than the annual price differences confronting the insured.

For basic insurance, prices of different insurers refer to exactly the same product so we are dealing with a homogenous product. The coverage of the basic package is set by the government and service levels play at most a limited role, since health providers send their bills directly to sickness funds. Customers face no risk of slow reimbursement by sickness funds. Schut and Hassink (2002) also argue that sickness funds offer a standardised product: “Since sickness funds offer standardised benefits and are only starting to employ managed care activities, it is unlikely that price variation can be explained by differences in quality or efficiency in purchasing or organising medical care.” Schut and Hassink (2002) p. 1017).

The same does not apply to supplementary insurance, where comparing policies is hampered by differences in coverage across policies and firms. We have tried to solve this by selecting a supplementary package that is more or less identical across insurers. However, we cannot exclude the possibility that price differences reflect to some extent real differences in quality (i.e. coverage) across supplementary policies. It is also possible that consumers interpret differences in these prices as indicators of differences in quality, even if in reality this is not the case. In that case we could in theory find a positive elasticity of residual demand.

Table 4.4 shows average prices for basic insurance, supplementary insurance and the total package. Note that price differentials are fairly small but not negligible. The maximum savings for the total package amount to about 100 euro per year. Since premium differentials are fairly stable over time (see below), this implies that the present value of switching to a cheaper fund could be substantial for some consumers.

Table 4.4 Prices: average annual premium 1996-2002 (euro)			
Sickness fund	Basic insurance	Supplementary insurance	Total
1	148.7	50.8	199.5
2	145.6	56.8	202.4
3	158.5	50.0	208.5
4	127.3	83.8	211.1
5	139.9	74.5	214.4
6	153.1	66.5	219.6
7	149.0	71.7	220.8
8	157.9	64.9	222.7
9	139.2	84.4	223.6
10	143.2	84.0	227.2
11	164.3	63.7	228.0
12	157.5	73.6	231.1
13	164.6	70.2	234.8
14	137.0	103.5	240.5
15	159.3	88.5	247.8
16	141.4	106.4	247.8
17	140.8	122.1	262.9
18	176.9	87.6	264.5
19	178.3	93.7	272.0
20	156.6	133.8	290.4

Source: iBMG database on health insurance premiums (unplublished)

It is important to note that differences in prices between firms are fairly stable over time, in the following sense: a firm that is relatively expensive in year t will also be relatively expensive in year t+1. This applies to both basic and supplementary insurance; especially in the latter case correlations over time are very high. We also note that the spread in supplementary premiums is much larger than the spread in premiums for basic insurance as measured by the coefficient of variation (100 x standard deviation / mean), possibly indicating heterogeneity in the coverage of supplementary insurance (see Table 4.5). This suggests that it is important to include firm specific coefficients in the model to be estimated.

Table 4.5 Prices: correlation over time and coefficient of variation				
	Correlation (t, t+1)	Basic insurance coefficient of variation	Correlation (t, t+1)	Supplementary insurance coefficient of variation
1996	0.58	1.90	NA	47.93
1997	0.98	9.54	0.96	43.20
1998	0.79	10.36	0.95	58.11
1999	0.91	7.45	0.98	55.90
2000	0.86	9.72	0.97	45.25
2001	0.82	15.69	0.95	50.17
2002		10.93		26.20

Source: calculated from time series on prices.

4.2.2 Age and gender

We included age and gender as explanatory variables. In addition, these variables were interacted with the price variables in order to assess whether price elasticities depend on these characteristics.

4.2.3 Firm specific coefficients

Apart from price, we include three sets of firm specific coefficients that should pick up unmeasured firm-specific effects. First, we include coefficients for each former monopolist (α_i). The coefficient for firm *i* is included if firm *i* is involved as a former monopolist. Second, we include a set of firm specific coefficients for firms that are not the former monopolist in a region (α_j). This coefficient is included if firm *j* is involved in a bilateral flow. Third, we include a coefficient for new sickness funds (α_{new}).

The coefficients α_i and α_j cannot be included simultaneously. This is because including both α_i and α_j would result in perfect collinearity between the firm specific effects and price.

4.3 Equations to be estimated

We will estimate six different models: three models with total premium as the price variable and three models with separate variables for the price of basic and supplementary insurance. For each of these choices of price variables, we present results for a model without any of the firm specific coefficients defined in section 4, with firm specific coefficients for the former regional monopolist (α_i) and with firm specific coefficients for entrants (α_j). In the first case, we also include a separate dummy for new firms (firms that were not active in any of the regional markets in 1993). In the second case, any effects of this dummy are picked up by the firm specific coefficients for entrants.

The estimated equations with total premium as the price variable read as follows:

$$F_{i\to j,\, j\neq i} = \alpha_1 + \alpha_{age} + \alpha_{male} + \alpha_{age}(\overline{P_j} - \overline{P_i}) + \alpha_{male}(\overline{P_j} - \overline{P_i}) \quad (1)$$

$$F_{i\to j,\, j\neq i} = \alpha_i + \alpha_{age} + \alpha_{male} + \alpha_{age}(\overline{P_j} - \overline{P_i}) + \alpha_{male}(\overline{P_j} - \overline{P_i}) + \alpha_{new} \quad (2)$$

$$F_{i\to j,\, j\neq i} = \alpha_j + \alpha_{age} + \alpha_{male} + \alpha_{age}(\overline{P_j} - \overline{P_i}) + \alpha_{male}(\overline{P_j} - \overline{P_i}) \quad (3)$$

Where:

$F_{i\to j, i\neq j}$ = The number of switchers from firm *i* to form *j* during 1993-2002, as a percentage of the total number of insured in the region where *i* was the former regional monopolist ((i=1,..,15, j=1,..20).

α_i, α_j	=	Coefficients for former regional monopolists (subscript *i*) and regional entrants (subscript *j*).
α_{new}	=	Coefficient which is included if $F_{i \to j, i \neq j}$ concerns a flow to a new firm.
$(\overline{P_j} - \overline{P_i})$	=	Total average premium entrant minus total average premium former regional monopolist, over the period 1996-2002.
α_{age}	=	Set of four coefficients, one for each age class (25-34, 35-44, 45-54, 55-64). The relevant coefficient is included if $F_{i \to j, i \neq j}$ consists of a flow of switchers in the relevant age class.
α_{male}	=	Coefficient which is included if $F_{i \to j, i \neq j}$ concerns a flow of male switchers.

The estimated equations with separate variables for the price of basic and supplementary Insurance are similar, except that we replace $(\overline{P_j} - \overline{P_i})$, by two other variables, $(\overline{P_j^B} - \overline{P_i^B})$ and $(\overline{P_j^S} - \overline{P_i^S})$ where *B* stands for basic insurance and *S* for supplementary insurance.

Note that we are including interaction terms with price for *each* age group. As a consequence, we do not (indeed, we cannot) include price separately (i.e., not interacted). In effect, we allow the coefficient(s) on price to vary freely across age groups.

Note also that we also impose the effect of gender to raise or lower the whole age profile of the price-coefficient by the same absolute amount, irrespective of age. Allowing the age profile of the price coefficient to differ between men and women in an unconstrained way leads to very similar results.

5 Results

5.1 Estimated coefficients

We report results based on both OLS and WLS (weighted least squares) since diagnostic tests indicate heteroskedasticity in most cases. The main effect of using WLS instead of OLS is a substantial increase in t-values.

Tables 5.1 and 5.2 show estimation results for equations with total premium as the price variable, while Tables 5.3 and 5.4 show estimation results for equations with the basic and supplementary insurance as separate explanatory variables. In order to avoid cluttering of tables we do not report estimated coefficients on firm specific coefficients or intercepts. We also omit the coefficient for new entrants, which was always significantly negative.

In all equations we find that the probability of switching falls with age and is larger for men than for women: this is in line with the descriptive statistics reported in Table 4.1.

Turning to the estimated coefficients for price, we find that including firm specific coefficients matters a great deal, in particular for supplementary insurance. Quite strikingly, including coefficients for regional entrants leads to insignificant results for total and supplementary insurance but not for basic insurance. This points to omitted firm characteristics and/or heterogeneity in the quality (probably coverage) of supplementary insurance. On the other hand, one fairly robust finding is the significantly negative coefficient on the price of basic insurance.

In general, we find plausible age patterns: the estimated coefficient is larger in absolute term for younger enrollees. In equations where we distinguish between basic and supplementary insurance, we find that men are more sensitive to price than women.

Table 5.1	**Estimation results, total premium, OLS**					
	No firm specific coefficients		With firm specific Coefficients α_i		With firm specific coefficients α_j	
Variable	Estimate	T-value	Estimate	T-value	Estimate	T-value
Age 35-44	– 0.317	– 2.696	– 0.315	– 2.802	– 0.316	– 3.282
Age 45-55	– 0.452	– 3.828	– 0.441	– 3.909	– 0.458	– 4.736
Age 55-64	– 0.568	– 4.758	– 0.552	– 4.832	– 0.612	– 6.255
Male	0.329	3.920	0.320	3.982	0.323	4.698
$(\overline{P_j} - \overline{P_i})$ 25-34	– 0.007	– 2.435	– 0.006	– 2.168	– 0.001	– 0.582
$(\overline{P_j} - \overline{P_i})$ 35-44	– 0.005	– 1.699	– 0.004	– 1.421	0.000	0.186
$(\overline{P_j} - \overline{P_i})$ 45-54	– 0.004	– 1.433	– 0.003	– 1.004	0.001	0.394
$(\overline{P_j} - \overline{P_i})$ 55-64	– 0.004	– 1.374	– 0.003	– 0.918	0.001	0.494
$(\overline{P_j} - \overline{P_i})$ male	– 0.002	– 0.940	– 0.002	– 0.946	– 0.002	– 1.033
Adj R^2	0.028		0.095		0.351	
N	2128		2128		2128	

Tabel 5.2	**Estimation results, total premium, WLS**					
	No firm specific coefficients		With firm specific coefficients α_i		With firm specific coefficients α_j	
Variable	Estimate	T-value	Estimate	T-value	Estimate	T-value
Age 35-44	– 0.338	– 2.466	– 0.115	– 1.953	– 0.116	– 3.222
Age 45-55	– 0.493	– 3.703	– 0.170	– 2.885	– 0.181	– 5.195
Age 55-64	– 0.625	– 4.906	– 0.233	– 4.052	– 0.230	– 6.845
Male	0.341	4.186	0.125	3.319	0.061	3.103
$(\overline{P_j} - \overline{P_i})$ 25-34	– 0.009	– 2.718	– 0.006	– 3.367	0.000	0.194
$(\overline{P_j} - \overline{P_i})$ 35-44	– 0.006	– 2.329	– 0.004	– 2.464	0.001	0.950
$(\overline{P_j} - \overline{P_i})$ 45-54	– 0.006	– 2.439	– 0.003	– 2.125	0.001	1.292
$(\overline{P_j} - \overline{P_i})$ 55-64	– 0.006	– 2.852	– 0.002	–1.580	0.001	2.294
$(\overline{P_j} - \overline{P_i})$ male	– 0.003	– 1.463	– 0.002	–1.487	– 0.001	– 1.075
Adj R^2	0.034		0.211		0.305	
N	2128		2128		2128	

Table 5.3 Estimation results, basic and supplementary premium, OLS

	No firm specific coefficients		With firm specific coefficients α_i		With firm specific coefficients α_j	
Variable	Estimate	T-value	Estimate	T-value	Estimate	T-value
Age 35-44	– 0.321	– 2.716	– 0.319	– 2.822	– 0.321	– 3.322
Age 45-55	– 0.450	– 3.792	– 0.438	– 3.862	– 0.455	– 4.700
Age 55-64	– 0.579	– 4.823	– 0.561	– 4.886	– 0.622	– 6.340
Male	0.345	4.085	0.335	4.150	0.338	4.903
$(\overline{P}_j^B - \overline{P}_i^B)$ 25-34	– 0.008	– 1.381	0.000	– 0.038	– 0.010	– 1.941
$(\overline{P}_j^B - \overline{P}_i^B)$ 35-44	– 0.004	– 0.671	0.003	0.602	– 0.006	– 1.167
$(\overline{P}_j^B - \overline{P}_i^B)$ 45-54	– 0.005	– 0.931	0.002	0.352	– 0.008	– 1.456
$(\overline{P}_j^B - \overline{P}_i^B)$ 55-64	0.000	0.003	0.007	1.220	– 0.002	– 0.441
$(\overline{P}_j^B - \overline{P}_i^B)$ male	– 0.010	– 1.936	– 0.009	– 1.956	– 0.009	– 2.284
$(\overline{P}_j^S - \overline{P}_i^S)$ 25-34	– 0.006	– 2.295	– 0.008	– 2.537	– 0.001	– 0.363
$(\overline{P}_j^S - \overline{P}_i^S)$ 35-44	– 0.005	– 1.693	– 0.006	– 1.907	0.001	0.248
$(\overline{P}_j^S - \overline{P}_i^S)$ 45-54	– 0.004	– 1.279	– 0.004	– 1.351	0.002	0.641
$(\overline{P}_j^S - \overline{P}_i^S)$ 55-64	– 0.005	– 1.547	– 0.005	– 1.562	0.001	0.343
$(\overline{P}_j^S - \overline{P}_i^S)$ male	– 0.001	– 0.394	– 0.001	– 0.390	– 0.001	– 0.387
Adj R^2	0.0274		0.0951		0.3536	
N	2128		2128		2128	

Table 5.4 Estimation results, basic and supplementary premium, WLS

	No firm specific coefficients		With firm specific coefficients α_i		With firm specific coefficients α_j	
Variable	Estimate	T-value	Estimate	T-value	Estimate	T-value
Age 35-44	– 0.340	– 2.559	– 0.114	– 1.901	– 0.140	– 3.711
Age 45-55	– 0.499	– 3.865	– 0.172	– 2.875	– 0.213	– 5.838
Age 55-64	– 0.630	– 5.100	– 0.233	– 4.005	– 0.266	– 7.377
male	0.345	4.259	0.125	3.298	0.060	2.827
$(\overline{P}_j^B - \overline{P}_i^B)$ 25-34	– 0.008	– 1.415	– 0.006	– 2.514	– 0.008	– 4.479
$(\overline{P}_j^B - \overline{P}_i^B)$ 35-44	– 0.005	– 0.953	– 0.003	– 1.704	– 0.004	– 3.488
$(\overline{P}_j^B - \overline{P}_i^B)$ 45-54	– 0.005	– 1.160	– 0.003	– 1.470	– 0.003	– 2.336
$(\overline{P}_j^B - \overline{P}_i^B)$ 55-64	– 0.002	– 0.438	– 0.002	– 1.052	– 0.002	– 1.694
$(\overline{P}_j^B - \overline{P}_i^B)$ male	– 0.009	– 2.142	– 0.002	– 0.922	0.000	0.420
$(\overline{P}_j^S - \overline{P}_i^S)$ 25-34	– 0.009	– 2.791	– 0.007	– 3.727	0.001	1.107
$(\overline{P}_j^S - \overline{P}_i^S)$ 35-44	– 0.006	– 2.547	– 0.005	– 3.123	0.001	1.823
$(\overline{P}_j^S - \overline{P}_i^S)$ 45-54	– 0.006	– 2.575	– 0.004	– 2.834	0.001	1.812
$(\overline{P}_j^S - \overline{P}_i^S)$ 55-64	– 0.007	– 3.436	– 0.004	– 2.555	0.001	2.455
$(\overline{P}_j^S - \overline{P}_i^S)$ male	– 0.002	– 0.766	– 0.002	– 1.362	– 0.001	– 1.117
Adj R^2	0.0385		0.2074		0.3126	
N	2128		2128		2128	

5.2 Choosing between models

In order to choose between the six estimated models, we performed two sets of F-tests:

- First, we determined whether the model without firm specific coefficients is rejected in favour of models with firm specific coefficients; this was always the case.
- Second, we assessed whether equality of coefficients on basic and supplementary premium is rejected in favour of including these as two separate explanatory variables; this was also the case.

Therefore, on statistical grounds we are able to narrow the number of models down to just two: with separate price variables basic and supplementary premium, either with firm specific coefficients for former regional monopolists (α_i) or with firm specific coefficients for firms that are new to the region (α_j).[6]

5.3 Sensitivity checks

In order to test the robustness of our findings to changes in the underlying assumptions, we did two sensitivity checks. First, in selecting the municipalities to be included in the analysis, we raised the cut-off point from 60% to 75% (see section 4). This lead to discarding 1,39 million observations and slightly higher (and statistically more significant) estimates. Second, rather than using average prices for the period 1993-2002 we included average prices for the period 2000-2002. Again this yielded essentially the same results, which is consistent with the fact reported in section 4 that differences in premium tend to be stable over time.

6 Looking at the adjusted R-squared, the equations with α_j consistently have the highest explanatory power.

6 Elasticities

6.1 Calculating elasticities

Because of the way we constructed our dataset, calculating elasticities on the basis of the estimated coefficient is rather complicated. The elasticity we are looking for is defined as follows:

$$\varepsilon_i = \frac{\Delta Q_i}{\Delta \overline{P}_i} \frac{\overline{P_i}}{Q_{i,1993}} \tag{4}$$

where

ΔQ_i = net change in the number of insured at firm *i* during the period 1993-2002.
$Q_{i,1993}$ = the number of insured at firm *i* at the beginning of the period 1993-2002.
$\Delta \overline{P}_i$ = change in the average price over the period 1993-2002 of firm *i*.
$\overline{P}_i$ = the average price over the period 1993-2002 of firm i.

If we define S_{ij} as the number of switchers from firm *i* to firm *j*, we can replace $\Delta Q_i / \Delta \overline{P_i}$ in eq. (4) by:

$$\Delta Q_i \ / \Delta \overline{P_i} = \left[\sum_{j=1,\, j\neq i}^{j=20} \Delta S_{ij} - \sum_{k=1,\, j\neq i}^{k=15} \Delta S_{ki} \right] / \Delta \overline{P_i} \tag{5}$$

Equation 5 indicates that the change in the number of insured of firm *i* as a consequence of raising its price, equals the change in each of the bilateral flows in which firm *i* is involved. The first term between brackets represents the change in the flows *from* firm *i* to other firms as a consequence of firm *i* raising its price. The second term between brackets indicates the flows *to* firm *i* from other firms as a consequence of firm i raising its price. These flows can occur only from one of the 14 other firms which held a regional monopoly in 1993.

Next we substitute F_{ij} for S_{ij} using the following equation (this follows directly from the definition of F_{ij} in eq. 1-3):

$$F_{ij} = 100 \cdot S_{ij} / Q_{i,1993} \tag{6}$$

where:

F_{ij} = $F_{i \to j,\, j \neq i}$, see equation(1)

$$\Delta Q_i / \Delta \overline{P_i} = \frac{Q_{i,1993}}{100} \left[\sum_{j=1,\, j\neq i}^{j=20} \Delta F_{ij} / \Delta \overline{P_i} - \sum_{k=1,\, j\neq i}^{k=15} \Delta F_{ki} / \Delta \overline{P_i} \right] \tag{7}$$

Since we are interested in the elasticity for an average firm, we replace $Q_{i,1993}$ by Q^*, defined as the number of insured of the average incumbent firm. Furthermore we can calculate dF_{ij}/dP_i and dF_{ji}/dP_i in eq. 7 by differentiating eq. 1-3 with respect to price (note that recall that $\Delta\overline{P_i}$ is defined as $(\overline{P_j} - \overline{P_i})$, implying that $\Delta F_{ji} / \Delta\overline{P_i} = -\Delta F_{ji} / \Delta\overline{P_i}$). Using the fact that the first term in eq. 7 is a summation over 19 identical terms while the second term is a summation over 14 identical terms we arrive at the following equation:

$$\Delta Q_i / \Delta\overline{P_i} = \frac{Q^*}{100}\left[\sum_{j=1,\, j\neq i}^{j=20}\alpha_{price} - \sum_{k=1,\, j\neq i}^{k=15} - \alpha_{price}\right] = \frac{33\alpha_{price}Q^*}{100} \tag{8}$$

where:

α_{price} = the (sum of the) estimated coefficient(s) on price in eq. 1 – 3 for the relevant group; for example, for 25-34 year old working males, we would add up the coefficients on the price for this age group and the coefficient on gender.

Substituting (8) in (4) and dropping the subscript *i* (since we are interested in the elasticity for the average firm) yields:

$$\varepsilon = 33 \cdot \overline{P} \cdot \alpha_{price} / 100 \tag{9}$$

This elasticity can be used to answer the following question: how much smaller (in %) would the number of insured of an average firm have been in 2002, if its average price during the period 1993-2002 had been 1% higher, assuming all other firms had kept their prices at the observed level?

6.2 Results

Table 6.1 presents elasticity estimates for our preferred specifications. We find plausible age effects: estimated elasticities fall (in absolute value) with age. Also we find consistently that men are more responsive to price than are women. Perhaps the most striking result is that these elasticities are so small. Recall that these elasticities measure the cumulative effect after 7 years of keeping price 1% above the price of competitors. However, this does not means that we can derive annual elasticities simply by dividing the elasticities reported in Table 6.1 by seven. Some insured may have switched to another firm early on (say in 1997), and stayed with that firm because this turned out to be the right choice in following years. Given the high correlation of prices over time noticed in section 4, this is a plausible scenario. This reasoning suggests that the elasticities in Table 6.1 should be interpreted as an upper limit for the annual elasticity.

We should also point out that, since we are restricting our data to those who where insured continuously during the period 1993-2002, we are omitting new customers: new entrants to the labour market and the self-employed. These groups are likely to be more price-sensitive (see Hassink and Schut (2002)).

Table 6.1 Elasticities

	Based on equations including α_i		Based on equations including α_j	
	Women	Men	Women	Men
Basic insurance				
Age 25-34	– 0.30	– 0.38	– 0.41	– 0.39
Age 35-44	– 0.17	– 0.25	– 0.22	– 0.19
Age 45-55	– 0.14	– 0.22	– 0.15	– 0.12
Age 55-64	– 0.10	– 0.18	– 0.10	– 0.08
Supplementary insurance				
Age 25-34	– 0.18	– 0.22	0.03	0.01
Age 35-44	– 0.13	– 0.17	0.03	0.01
Age 45-55	– 0.12	– 0.16	0.03	0.01
Age 55-64	– 0.10	– 0.14	0.04	0.02

7 Conclusions

The main findings are as follows: the price elasticity of residual demand for social health insurance was low during the period 1996-2002. We find small but significant effects of the price of basic insurance but no robust effect of the price of supplementary insurance. Young enrollees are more price sensitive than older enrollees. However, these findings are conditional on the limited variation in price observed in our data. At larger price differentials, the elasticity may well be higher. We also stress that we had to make quite a few assumptions in order to estimate price elasticities from the data at hand. Although we believe that these assumptions are realistic, they are not entirely correct for all our observations. As a result, our estimates may be biased. Clearly, it would have been preferable to work with real paneldata, but these are only available at an aggregate level. Given the observed price differences, the large market shares of all former regional monopolist in their former designated regions are consistent with a low price elasticity of residual demand.

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