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Risk adjustment in the Netherlands

An analysis of insurers' health care expenditures

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Abstract in English

As of 2006, the Dutch healthcare system will be run by regulated competition. An important part of regulated competition is a system of risk adjustment. This paper presents an empirical analysis of the effects of risk adjustment in the Dutch social health insurance system covering the years 1991-2001. By comparing insurers' health care expenditures with their risk adjusted premiums, our analysis estimates the impact of risk adjustment over a number of years. Results indicate that the risk-adjustment system has improved substantially. Whereas in the beginning of the nineties prospective risk adjustment could explain about 20% of the variation in health care expenditure differentials between insurers, this figure rose to 55% in 2001. The explanation of the same variation after retrospective payments did not show a clear upward or downward trend, and has varied since 1995 around 85%. The remaining variation in insurers' health care expenditure differentials are determined more by structural than random factors. One such factor may be related to the low ex-ante projections of the government's total health care expenditures, which favour insurers with a population of relatively good health risks. Results show that new entrants in the Dutch health insurance market had significantly lower health care expenditures. Furthermore, economies of scale do not seem to have played a role during the sample period: the expenditures of large insurers were not significantly lower than those of the smaller insurers.
Key words: risk adjustment, health care expenditure, health care insurers

Abstract in Dutch

Deze studie onderzoekt in hoeverre verschillen tussen gezondheidszorguitgaven van ziekenfondsen worden verklaard door risicoverevening gedurende 1991-2001. We vinden dat het systeem van risicoverevening sterk is verbeterd. Kon prospectieve verevening in het begin van de jaren negentig ongeveer 20% van de variatie in de gezondheidszorguitgaven tussen ziekenfondsen verklaren, in 2001 liep dit percentage op tot ongeveer 55%. De verklaring van dezelfde variatie na retrospectieve verevening bleef na 1995 ongeveer gelijk en schommelde rond de 85%. De studie laat zien dat de overgebleven variatie in de uitgaven voor een groter gedeelte nog bepaald worden door structurele dan willekeurige factoren. Een structurele factor, die hier mogelijk een rol speelt, is de lage projectie van het macrobudget waardoor verzekeraars met een relatief gezondere populatie bevoordeeld worden. De studie laat verder zien dat nieuwe toetredende ziekenfondsen gedurende 1991-2001 significant lagere uitgaven hadden. We vinden geen aanwijzingen dat schaalgrootte een belangrijke rol heeft gespeeld: grote ziekenfondsen hadden geen significant lagere uitgaven dan kleine ziekenfondsen.
Steekwoorden: risicoverevening, gezondheidszorguitgaven, ziekenfondsen

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Contents

Summary

The Dutch government has decided that the new health care system in 2006 will be one of regulated competition. Some of the key elements are compulsory health insurance for the entire population, one basic benefit package, open enrolment, equal nominal premiums for insured persons affiliated with the same health care insurer and competing health care insurers and health care providers. One important element of regulated competition is a system of risk adjustment. Risk adjustment refers to the practice of paying insurers prospectively a subsidy per person that is related to the expected health care expenditures of that individual. The system of risk adjustment levels health care expenditure differentials between insurers that are due to differences in their population mix. This is important, since insurers are obliged to accept all enrolees for the same flat rate premium. In the absence of risk adjustment, insurers with a less healthy population will have a competitive disadvantage, as they must charge higher nominal, community-rated, premiums. A second advantage of risk adjustment is that it increases efficiency by reducing risk-selection incentives. If after risk adjustment the expected health expenditures of persons are similar, then insurers have no incentives to attract favourable enrolees.

Ideally, the risk-adjustment system includes all relevant factors for which society desires solidarity (such as age, gender and health status of the insured population), and excludes all factors that insurers are expected to influence under regulated competition (such as efficiency efforts). In the ideal case, the incentives for efficiency are optimal, and the incentives for risk selection are minimal.

Yearly payments to insurers consist of two parts. At the beginning of the year insurers receive prospective risk-adjusted payments; at the end of the year, after realisation of health care expenditures, they receive retrospective payments. By comparing insurers' health care expenditures with their prospective (and retrospective) payments, we estimate the impact of risk adjustment over a number of years. Results indicate that the risk-adjustment system has improved substantially. Whereas in the beginning of the nineties prospective risk adjustment could explain about 20% of the variation in health care expenditure differentials between insurers, this figure rose to 55% in 2001. The explanation of the same variation after retrospective payments did not show a clear upward or downward trend, and has varied since 1995 around 85%. The latter result suggests that the increase in insurers' financial risk coincided with improvement of the risk-adjustment system.

The remaining expenditure variation between insurers can be substantial. Consider for a moment only expenditure differentials after risk-adjusted prospective payments. During 1997-2001, the average insurer spent about 244 euros per capita per year (2001 prices) more than the lowest expenditure insurer (or 19% of 1300 euros, the mean per capita health care expenditures of the total population). Omitting the five lowest expenditure insurers, this figure declines

drastically to about 61 euros per capita. The same numbers, of course, decline after the retrospective payments. Then we find 69 euros and, respectively, 20 euros.

Are the remaining expenditure differentials between insurers structural or random? The latter case presents no problem, since lucky outcomes for an insurer in some years will be compensated by unlucky outcomes in others. Our findings suggest, however, that differences between insurers are more structural than random, which implies that some insurers have structurally lower health care expenditures.

One structural factor may be related to the practical implementation of risk adjustment: the prediction of the so-called ‘macro budget’. In the Netherlands this macro budget is determined by the government and is used as an ex-ante projection of total expected health care expenditure for the coming year. We show theoretically that if this projection turns out to be too low ex-post, then it favours insurers with a population of relatively good health risks. During the sample period the ex-ante projection turned out to be structurally too low. For example, in 2001 a 1.4% prediction error of the macro budget resulted in differences of mean per capita prospective payments between insurers ranging from –4.8 euros to 3.2 euros per capita. A simple solution for this problem would be to recalculate the prospective risk-adjusted subsidies retrospectively, and then correct for these differences.

We also explored the possibility that efficiency in purchasing medical care may have played a role by testing whether economies of scale are important in our sample period. Estimations of panel data models showed that new small entrants on the insurance market have on average (much) lower health care expenditures. Precisely why they have lower expenditures remains unclear, although the fact that new entrants since 1991 were almost all private health insurers may have played a role. After omitting the entrants from the panel, we found no evidence that economies of scale played a role during 1991-2001.

1 Introduction

The process of restructuring the health care sector has been fraught with struggle. Policymakers face difficult trade-offs between solidarity, quality, efficiency and macro expenditure containment. As in many countries, also in the Netherlands policymakers are pursuing a route to introduce more incentives for efficiency without harming, insofar as possible, other aspects of health care. Although some of the reforms have yet to take place, the decision has been taken that the new health care system will be one of regulated competition. Some of the key elements are compulsory health insurance for the entire population, one basic benefit package, open enrolment, equal nominal premiums for insured persons affiliated with the same health care insurer and competing health care insurers and health care providers (VWS, 2002).

One important element of regulated competition is a system of risk adjustment. Risk adjustment refers to the practice of paying insurers prospectively a subsidy per person that is related to the expected health care expenditures of that individual. The system of risk adjustment levels the health care expenditure differentials between insurers that arise from differences in their population mix. This is important, since insurers must accept all enrolees for the same flat-rate premium. Without risk adjustment, insurers with a less healthy population will face a competitive disadvantage, as they have to charge higher nominal, community-rated, premiums. A second advantage of risk adjustment is that it increases efficiency by reducing risk-selection incentives. If after risk adjustment the expected health care expenditures of persons are similar, then insurers have no incentives to attract favourable enrolees.

Ideally, the risk-adjustment system includes all relevant factors for which the society desires solidarity (such as the age, gender and health status of the insured population), and excludes all factors that insurers are expected to influence under regulated competition (such as efficiency efforts). In the ideal case the incentives for efficiency are optimal, and the incentives for risk selection are minimal. The aim is thus not to minimise expenditure differentials between insurers in general, but to minimise only those expenditure differentials between insurers that can be explained by risk factors for which society desires solidarity (Van de Ven and Ellis, 2001).

Although risk adjustment is present in many countries, the effectiveness of existing risk-adjustment models remains uncertain. In a hypothetical experiment, Shen and Ellis (2001) show that insurers could make significant profits, even after risk adjustment, if they could dump their unprofitable enrolees. Van de Ven *et al.* (2004) show that in the Dutch social health care system the average predicted losses in year *t* for the ten percent of patients with the highest expenditures in year *t*-1 is still higher than 1000 euros, even when considering sophisticated risk adjustment models. Van de Ven *et al.* (2002) studied risk adjustment in five countries (Belgium, Germany, Israel, the Netherlands and Switzerland), and concluded that as of 2001, it is still imperfect. The authors found that sickness funds have financial incentives for risk selection that may threaten solidarity, efficiency, quality of care and consumer satisfaction. Given the impossibility of perfect risk adjustment, the incentives for risk selection will persist to a certain extent.

While in many countries incentives for risk selection may be (partly) neutralized by risk adjustment there are also hurdles placed in the path of insurers attempting to apply risk selection. For example, institutional arrangements such as open enrolment may make it difficult for insurers to refuse unprofitable enrolees. Also reputation mechanisms may play a role; an insurer may be put in the pillory if creaming, skimping or dumping of enrolees becomes publicly known. A recent study of the Dutch social health insurance market concludes that no significant risk-selection activities by insurers have ever taken place (CPB, 2003). Currently, public insurers in the Netherlands may also see no need for selection activities, as they are already at the maximum permitted level of their financial reserves, or they may have easier instruments at their disposal to create profit, such as raising the out-of-pocket premiums. The latter argument may stem from the fact that Dutch consumers show a low sensitivity to differences in out-of-pocket premiums (Schut and Hassink, 2002). As in most countries, also in the Netherlands we observe low consumer mobility, which suggests that expenditure differentials between insurers will be for the most part determined by the majority of enrolees who stay put and do not switch from insurer.

This paper studies the magnitude of these health care expenditure differentials between insurers, and how these differentials are influenced by risk adjustment. Even if risk-selection activities are relatively unimportant, for insurers it is vital that risk-adjusted payments are adequate. Even the most efficient insurers may not be able to bring their out-of-pocket premiums down to the same level as their competitors, and may thus fail to survive in the long run, if they are seriously harmed by imperfect risk adjustment.

The subject is highly relevant in practice, since in many countries insurers (or hospitals) complain to the government if they believe that their risk-adjusted payments are incorrect or are too low in comparison with individual risk profiles of the entire market. In the Netherlands, insurers report complaints to the government, which has established an independent research group to study complaints and to improve the risk-adjustment system. Another example comes from Australia. The Alfred Hospital performed a benchmark study of ten hospitals to show the payment gap, due to imperfect risk adjustment, between the Alfred Hospital and other hospitals. This resulted in more equitable provision of grants for certain diseases across the entire hospital industry (Antioch and Walsh, 2000).

Yearly payments to insurers consist of two parts. At the beginning of the year insurers receive prospective risk-adjusted payments; at the end of the year, after realisation of health care expenditures, they receive retrospective payments. By comparing insurers' health care expenditures with their prospective (and retrospective) payments, we estimate the impact of risk adjustment over a number of years. Results indicate that prospective risk-adjusted payments have improved substantially. The main reason for this improvement is the inclusion of new risk adjusters which has leaded to better predictions of insurers' expected health care expenditures.

This development may reduce the negative side effects of regulated competition, such as risk selection, and enable the government to increase gradually sickness funds' efficiency incentives. The paper explores how large the remaining differences between insurers' health care expenditures are after risk adjustment and whether they are determined randomly or by structural factors. The conclusion is that structural factors still play a dominant role. The possible source for the remaining structural factors may be related to a myriad of differences between insurers, such as imperfect risk adjustment or efficiency differences. One structural factor, in particular, may be related to the low ex-ante projections of total health care spending of the government, which favour insurers with a population of relatively good health risks. Although difficult to determine, other possible structural factors are also considered.

This paper contributes to the literature by analysing the quantitative consequences of (imperfect) risk adjustment at the insurer level. Our data includes *all* insurers on the Dutch social health care market. In the literature we could not find any similar type of study that considers also *all* competing insurers on a health care market. Acquiring data at the insurer level may be difficult in practice, may be inaccessible, or may be kept in private hands. We obtained the data (not publicly available) from the government institution responsible for the risk-adjustment system in the Netherlands.

The paper is organised as follows. Section 2 describes the Dutch social health insurance system, the risk-adjustment system and the data. A measure for risk-adjustment performance is defined and computed in section 3. Section 4 shows that the low ex-ante predictions of total health care expenditures of the government may favour insurers with a relatively healthy population. After a comparison of health care expenditure patterns of individual insurers in section 5, section 6 concludes.

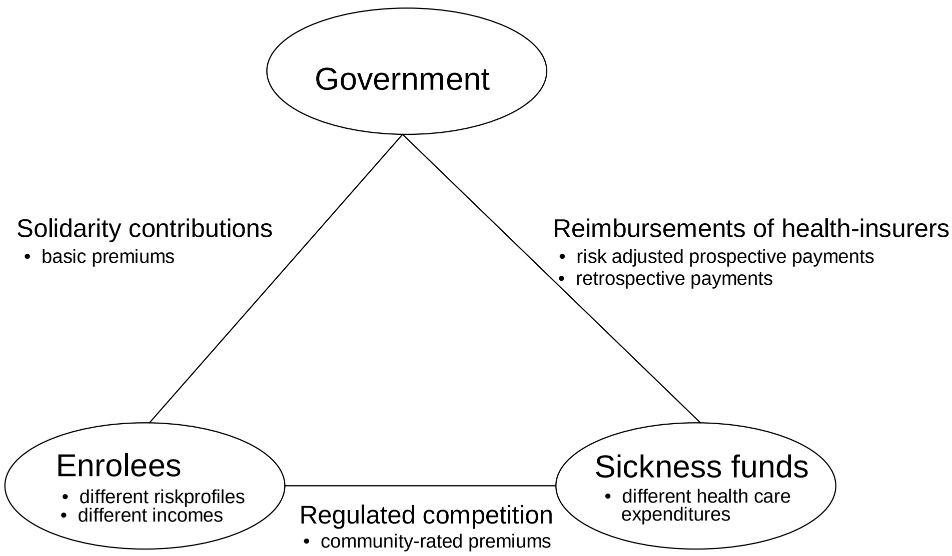
2 Risk adjustment in the Netherlands

Before turning to the empirical results of the study, we first explain some institutional characteristics of the Dutch social health insurance system, with a particular focus on the system of risk adjustment.[1] Finally, the last subsection describes the data.

2.1 The Dutch social health insurance system

Figure 2.1 illustrates the public scheme for health insurance as administered by social health care insurers (sickness funds) in the Netherlands. This public scheme regulates insurance for those with labour income below a certain threshold (about 31,000 euros in 2002). Insurance is obligatory for those who are eligible, and covers about two-thirds of the Dutch population (about 10 million people). The scheme covers health care expenditure on basic cure services including hospital care, care delivered by general practitioners and pharmaceuticals. Depending on the year, it is administered by 24-30 independent sickness funds. Enrolees may change yearly from insurer, and acceptance is obligatory. Enrolees face equal basic benefit packages, as designed by the government, and two types of premiums: a basic and a community-rated premium.

Figure 2.1 The health insurance system as administered by Dutch sickness funds



[1] It is almost inevitable that this section contains some overlap with previous papers about Dutch social health insurance, such as Lamers *et al*. (2003) and Schut and Hassink (2003).

The basic premium is uniform across sickness funds and is a function of the income of the insured person. This contributes to income solidarity.[2] In figure 2.1 these payments reflect the solidarity contributions of the consumer to the Dutch government. The government collects these contributions and reimburses them across sickness funds. Two types of payments by the government can be distinguished: risk-adjusted prospective payments and retrospective payments. These will be explained in greater detail in the next section.

Thus far we have explained the upper part of figure 2.1; basic premiums of the consumer flow through the government to the health insurer. The main point is that the government regulates these flows. The government is thus able to introduce a variety of solidarity issues into the insurance system, such as risk and income solidarity across enrolees and expenditure solidarity between insurers.

However, in order to stimulate insurers to implement more efficient activities, the Dutch government introduced elements of regulated competition. This is visible in the lower part of figure 2.1. Health insurers are also allowed to raise a flat rate (community-rated) premium.

These nominal premiums are set by individual sickness funds. However, sickness funds are not allowed to differentiate across different risk categories. All enrolees contracted by the same insurer, except children under eighteen, pay identical nominal premiums. Since 1996, nominal premiums are allowed to differ across insurers, which enables insurers to compete on price. This element of regulated competition is expected to increase the efficiency efforts of insurers. Less efficient insurers will arguably have higher expenditures and will therefore have to set higher community-rated premiums than more efficient insurers.

2.2 Prospective risk adjustment

For the Dutch government, the process of prospective risk adjustment contains three crucial steps:

- Determining a payment formula to predict individual risks.
- Predicting total expected expenditure for the next year (the so-called ‘macro budget’).
- Determining risk-adjusted subsidies for insurers by combining (1) and (2).

The payment formula to predict individual risks uses so-called risk adjusters. Table 2.1 presents an overview of risk adjusters used over the years. Risk adjustment started in 1991with ‘historical expenditures’ or prior yearly expenditures of sickness funds as the only risk adjuster.[3] In 1992, the normative risk adjusters ‘age’ and ‘gender’ were only used for 20% of

[2] This element is not central in a scheme of regulated competition, however. Indeed, income solidarity can also be achieved in other ways, including through the tax system.
[3] The prior yearly expenditures were determined at the insurer level. The disadvantages of this risk adjuster are well known (see Van de Ven and Ellis, 2002).

Table 2.1	**Risk adjusters in the Dutch prospective risk-adjustment system**
Year	Risk adjusters
1991	historical expenditure
1992	historical expenditures, age, gender
1993-1994	age, gender
1995-1998	age, gender, region, disability status
1999	age, gender, region, employment or social security status
2000-2001	age, gender, region, employment or social security status, historical expenditures

the budget and 80% was still explained by 'historical expenditures'. The first two years were seen as probationary years, and after 1992 only normative risk adjusters were, as much as possible, taken into consideration. From 1993-1994 'historical expenditures' were abandoned, leaving 'age' and 'gender' as the only risk adjusters used. Further refinements took place gradually. First of all, expenditures were split up into different categories. For example, in 1998 four types of different budgets could be distinguished: variable and fixed hospital outlays, inpatient medical specialist care and other medical expenses.[4] For each of these budgets the same risk adjusters were used, but with different weights. The number of risk adjusters increased as well. During 1995-2001, besides 'age' and 'gender', three additional risk adjusters were introduced. The regional risk adjuster was based on the degree of urbanisation of a region. The risk factor explaining the disability status of a person was updated in 1999 when also the employment and social security status of persons were considered. Since 2000, 'historical expenditures' was introduced as the risk adjuster for the budget 'other medical expenses'. 30% of this budget (on the insurer level) is determined by the average expenditures of medical outlays of the three previous years. However, in 2003 the historical expenditures risk adjuster was abolished, since it hampered the incentives for efficiency and seemed to add little to the explanatory power of the system.

Although not included in our sample, further refinements have already taken place. In 2002, the inclusion of pharmacy expenditure groups improved the predictions for especially the chronically ill, and in 2004, diagnosis expenditure groups (DCG) were added to the risk-adjustment model (Lamers *et al*., 1999, Van de Ven *et al.* 2004).[5]

Risk-adjusters, with corresponding weights, can thus help to explain health care expenditures in the past. Determination of the risk-adjusted subsidies for the coming year, however, requires additional information. Therefore the Dutch government decides each year the total amount of money for financing the insurers, which is related to the basic benefit package (or also called acceptable expenditures; see Van de Ven and Ellis, 2000). This so-called 'macro budget' will be equal to the sum of the individual budgets of the social insurers. In order to obtain the adequate payment distribution between insurers, the weights of the risk

[4] More information can be found in the Appendix.

[5] Following US policy, which introduced DCG's on 1 January 2000 in the monthly payments from medicare to HMOs (see e.g. Pope *et al.* 2000)

adjusters are scaled so that the ex-ante macro budget exactly equals the sum of the prospective budgets of insurers.[6]

2.3 Retrospective payment schemes

If prospective risk adjustment were the only form of risk adjustment, then sickness funds would be fully financially responsible for their health care expenditures. As perfect risk adjustment is still a long way off, governments have introduced retrospective payment schemes to reduce selection. These payment schemes imply that sickness funds are retrospectively reimbursed for some of their expenditures. Various retrospective payment schemes were introduced in the Netherlands during 1991-2001:

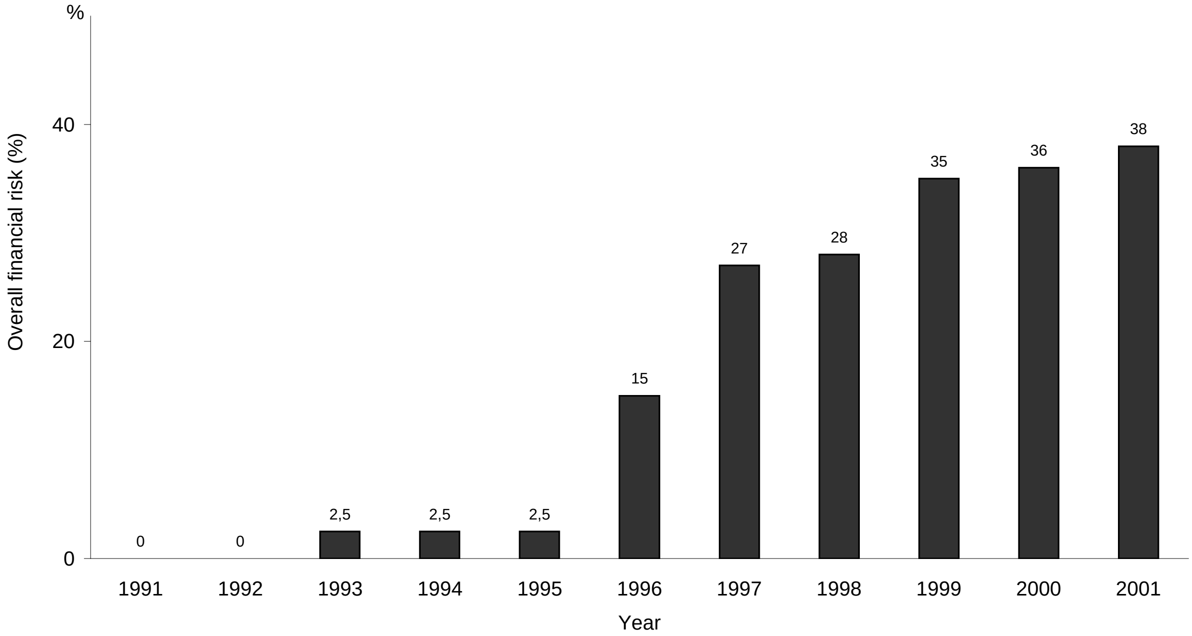
- *Retrospective equalisation*. The idea of retrospective equalisation is that some sickness fund expenditures are transferred to a pool. The money in this pool is shared retrospectively, to a certain percentage, by insurers. Losses of some insurers will thus be compensated by gains of others. Retrospective equalisation is budget-neutral for the government, and therefore involves no extra subsidies.
- *High-risk equalisation*. This is special form of retrospective equalisation. High-risk equalisation implies that individual expenditures (above a certain threshold) will be transferred to a pool. The money in this pool will again be shared, to a certain percentage, across insurers.
- *Retrospective compensation*. Each sickness fund receives from or pays to the government a fixed percentage of some of their losses or profits. Retrospective compensation is a form of risk sharing that leads to additional losses or profits for the government.

Once actual expenditures become known, the government applies the retrospective payment schemes in the same order as explained above. More detail regarding the three retrospective payment schemes can be found in the Appendix .

If differences in expenditures between insurers are related to differences in the risk pool, which are not captured by the risk-adjustment scheme, or to unlucky outcomes, then retrospective payment schemes may sound sensible. After all, one can argue that an insurer should not be punished for imperfect risk adjustment or bad luck. On the other hand, retrospective payments seem less sensible if differences in expenditures are due to differences in efficiency, since insurers should not be penalized for their efficient behaviour. In terms of incentives, a greater role for retrospective compensation means stronger disincentives for both efficiency and risk-selection strategies. This marks the trade-off between efficiency and risk selection (Newhouse, 1996).

[6] See Lamers *et al.* (2003) for more information on this subject.

Figure 2.2 Sickness funds' overall financial risk (in %) in the period 1991-2001



Source: Van de Ven et al. (2004).

The more expenditures are retrospectively reimbursed, the less risk borne by insurers. By combining the three different types of retrospective payment schemes, one can calculate the percentage of overall financial risk borne by insurers in a regulated competition environment during the period examined. An example: in 1993, 75% of the expenditures were retrospectively equalised, which reduced the overall financial risk of sickness funds to 25%. However, 90% of these expenditures were retrospectively compensated, which reduced overall financial risk to 2.5%. Overall financial risk rose to 38% in 2001 (see Figure 2.2).

Although falling outside of our sample, overall financial risk in 2004 has already increased to 52%.[7]

2.4 Description of the data

Risk adjustment for Dutch social health insurers is mandatory, and the CVZ ('College voor Zorgverzekeringen') is responsible for the implementation of risk adjustment. CVZ provided us with data on all insurers for 1991-2001; this data includes for each insurer the following: prospective and retrospective payments, actual expenditures and number of enrolees.

Some characteristics of the social insurance market are presented in table 2.2. We observe that the number of insurance companies fluctuated, although a declining trend seems to have started in the new Millennium. After an initial increase, the number of small insurance

[7] Many researchers claim a link between this result and the increase in the variation of out-of-pocket premiums (see e.g. Douven and Westerhout (2001), Lamers *et al.* (2003) and Schut and Hassink (2002)).

companies seemed to decrease as well. Almost all insurers leaving during the sample period merged with other insurers, while new insurers entering the market started always with a

Table 2.2 Some characteristics of the Dutch social health insurance market

Year	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
Population size (millions)	9.2	9.3	9.4	9.6	9.7	9.8	9.9	9.9	9.9	10.3	10.3
Total number of insurers	30	25	25	25	26	27	29	29	29	26	24
Number of insurers leaving	n.a.	6	2	0	1	0	0	2	0	3	2
Number of insurers entering	n.a.	1	2	0	2	1	2	2	0	0	0
Insurers with population < 10,000	0	1	2	1	3	3	5	6	6	2	1
Insurers with population < 50,000	0	1	3	3	4	5	7	8	8	6	5
Insurers with population < 100,000	2	2	3	4	5	6	8	10	10	7	6
Insurers with population > 500,000	4	5	7	7	7	7	7	8	8	9	9
Insurers with population > 800,000	0	1	1	1	2	2	2	3	3	4	4
HHI	414	539	572	566	625	622	622	663	663	707	711

relatively small population. The data show also a clear trend of insurance companies becoming larger. This is confirmed in the last row by the concentration measure HHI (Herfindahl-Hirschmann index), which shows an upward trend.[8]

To obtain comparable yearly expenditure indicators, we expressed all expenditures in 2001 prices, where inflation of prices is assumed to follow inflation of average actual health expenditures per capita. Per capita health care expenditures in 2001 were 1297 euros. Table 2.3 shows our measure for price inflation.

Table 2.3 Our measure for price inflation

Year	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
Price (2001=1)	0.62	0.53	0.55	0.57	0.55	0.74	0.79	0.84	0.87	0.91	1.00

Note that we do not observe a continually increasing trend in per capita expenditures over the years. For example, per capita expenditure was lower in 1992-1995 than in 1991. A number of reasons can be cited, the most important of which is the variableness over time of the basic benefit package. A broader basic benefit package, moreover, is more expensive than a smaller one. For example, a part of pharmaceutical expenditures, audiology, revalidation, and genetic research went out of the basic benefit package of sickness fund insurance in 1992 and was transferred into the so-called AWBZ or Exceptional Medical Expenses Act. These allowances were transferred back again, however, in 1996. Other reasons may be regulatory changes, such as changes in tariffs by the National Health Tariffs Authority.

[8] The HHI is equal to the sum of the squared market shares of the sickness funds.

3 Performance of risk adjustment at the insurer level

The larger the differences are in the population mix (or health care profiles) between insurers, the greater the need for risk adjustment at the insurer level. Differences in insurers' health care profiles can be obtained through a comparison of risk-adjusted prospective budgets. We first define the following statistics for insurer *i* in year *t*:

B_{it} : risk-adjusted prospective budget, B_t :total yearly budget or 'macro budget' ($B_t = \sum B_{it}$).
N_{it} : population of insurer *i*, N_t :total population ($N_t = \sum N_{it}$)

Subsequently, we define the following per capita variables $b_{it} = B_{it} / N_{it}$ and $b_t = B_t / N_t$. An empirical measure for the health care profiles (*hcp*) of the population of insurer *i* can now be obtained by $hcp_{it} = b_{it} / b_t$. An hcp_{it} equal to one would now imply that the population of insurer *i* in year *t* has a health care profile that is equal to the average health care profile of the total population. Table 3.1 shows the minimum and maximum for hcp_{it} for the sample years. Health care profiles among insurers differ. The largest difference seemed to occur in 1994, when the ex-ante health care expenditures of one insurer were estimated as 1.75 times higher than the average health care profile. Small insurers will generally exhibit higher levels of variation in health care profiles than larger insurers. Indeed, if we consider only those insurers with a population greater than 50,000, then the difference between the minimum and maximum health care profiles declines substantially.

Table 3.1 Insurers health care profiles (hcp_{it})

Year	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
MAX hcp_{it}	1.25	1.25	1.05	1.75	1.69	1.29	1.49	1.27	1.47	1.18	1.14
MIN hcp_{it}	0.89	0.88	0.74	0.89	0.85	0.86	0.69	0.69	0.66	0.66	0.69
MAX* hcp_{it}	1.25	1,25	1.05	1.09	1.69	1.29	1.45	1.21	1.17	1.18	1.14
MIN* hcp_{it}	0.89	0.88	0.96	0.94	0.92	0.95	0.93	0.94	0.95	0.93	0.94

*Omitting insurers with population < 50,000.

3.1 Prospective risk-adjustment performance

The literature offers two common indicators for measuring the effect of risk adjustment on the insurer level. The first is a quadratic function of ex-ante per capita subsidies, or predicted expenditures by the regulator, minus per capita actual expenditures (Ellis and Ash (1995) and Rosenkranz and Luft (1997)). The second indicator is a mean absolute function of the two variables (Ettner *et al.*, 1998). Rosenkranz and Luft (1997) explain the inappropriateness of the first indicator: '…a few data points with large deviations between actual and predicted values can overwhelm the resulting composite measure'. Similarly, Van Barneveld (2000) asserts that this indicator gives too much weight to outliers, and that it is by no means clear that insurers weight outliers in this way.[9] Given these considerations, the second indicator is our choice for this study. We now introduce definitions for the mean absolute result (*MAR*) and risk-adjustment performance (*RAP*) as follows:

$$MAR_t = \frac{1}{K_t}\sum_{i=1}^{K_t} | e_{it} - b_{it} | \ , \qquad MART_t = \frac{1}{K_t}\sum_{i=1}^{K_t} | e_{it} - b_t | \ , \qquad RAP_t = 1 - \frac{MAR_t}{MART_t} \ ,$$

where
E_{it} : actual health care expenditures, K_t : number of insurance companies,
e_{it} : actual mean per capita health care expenditures of insurers ($e_{it} = E_{it} / N_{it}$).

The value of *RAP* shows similarities with an R^2- measure in which *MART* is a measure for total variation in the data (without risk adjustment) and *MAR* is the variation left after risk adjustment.[10] Thus, *RAP* measures the extent to which risk adjustment reduces the variation, compared to a situation without risk adjustment that allocates to each enrolee the same subsidy. Before turning to the outcomes, however, we must clarify why the predicted ex-ante risk-adjusted budget, b_{it} , and actual expenditure, e_{it} , may deviate. We consider the following possibilities:

1. Unexpected events. Even perfect risk adjustment cannot predict luck or unexpected events. In general, one could say that the occurrence of unexpected events will complicate the prediction of individual health care expenditures and will result in a higher *MAR* and *MART*, but not necessarily a lower *RAP* statistic.
2. Imperfect risk adjustment. Imperfect risk adjustment may lead to prediction errors of individual health expenditures, which may result in a higher *MAR* and (crucial for our analysis) a lower *RAP* statistic. Note that as in (1) this is not necessarily the case, since prediction errors of the risk-adjustment system at the individual level can cancel out at the insurer level. For example, it depends also on how these prediction errors are distributed between insurers.
3. Management differences between insurers. All else being equal, if one insurer is more efficient, or produces lower quality care than another, then higher *MAR* and *MART* statistics generally result. Management differences between insurers were probably small during the sample period, but empirical evidence on this issue is scarce.[11]

[9] An interesting technical point here is that the weights in the risk-adjustment formulas are determined by least squares types of estimation methods. If we are more interested in absolute differences than quadratic differences, then it may also be more appropriate to determine the weights with absolute differences types of estimation methods.
[10] Note that we do not argue here that a *RAP* of 1 is optimal. The advantage of the *RAP* statistic is that it becomes possible to compare the performance of risk adjustment at the insurer level over a number of years.
[11] Schut and Hassink (2003) argue: 'Sickness funds ….. 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', The authors claim, furthermore, that nominal premiums are affected by unobserved fund heterogeneity, which may be caused by differences in administrative efficiency and differences between the price of health care suppliers in the region.

4. Mobility issues. Over time, expenditure differentials between insurers may be influenced by switchers (for example due to (adverse) selection), if switchers are cheaper than non-switchers within the same risk category. In general, the magnitude of this effect is probably low, since the population of insurers did not change substantially over time, and studies report no significant selection activities by insurers (CPB, 2003).[12] Another possibility to consider is that non-mobility leads to adverse retention, which is the tendency for people who stay put to magnify expenditure differentials between insurers; this they will do if they differ in age, and expenditures are more linear with age (Altman *et al.*, 1998). The presence of risk adjustment, however, will diminish this factor.
5. Changing market and policy characteristics. Market characteristics related to the insurance markets, such as mergers and new entrants may not be *MAR* or *RAP* invariant. Also policy characteristics, such as changes in the income threshold, changes in the basic benefit package, or changes in the prediction quality, etc., may influence our statistics.

Table 3.2 presents the outcomes of our statistics. *RAP* is relatively high in 1991 and 1992 (see also section 4.1). As explained in section 2.2, these were probationary years: in 1991 the only risk adjuster was historical expenditures, and in 1992 historical expenditures covered 80%, and age and gender 20%. The outcomes show that historical expenditures performed well as a risk adjuster. When it was dropped in 1993, and age and gender were the only risk adjusters, the *RAP*'s declined substantially. After 1993, however, risk adjustment improved from about 20% in 1993 to around 55% in 2001. The *MAR* in table 3.2 shows that the average deviation from the actual expenditures to the risk-adjusted budgets was 69 euros in 2001, and reduced to 44 euros after omitting the five smallest insurers. This aspect will be studied further in section 5.

Table 3.2 Prospective risk-adjustment performance

Year *t*	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
MAR (in euros)	34	65	119	98	109	63	74	72	62	97	69
MAR * (in euros)	34	65	80	78	68	32	27	43	25	38	44
RAP	0.58	0.27	0.19	0.23	0.32	0.39	0.45	0.43	0.59	0.46	0.56

*Omitting insurers with population < 50,000. Numbers are expressed in 2001 prices and are rounded off.

[12] Mobility was rather low in the Netherlands. On average, market shares of sickness funds fluctuated only about 0.4% between 1995 and 2000 (Douven and Sahin, 2003). Low mobility also suggests that expenditure differentials between insurers are mainly determined by enrolees who stay put and do not switch from insurer.

Lamers *et al.* (1999) shows that in the Dutch social health insurance sector not even 10% of expenditure variation on an individual level can be explained by risk adjustment, using the demographic risk adjusters age and gender. After including more sophisticated risk adjusters (such as pharmaceutical expenditure groups), the authors show that the percentage increases to about 20%.[13] The fact that expenditure variation on the individual level is larger than on the insurer level can, of course, be explained by the fact that we consider average per capita expenditures of populations instead of individual expenditures, which, according to the law of large numbers, leads to a lower variation.[14]

3.2 Retrospective risk-adjustment performance

As shown in section 2.3, the government diminished the role of retrospective payment schemes during the sample period and influenced thereby not only expenditure differentials between insurers, but also their incentives for efficiency and risk-selection efforts. This section calculates and discusses the following statistics:

$$MAR_t^r = \frac{1}{K_t}\sum_{i=1}^{K_t} | e_{it} - b_{it}^r |, \quad RAP_t^r = 1 - \frac{MAR_t^r}{MART_t}, \quad MARG_t^r = \frac{1}{K_t}\sum_{i=1}^{K_t} | b_{it} - b_{it}^r |$$

Where MAR^r represents the variation left in health care expenditure after insurers obtained their (prospective and) retrospective payments B_{it}^r (where $b_{it}^r = B_{it}^r / N_{it}$). Note that risk-adjusted prospective payments mainly filter out structural variation of expected health care expenditures (which are related to the risk adjusters in table 2.1), while retrospective payments may filter out all types of variation: both structural variation (for example, related to not only imperfect risk adjustment but possibly also to management differences) and random variation (for example, related to unexpected events). RAP^r is now similar to *RAP,* but measures risk-adjustment performance after insurers obtain their prospective and retrospective payments. We included also $MARG^r$ in our analysis. $MARG^r$ measures variation of the retrospective payments that the government transfers at the end of the year to the insurers.

Table 3.3 shows that the remaining expenditure variation (MAR^r) is rather small, but increases slightly due to the policy to increase insurers' financial risk (see fifth row of table). We conclude, however, that retrospective risk-adjustment performance (RAP^r in the fourth row of the table) did not show a clear upward or downward trend after 1995, and varied around 85%, with one outlier in 1997 of 77%.

[13] Note that these percentages correspond to conventional R^2 .

[14] In general, it can also be said that the larger insurance companies become (and thus the fewer of them operating on a market), the lesser expenditure differentials between insurers will be.

Predictably, the sum of the variation of government transfers ($MARG^r$) and the unexplained variation after retrospective risk adjustment (MAR^r) is a good proxy for the unexplained variation after prospective payments (MAR) .

Table 3.3 Retrospective risk-adjustment performance

Year	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
MAR (euros)	34	65	119	98	109	63	74	72	62	97	69
MAR^r (euros)	3	15	13	10	8	16	21	33	25	25	27
$MARG^r$ (euros)	31	52	115	92	103	57	62	59	49	80	53
RAP^r	0.96	0.81	0.91	0.95	0.85	0.85	0.77	0.84	0.84	0.86	0.83
Financial risk	0.00	0.00	0.03	0.03	0.03	0.15	0.27	0.28	0.35	0.36	0.38

4 Prediction of the macro budget

As explained in section two, even if we have a perfect risk-adjustment formula, we still need an adequate prediction of expected health care expenditures for the coming year. What happens when this prediction turns out to be wrong?

Table 4.1 shows two indicators: absolute and percent differences (*pd*). We observe that the macro budget is generally set too low. The largest prediction difference occurred in 1992: the percentage difference was 4.7%, which corresponds to an on-average per capita expenditure difference of 61.2 euros.

Table 4.1 Absolute and percent differences between actual expenditures and the macro budget

Year	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
Differences (per capita, euros)	33.6	61.2	57.2	42.8	– 26.6	21.4	17.1	37.4	1.2	– 3.6	17.8
% differences (*pd*)	2.6%	4.7%	4.4%	3.2%	– 2.1%	1.7%	1.3%	2.9%	0.1%	– 0.3%	1.4%

If the prediction of the macro budget does not equal actual expenditures, then two types of problems may arise. The first is a technical problem, and related to the use of our *RAP* statistics; the second, more fundamental, problem shows that a prediction error may generate distribution errors between insurers. Both problems will be discussed in greater detail below.

4.1 A correction of our *RAP* statistics

The *RAP* statistics that are used in the previous section, which are based on the fact that insurers receive fixed subsidies, fail to take into account that insurers also raise nominal, community-rated, premiums. With regard to nominal premiums, an important observation is that a mismatch between the projected macro budget and actual expenditures is not necessarily a problem. For example, if every insurer would receive for each person the correct individual risk-adjusted subsidy minus a fixed amount of money, then we would not see this as a problem. Each insurer would be able to correct this mistake through raising their nominal premium by the same fixed amount of money, and mutual competition would not be disturbed.[15] Our *RAP* statistic, however, is not invariant for this possibility.

Consider the following example. Suppose that the correct health care profiles between insurers A and B would be as $hcp_A : hcp_B$ = 0.9:1.1. Suppose, furthermore, that the government correctly predicts next year's expected per capita health care expenditures at 1000 euros and sets the macro budget accordingly. Risk adjustment now implies that insurer A receives 0.9*1000=900, and insurer B 1.1*1000=1100 euros per capita. If actual expenditures

[15] It will be somewhat more complicated in practice, since subsidies to insurers have to be adjusted for the fact that children under the age of eighteen do not have to pay out-of-pocket premiums. Furthermore, a lower macro budget implies lower income-related basic premiums and higher nominal premiums (see section 2.1), which may have welfare consequences.

are 900 (1100) euros, then risk adjustment is perfect and our *RAP* statistic would generate a value of one. Now let's assume that the government opts for a lower macro budget and, for example, subsidises insurer A with 700 euros per capita and insurer B with 900 euros. Risk adjustment would be perfect as well, since both insurers can obtain their additional expenditures by raising their nominal premiums by 200 euros. However, if we substitute the values of 700 and 900 euros in our *MAR* statistics, we would obtain a *RAP* statistic totally different from one! The combination of risk-adjusted payments and community-rated premiums implies, thus, that we need a measure for *MAR* and *RAP* that is invariant for possible (fixed per capita) deviations of the macro budget from total actual expenditures. We therefore repair our statistics by adding a fixed amount to b_{it} in MAR_t (to b_t in $MART_t$), such that the adjusted budgets of insurers always add up to actual expenditures. This leads to the following update:[16]

$$MAR_t^f = \frac{1}{K_t}\sum_{i=1}^{K_t} | e_{it} - (b_{it} + (e_t - b_t)) |, \qquad MART_t^f = \frac{1}{K_t}\sum_{i=1}^{K_t} | e_{it} - (b_t + (e_t - b_t)) |,$$

$$RAP_t^f = 1 - \frac{MAR_t^f}{MART_t^f}$$

These statistics capture the effect of community-rated premiums and are invariant to possible prediction errors, since the sum over the adjusted per capita budgets always adds up to total actual expenditures.[17]

Table 4.2 shows the new results together with the *MAR* and *RAP* statistics from the previous section. We observe differences, of course, only if a prediction error occurs. The new RAP^f statistics seem to be more plausible. For example, in 1992 the RAP^f was 60%, which is not much lower than the RAP^f in 1991. Hence, in 1991 and 1992 historical expenditures seemed to be a very good predictor of current health care expenditures. In RAP^f the trend of improving risk-adjustment performance from 1993-2001 becomes even slightly more apparent.

Table 4.2 Prediction differences (*pd*), MAR^f, MAR, RAP and RAP^f

Year	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
% Differences (*pd*)	2.6%	4.7%	4.4%	3.2%	– 2.1%	1.7%	1.3%	2.9%	0.1%	– 0.3%	1.4%
MAR^f (in euros)	26	32	122	92	101	63	73	68	62	96	66
MAR (in euros)	34	65	119	98	109	63	74	72	62	97	69
RAP^f	0.69	0.60	0.18	0.25	0.35	0.41	0.47	0.52	0.59	0.46	0.58
RAP	0.58	0.27	0.19	0.23	0.32	0.39	0.45	0.43	0.59	0.46	0.56

16 In this equation e_t is defined similarly as b_t, namely $e_t = E_t / N_t$ with $E_t = \sum E_{it}$.

17 Note that $\sum N_{it}(b_{(i)t} - (e_t - b_t)) = E_t$.

4.2 A distribution error between insurers

A more fundamental problem of a wrong prediction of the macro budget is the possible distribution error between insurers. A simple example will illustrate the problem. Suppose again that the correct health care profiles between insurers A and B would be as $hcp_A : hcp_B$ = 0.9:1.1. Suppose now that the government wrongly predicts next year's expected per capita health care expenditures: say $b_t = 800$ euros, whereas actual per capita expenditures, e_t, turn out to be 1000 euros. Risk adjustment now implies that insurer A receives 0.9*800=720 euros per capita, whereas his actual expenditures are 0.9*1000=900 euros. Insurer B receives 1.1*800=880 euros, whereas his actual expenditures are 1100 euros. Insurer B has to pay 220 euros extra, whereas insurer A 'only' has to pay 180 euros. Although both insurers can raise their nominal premium, insurer B will incur an additional loss per capita of 40 euros compared to insurer A. Note that if the government predicts next year's expenditures too high, say $b_t = 1200$ euros per capita, then it is just the opposite. In that case, insurer A has a smaller profit per capita of 40 euros compared to insurer B. In the sequel, we call this error the 'distribution error between insurers of a wrong prediction of the macro budget'.

Assume again that the government had correctly predicted the health care profiles but had wrongly predicted the macro budget.[18] In that case, insurers have received a subsidy of b_{it} per capita, whereas they should have received $(E_t / B_t)b_{it} = (e_t / b_t)\, b_{it} = b_{it} + (e_t - b_t)hcp_{it}$ per capita. To calculate what the actual distribution errors for insurers are, we have to compare this subsidy with the adjusted subsidies, as explained in the previous section.[19] The yearly per capita distribution error (de) now becomes: $de_{it} = [\, b_{it} + (e_t - b_t)hcp_{it}\,] - [\, b_{it} + (e_t - b_t)\,] = (hcp_{it} - 1)(e_t - b_t)$. Note that the variation of the distribution errors increases with the variation in the insurer's health care profiles and the size of the prediction error.

Table 4.3 represents for each year *t* the minimum, maximum and median (the average is zero) of de_{it} , together with the percentage prediction error, *pd*, of the macro budget. Table 4.3 gives an impression of the per capita differences between insurers. For example, in 2001 a 1.4% prediction error of the macro budget resulted in differences of per capita payments between insurers ranging from –4.8 euros to 3.2 euros, with a median of 0.4 euros. Table 4.3 shows also that the distribution error between insurers increases with the size of the prediction error of the macro budget.

[18] Of course, health care profiles may not be predicted correctly. For example, one would expect that after risk adjustment the remaining variation in $e_{it} - b_{it}$ would be orthogonal to hcp_{it} . However, for 1998-2001 we still find relatively high positive correlation coefficients of around 0.5.

[19] Consider again the numerical example. Insurer A received 720 euros, but should have received (1000/800)*(0.9*800) = 900 euros. This amount should be compared with 720+200= 920 euros, where 200 euros is the per capita prediction error of the macro budget. Hence, insurer A has received 20 euros per capita too much (and, similarly, insurer B, 20 euros too little).

Table 4.3	Differences between insurers as a result of a distribution error (after prospective payments, per capita, in 2001 prices, euros)										
Year	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
% Differences (*pd*)	2.6%	4.7%	4.4%	3.2%	– 2.1%	1.7%	1.3%	2.9%	0.1%	– 0.3%	1.4%
Minimum	– 3.6	– 7.4	– 13.6	– 6.9	– 16.6	– 3.3	– 6.1	– 11.3	– 0.4	– 0.8	– 4.8
Maximum	8.6	15.3	3.9	29.7	5.6	6.1	7.7	10.6	0.6	1.1	3.2
Median	– 0.7	– 0.5	0.8	– 2.6	1.5	– 0.5	– 0.8	– 0.5	– 0.0	– 0.1	0.4

By calculating Spearman's correlation coefficients of de_{it}, we can observe whether in two successive years the same insurers are hit by a wrong prediction of the macro budget.[20] Table 4.4 shows relatively high (in an absolute sense) correlation coefficients, suggesting that often the same insurers suffer from a wrong prediction of the macro budget. In four cases we observe a negative correlation coefficient, suggesting that the lucky insurers in a given year may be the unlucky ones in the following (or preceding) year. These negative signs correspond exactly with those years in which an over prediction is followed or preceded by an under prediction of the macro budget. As is shown in table 4.3, 1995 and 2000 are the only two years in the sample that feature an over prediction of the macro budget.

Table 4.4	Spearman's correlation coefficients of the insurers' distribution errors									
Year to year	1991/92	1992/93	1993/94	1994/95	1995/96	1996/97	1997/98	1998/99	1999/00	2000/01
Correlation coefficient	0.78	0.41	0.54	– 0.52	– 0.80	0.91	0.87	0.92	– 0.60	– 0.79

Lacking precise information about the various retrospective compensation schemes, we cannot obtain similar outcomes for per capita expenditure differences after retrospective payments.[21]

4.3 Alternative possibilities of handling the macro budget

Here, we propose two possible solutions to handle a wrong prediction of the macro budget. The first solution is rather straightforward, the second more speculative.

One possible and rather simple solution would be to calculate the prospective risk-adjusted subsidies again retrospectively (thus, after the actual expenditures become known), and then correct for these differences. This procedure entails a redistribution of money between insurers and does not require additional government funding.[22] Technically, this correction works as follows. Assume ex-post it turns out that the macro budget is incorrect ($e_t \neq b_t$). This implies

[20] This statistic is calculated by first replacing the mean observations by their ranks. Spearman's correlation coefficient is obtained through computing the ordinary correlation coefficient of two rank-vectors of two successive years.

[21] We also computed RAP^f and MAR^f statistics including $b_{it} + (e_t - b_t)hcp_{it}$, instead of $b_{it} + (e_t - b_t)$, but this did not yield a significant improvement. This result may be due to the fact that the prediction errors are relatively small and health care profiles are close to one. Also, random errors in the data may distort the picture.

[22] Additionally, since new information is available, one could also update the weights and risk adjusters. See Newhouse *et al.* (1997) for more information on this subject.

that each insurer should receive an additional payment per capita of $hcp_{it}(e_t - b_t)$. Since this would require additional government funding, we subtract for each enrolee a fixed amount. The final payments for insurer *i* now become $(hcp_{it} - 1)(e_t - b_t)$. This formula shows that if the macro budget is set too low $e_t > b_t$, and if insurer *i* has a healthier population than average, $hcp_{it} < 1$, then this insurer has to transfer money to an insurer with a less healthy population than average. Since the payments are neutral with respect to the government budget, this is simply a form of retrospective equalisation. In the example in section 4.2, it turned out that, ex-post, insurer B incurred an additional loss per capita of 40 euros compared to insurer A. To correct for the wrong prediction of the macro budget, ex-post, insurer A, with a healthier population, will have to pay 20 euros per capita to insurer B.[23] The problem of this solution may be one of practicability. In practice, it takes a year (or sometimes two) before all the necessary data become available; until that time insurers face uncertainty about whether they have to receive or pay money.

A second solution transfers the responsibility of predictions for health care expenditures from the government to the insurers. This solution entails that the government *ex-ante* determines b_t , which does not need to be related to total expected health care expenditures for the coming year, and determines the health care profiles hcp_{it} . Insurers predict now total expected expenditures, say e_t^e . The per capita payments equal $hcp_{it}e_t^e - (e_t^e - b_t)$ for each insurer *i*. Consider again the example in section 4.2. The government sets the macro budget at 800 euros per capita. Assume that the insurers predict (correctly) expenditures at 1000 euros per capita. Subsidies are now distributed as follows. Insurer A receives 0.9*1000-(1000-800)=700 euros per capita, and insurer B: 1.1*1000-(1000-800)=900 per capita. By subtracting a fixed amount per capita, insurers obtain a similar per capita loss, which insurers can retrieve by setting their nominal premiums 200 euros higher. The main point of this procedure is that it levels the competitive playing field for insurers. Two interesting elements in this procedure are as follows:

a) The government can determine its own budget independent of precise expenditure developments, and is not responsible for wrong expenditure predictions.

b) Incentives by insurers to game the system can be diminished by a coordinator. This argument is explained by the fact that insurers with a more healthy population will have an incentive to underestimate expected expenditures, e_t^e, since this will increase their share of the pie. However, insurers with a less healthy population will for similar reasons have an incentive to overestimate health care expenditures e_t^e. In a competitive environment, it may not be easy to design a non-cooperative game that would yield a truthful prediction of total health care expenditures as outcome. This requires a coordinator. A practical implementation could be to transfer the responsibility of predicting expected expenditures to an umbrella organisation of *all* insurers. Another option would be to transfer the responsibility to a third institution that has access to all relevant information, and is able to generate better predictions.

[23] In this example we assume that both insurers have equal population size.

5 Consequences for competition

Although in the previous sections we presented averages over sickness funds for each year, no information was provided with regard to individual health care expenditure patterns of insurers. Risk adjustment is important, from a competing insurer's perspective, but even more vital may be the question of how risk adjustment affects different competitors' expenditure patterns. For example, if failures of the risk-adjustment system at the individual level through pooling almost cancel out at the insurer level, then risk adjustment may be called inadequate, but at the insurer level such an outcome does not necessarily imply distorted competition between insurers.

5.1 Individual expenditure structures of insurers

A preliminary overview of individual expenditure structures of insurers can be obtained with the following characterisations:

$$loss_{it} = e_{it} - b_{it}\,, \quad d_{it} = \left| \, loss_{it} - \overline{loss}_i \, \right|, \quad loss^r_{it} = e_{it} - b^r_{it}\,, \quad d^r_{it} = \left| \, loss^r_{it} - \overline{loss}^r_i \, \right| .$$

Here, $loss_{it}$ (respectively $loss^r_{it}$) represents per capita health care expenditure losses, after prospective (and retrospective) subsidies, of sickness fund *i* at year *t*. Both d_{it} and d^r_{it} are distant measures for individual variation. A graphic presentation can be obtained from figure 5.1 which presents for each insurer the interval $[\,\overline{loss}_i - \overline{d}_i, \overline{loss}_i + \overline{d}_i\,]$, where $\overline{loss}_i$ and $\overline{d}_i$ represents means over *t* of $loss_{it}$ and d_{it}. Similarly, figure 5.2 represents the intervals $[\,\overline{loss}^r_i - \overline{d}^r_i, \overline{loss}^r_i + \overline{d}^r_i\,]$. In figure 5.1 the insurers are ranked according to the mean $\overline{loss}_i$, where insurer 1 has the lowest mean, etc. For reasons of comparability we computed the mean only over the years 1997-2001, since for these years the overall financial risk of insurers became substantial and varied only slightly over the years (from 27% in 1997 to 38% in 2001; see figure 2.2), and also the *RAP*s did not vary strongly. To obtain at least four observations for each insurer we plotted only those insurers that were in the market in the year 2001.

Both figures yield some interesting observations. Figure 5.1 clearly shows individual differences between insurers. Some insurers (the lower numbers) make (on average) large profits, and some (the higher numbers) make losses. Individual differences can be large; insurer 1 gains on average about 330 euros per capita, whereas sickness fund 24 loses on average about 80 euros per capita—a difference of 410 euros, which is about 32% of 1297 euros, the mean per capita health care expenditure of the total population. Variation can fluctuate strongly, although it seems to be smaller for insurers with a higher number.

Figure 5.1 **Per capita loss and its variation after prospective payments for 24 insurers (1997-2001).**

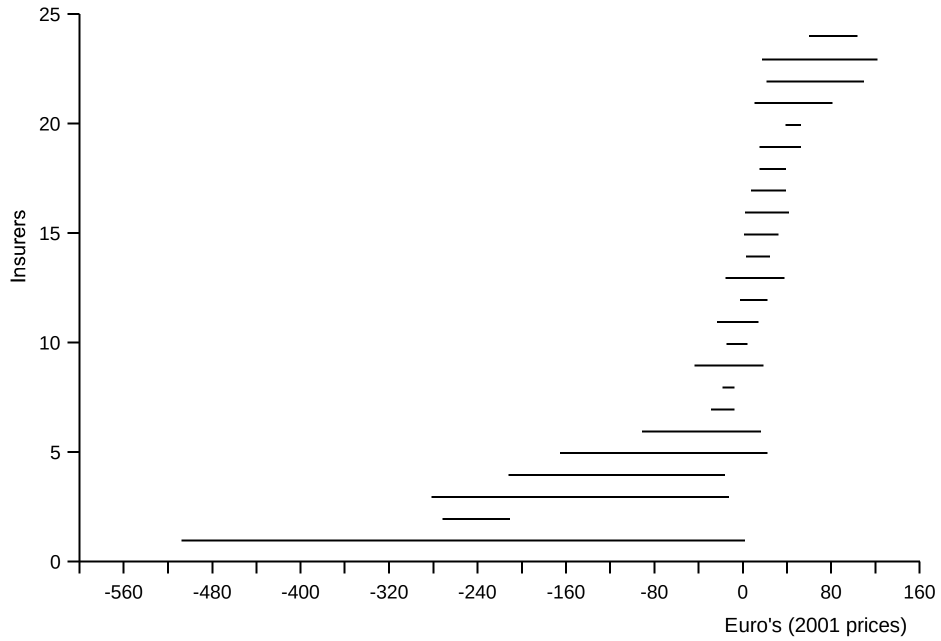


Figure 5.2 **Per capita loss and its variation for 24 insurers after retrospective payments (1997-2001)**

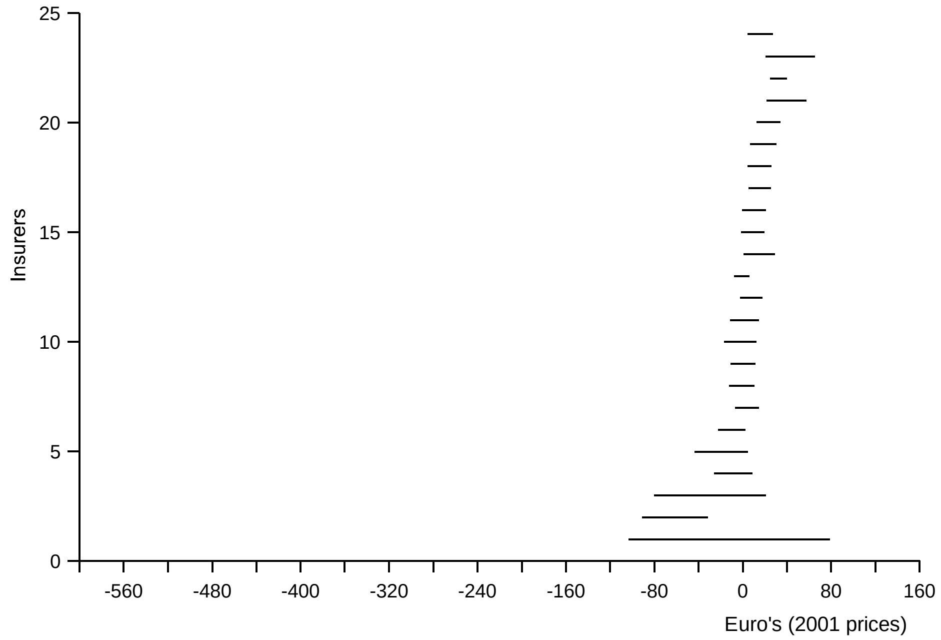


Figure 5.2 ranks the insurers in the same order as in figure 5.1, after final (retrospective) expenditures have emerged. As expected, retrospective payments decrease the mean and the variation of health care expenditures. Insurer 2 receives on average the most profits (about 60 euros per capita), and the largest losses are for insurer 23 (about 40 euros per capita).

The ranking of the means in figure 5.1 is strongly correlated with ranking of the means in figure 5.2. We computed a Spearman's (or the rank) correlation coefficient of 0.95. Hence, retrospective payments seem to have only a limited impact on the ranking of insurers.

An interesting statistic in a competitive environment is the average distance to the lowest mean expenditure insurer over the years 1997-2001. Table 5.1 shows that the lowest expenditure insurer is on average 244 euros per capita better off after prospective payments than its competitors. This number changes only slightly if the five highest expenditure insurers are omitted. Omitting the five lowest expenditure insurers from the computation, the sixth lowest expenditure insurer is "only" 61 euros per capita better off. Thus, if the insurers are sorted according to their means $\overline{loss}_i$ in increasing order (as is done in figure 5.1), then the variation between the low-expenditure insurers is much higher than between the high-expenditure insurers. Note the striking result in the last two rows of table 5.1. The same number of 61 in the first column suggests that the five insurers with the lowest mean expenditures after prospective payments are also the five with the smallest population. We will discuss this finding further in section 5.3. Smaller numbers show up in table 5.1 after insurers received their retrospective payments. A comparison of the average distance numbers after prospective expenditures with final (retrospective) expenditures shows that the average distance numbers decrease to about 30%, which corresponds again with the 33% of the mean overall financial risk during 1997-2002.

To break even with the lowest expenditure insurer, all else equal, other insurers must raise out-of-pocket premiums, on average, even more than 69 euros per year (which is about 5% of the mean per capita health care expenditures of the total population).[24]

Table 5.1 Average per capita distance to the lowest expenditure insurer (in euros, 1997-2001)[a]

Group of insurers considered	After prospective payments	After retrospective payments
(1) All 24 insurers in market at 2001	244	69
As (1), omitting five highest expenditure insurers	225	62
As (1), omitting five lowest expenditure insurers	61	20
As (1), omitting the five insurers with a population < 50,000	61	23

[a] First, we computed the mean $\overline{loss}_i$, respectively $\overline{loss}_i^r$. Next, we sorted out the insurer with the lowest mean and calculated the average distance to this mean of the other insurers.

[24] The reason is that not every person needs to pay an out-of-pocket premium. Children under the age of eighteen do not need to pay these premiums.

5.2 Luck or structural differences?

Section 3.1 discussed various reasons for health care expenditure differences between insurers. One important and unanswered question is to what extent differences are related to unexpected events (luck) or to structural differences. The latter may be related to structural factors such as population differences that are not compensated by the risk-adjustment scheme, or structural efficiency or quality differences between insurers. A rough estimate can be obtained by calculating Spearman's correlation coefficients (see footnote 19) of $loss_{it}$ and $loss_{it}^{r}$.

Table 5.2 Prospective and retrospective Spearman's correlation coefficients of $loss_{it}$ and $loss_{it}^{r}$.

Year to year	1991/92	1992/93	1993/94	1994/95	1995/96	1996/97	1997/98	1998/99	1999/00	2000/01
Correlation (prospective)	– 0.12	0.21	0.80	0.36	0.31	0.21	0.52	0.44	0.65	0.89
Correlation (retrospective)	– 0.10	0.21	0.75	0.35	0.47	0.88	0.49	0.42	0.78	0.89

In table 5.2, the correlation coefficients are obtained by correlating the yearly rankings of those insurers that were both years on the market. Remarkably, Spearman's correlation coefficients for the first two figures are rather low, suggesting few structural differences between insurers. A possible explanation for the findings for 1991/92 and 1992/93 is the use of the risk adjuster 'historical expenditures' at the insurer level. A risk adjuster such as 'historical expenditures' compensates a high expenditure (low-ranked) insurer in the previous year with higher subsidies (and thus a higher ranking) in the current year. This explains also one of the reasons why the risk adjuster 'historical expenditures' was abandoned in the year 2003: a system of premium subsidies based on historical expenditures pays insurers without regard to the appropriateness of care (McClure, 1984). Disregarding the two probationary years, we observe a wide spectrum of positive correlation coefficients. A comparison of the prospective and retrospective correlation coefficients yields similar results, except for 1996/97.[25] The average prospective and retrospective Spearman correlation coefficient over 1993/94-2000/01 is about 0.6, which suggests that differences between insurers are related more to structural than to random differences. From these results it cannot be seen to what extent remaining structural differences are connected with differences in efficiency between insurers.

5.3 Do small insurers have lower per capita expenditures?

In table 5.1 of section 5.1 we suggested that the smallest five insurers were also those with the lowest expenditures during 1997-2001. This suggests that the market share of an insurer matters, and that small insurers have lower per capita expenditures. Although this observation may be correct, we will show in this section that a better explanatory variable for low per capita

[25] The reasons for these findings in 1996/97 are unclear.

expenditures is whether or not an insurer entered the market. This result ensues from estimating the following two equations for $loss_{it}$ and $loss^r_{it}$:

$$loss^{(r)}_{it} = \beta_1 \; entrant_{it} + \beta_2 \, exit_{it} + \beta_3 \, market\ share_{it} + \alpha_i + \mu_t + \varepsilon_{it}$$

The explanatory variables are defined as follows:

$entrant_{it}$: 1 for all t if sickness fund i entered the market during 1991-2001, 0 otherwise.

$exit_{it}$: 1 for all t if sickness fund i left the market during 1991-2001, 0 otherwise.

$market\ share_{it}$: market share (in percentage points) of sickness fund i in year t.

α_i: dummies capturing unobserved fund effects (in case of pooled OLS, $\alpha_i = \alpha$).

μ_t: dummies capturing annual changes in per capita health care expenditures.

ε_{it}: errors that are assumed to be independently identically distributed.

Table 5.3 shows eight estimation results: four for the dependent variable per capita loss, $loss_{it}$, and four after retrospective risk adjustment $loss^r_{it}$. For each dependent variable we considered two different time periods, the whole period 1991-2001 and 1997-2001. The latter time period was chosen for reasons similar to those delineated in section 5.1. For each time period two types of estimators are presented: a pooled OLS estimator and a fixed-effect estimator.[26]

Table 5.3 **Estimation results of pooled OLS and fixed-effect (FE) equation**[a]

	Dependent variable $loss_{it}$				Dependent variable $loss^r_{it}$			
	Time period 1991-2001		Time period 1997-2001		Time period 1991-2001		Time period 1997-2001	
Variables	OLS	FE	OLS	FE	OLS	FE	OLS	FE
$Entrant_{it}$	– 200 (22)	-	– 178 (28)	-	−34 (5)	-	−45 (9)	-
$Exit_{it}$	28 (18)	-	68 (32)	-	10 (4)	-	18 (10)	-
Market $share_{it}$	– 1.1 (2.5)	0.5 (5.8)	– 1.8 (3.5)	– 6.3 (15.5)	– 0.6 (0.6)	– 0.5 (0.5)	– 1.2 (1.1)	– 0.8 (5.1)
R^2	0.35	0.59	0.31	0.63	0.34	0.50	0.33	0.57

[a] All results follow from incomplete panels, since some insurers finished business, or merged during the observation period. The number of observations is 317 for the period 1991-2001, and 147 for the period 1997-2001. The standard errors appear between parentheses.

Inclusion of the dummies α_i in the fixed-effect estimator eliminates any time-invariant variables from the model. The effect of the variables $entrant_{it}$ and $exit_{it}$ is thus negated. The effect of the variable $market\ share_{it}$ is included, however, in all estimation results.

The pooled OLS regressions confirm that the ten sickness funds that entered the market during 1991-2001 had less per capita expenditures than those that were already on the market in 1991. The results suggest that the per capita health care expenditures of newcomers were, on average, 200 euros less during 1991-2001, and 178 euros less during 1997-2001. The result is strongly significant and holds also (albeit to a smaller extent) after retrospective payments.

[26] The outcomes of the random-effect model are comparable to the pooled OLS model.

The pooled OLS regressions show positive estimates of the variable $exit_{it}$, and are twice significant at the 5% level. This suggests that the 16 sickness funds that left the market and often merged with other funds during 1991-2001 had higher expenditures than average.

In the pooled OLS- and fixed-effect estimations, the variable market share is often negative but highly insignificant. Inclusion of the unobserved funds effects α_i in the fixed-effect model, and inclusion of the variables $entrant_{it}$ and $exit_{it}$ in the pooled OLS, diminish the effect of market share.[27] This shows also the problem of our estimators. In the fixed-effect model the unobserved funds effects α_i , and in the pooled OLS especially the explanatory variable $entrant_{it}$, are likely to be strongly correlated with the variable $market\ share_{it}$. The reason is that new insurance companies entering the market start often with a small population, which is especially the case in our sample. To control for this aspect, we balance the sample and consider only the twenty insurers that were on the market during the entire period. Performing fixed and random-effect estimators for several time periods did yield sometimes positive and sometimes negative estimators for $market\ share_{it}$, but all these estimators proved to be highly insignificant. To conclude, for the period examined we found no evidence supporting the argument that economies of scale play a role and lead, for example, to more efficiency and thus to lower per capita health care expenditures.

Why entrants on the insurance market have on average lower health care expenditures remains unclear. New entrants on the social security market since 1991 have almost all been established by private health insurers (Schut and Hassink, 2003). It may be the case that private health insurers are more efficient, or keener to perform some kind of risk selection, compared to social health insurers. Private insurers, for example, may be particularly attractive for employers with favourable health care expenditures. [28] Some newcomers may have used other selection strategies directed at favourable groups. These will be effective only if a health care insurer can beat the risk-adjustment system. For example, selection may have played a role in 2000 when the Dutch government brought some 375,000 lower-income self-employed persons (and dependents) under the mandatory social health insurance scheme. Ex-post, the data show that comparatively more self-employed opted for sickness funds that entered the market during 1991-2001. Ex-post, the data show also that self-employed persons turned out to be more profitable than other persons in the same risk group. Another argument that may be relevant is that sickness funds that enter the market attract mobile persons. These persons may be more profitable than others in the same risk group[29].

[27] Pooled OLS without the variables *entrant* and *exit* yields a significant negative effect for the variable *market share*.

[28] Lamers *et al.* (2003a) cite an example of the establishment of a new sickness fund by a large employer primarily for its own employees. The data showed that this sickness fund had substantially lower health care expenditures than others. Therefore, a new risk adjuster 'yes/no being employed' was added to the risk adjustment system in 1999.

[29] The literature suggests that mobile persons are healthier (and more profitable) than their non-mobile counterparts (see, for example, Strombom *et al*., 2002).

6 Conclusions

This paper presents a risk-adjustment performance indicator at the insurer level that compares the variation of insurers' per capita health care expenditures both with and without risk adjustment. This indicator shows that prospective risk adjustment improved from 1993-2001. While risk adjustment could explain initially about 20% of the variation of health care expenditures, improvements in the risk-adjustment system by the inclusion of additional risk adjusters caused a rise in this figure to about 55% in the year 2001. Retrospective risk-adjustment performance does not show a clear upward or downward trend and varied around 85%. This implies that the increase in insurers' financial risk coincided with the improvement of the risk-adjustment system. This still leaves unexplained, however, about 45% prospectively, and about 15% retrospectively, of the variations in insurers mean per capita health care expenditures.

In monetary terms (2001 prices), the remaining variations can be substantial. Considering for a moment only prospective subsidies, we find during 1997-2001 that the average insurer spent about 244 euros per capita per year more than the lowest expenditure insurer (or 19% of 1300 euros, which represents the mean per capita health care expenditures of the total population). Disregarding the five lowest expenditure insurers, this figure declines drastically to about 61 euros per capita. The same numbers, of course, decline after the occurrence of retrospective payments. Then we find 69 euros and, respectively, 20 euros.

The first question to answer here seems to be whether these differences between insurers are structural or random. The latter case presents no problem, since lucky outcomes for an insurer in some years will be compensated by unlucky outcomes in other years. We find that differences between insurers are for a greater part structural than random, which suggests that some insurers have structurally lower health care expenditures.

The next question to answer is what could possibly explain these structural differences? If they could be explained by efficiency arguments alone, then the risk-adjustment system would be working perfectly, since efficient insurers could keep their gains. Unfortunately, our dataset is too limited to lead to an empirical answer to the question of the efficiency argument. Harnessing the observations of previous researchers, however, we may conclude that during our sample period it not very likely that all (or even most) of the structural variation can be explained by efficiency alone. First is the observation that during our sample period sickness funds were only starting to employ managed-care activities; as a result, differences in quality or efficiency in purchasing or organising medical care are probably low. Second, risk adjustment in the Netherlands is still considered to be imperfect. The cited paper in the introduction, Van de Ven *et al.*(2004), shows that insurers still make high predicted losses in year *t* of the 10 percent of people with highest expenditures in year *t*-1. This result implies also that the need for retrospective payment schemes, such as retrospective equalisation for high risks, will remain.

We have addressed one aspect of imperfect risk adjustment, which is related to the practical implementation of risk adjustment: the prediction of the so-called ‘macro budget’. In the Netherlands this macro budget is determined by the government and is used as an ex-ante projection of total expected health care expenditure of the coming year. We show theoretically that if this projection turns out to be too low ex-post, then insurers with a population of relatively good health risks are favoured. If this projection turns out to be too high, then it is the other way around, and insurers with a population of relatively bad health risks will be favoured. During the sample period the ex-ante projection turned out to be structurally too low. In 2001, for example, a 1.4% prediction error of the macro budget resulted in differences of mean per capita prospective payments between insurers ranging from –4.8 euros to 3.2 euros per capita. Also important: the payment errors to insurers increase as the prediction error of the macro budget gets larger. A simple solution for this problem would be to calculate the prospective risk-adjusted subsidies again retrospectively, and correct for these differences.

We also examined the possibility that efficiency in purchasing medical care may have played a role by testing whether economies of scale are important in our sample period. Estimations of panel data models showed that new small entrants on the insurance market have on average (much) lower health care expenditures. The precise reason why entrants have lower health care expenditures remains unclear. One influence might have been that new entrants on the social security market since 1991 were almost all private health insurers. After disregarding the entrants, we could find no evidence that economies of scale played any role during 1991-2001.

As mentioned at the beginning of this concluding section, two important occurrences since 2001 may have changed the picture that follows from this paper: new health-related risk adjusters have been incorporated in the risk-adjustment system, and insurers’ financial risks have increased. It seems likely that prospective risk adjustment has improved further. Even larger changes will occur in 2006, after the introduction of a new health care system. The main idea is that more competition between insurers will lead to greater efficiency. If risk adjustment at some point in the future can be considered (practically speaking) as almost perfect, then management or efficiency differences should be the only structural factor left explaining expenditure differentials between insurers.

References

Altman, D., D.M. Cutler and R.J. Zeckhauser, Adverse selection and adverse retention, *American Economic Review* 88(2), pp. 122-126, 1998.

Antioch, K. and M. Walsh, Funding issues for Victorian hospitals: the risk-adjusted vision beyond casemix funding, *Australian Health Review* 23 (3), pp 145-153, 2000.

CPB, Zorg voor concurrentie, een analyse van het nieuwe zorgstelsel, CPB Document 28, The Hague, 2003.

CVZ, Evaluatie budgettering verstrekkingen ZFW 2000, Publicatienummer 87, Diemen, 2001.

Douven, R. and H. Sahin, Deliberate consumer choice in health insurance, *CPB Report* 2003/1, pp 48-51, 2003.

Douven, R. and E. Westerhout, Reimbursement systems in Dutch sickness funds, *CPB Report* 2000/3, pp 51-55, 2000.

Ellis, R.P. and A. Ash, Refinements to the diagnostic expenditure group model, *Inquiry* 32, pp 418-429, 1995.

Ettner, S.L., R.G. Frank and T.G. McGuire, Risk adjustment of mental health and substance abuse payments, *Inquiry* 35, pp 223-239, 1999.

Lamers, L.M.,R.C.J.A.Van Vliet and W.P.M.M. Van der Ven, Farmacie Kosten Groepen: een verdeelkenmerk voor normuitkeringen gebaseerd op medicijngebruik in het verleden, iBMG, Erasmus University Rotterdam, 1999.

Lamers, L.,R.C.J.A. Van Vliet and W.P.M.M.Van der Ven, Risk adjustment premium subsidies and risk sharing: key elements of the competitive sickness fund market in the Netherlands, *Health Policy* 65 (1), 49-65, 2003a.

McClure, W., On the research of risk adjusted capitation rates, *Inquiry* 21, pp 205-213. 1984.

Newhouse J.P., Reimbursing health plans and health providers: efficiency in production versus selection, *Journal of Economic Literature* 34, pp 1236-1263, 1996.

Newhouse, J.P., M.Beeuwkes Buntin and J.D. Chapman, Risk adjustment and medicare: Taking a closer look, *Health Affairs* 16(3), pp 26-43, 1997.

Pope, G.C., R.P. Ellis, A.S. Ash *et al.*, Principal inpatient Diagnostic Expenditure Group model for Medicare risk adjustment, *Health Care Financing Review* 21(3), pp 93-118, 2000.

Rosenkranz, S.L. and H.S. Luft, Expenditure models for prospective risk adjustment: Choosing the measure appropriate for the problem, *Medical Care and Research Review* 54, pp 123-143, 1997.

Schut, F.T. and W.H.J. Hassink, Managed competition and consumer price sensitivity in social health insurance, *Journal of Health Economics* 21, pp 1009-1029, 2002.

Shen, Y. and R.P. Ellis, How profitable is risk adjustment, Discussion Paper, Boston University, http//econ.bu.edu/Ellis/Papers/shen2.pdf, 2001.

Strombom, B.A., T.C. Buchmueller and P.J. Feldstein, Switching costs, price sensitivity and health plan choice, *Journal of Health Economics* 21, pp 89-116, 2002.

Van Barneveld, E.M., Risk sharing as a supplement to imperfect capitation in health insurance: A tradeoff between selection and efficiency, Ph.D. Thesis, iBMG, Erasmus University Rotterdam, the Netherlands, 2000.

Van der Ven, W.P.M.M. *et al*. Risk adjustment and risk selection on the sickness fund insurance market in five European countries, *Health Policy* 65 (1), 75-98, 2002.

Van der Ven, W.P.M.M., R.C.J.A. van Vliet and L.M. Lamers, Health-adjusted premium subsidies in the Netherlands, *Health Affairs* 23(3), pp 45-55, 2004.

Van der Ven, W.P.M.M. and R.P. Ellis, Risk adjustment in competitive health plan markets. In Cuyler A.J. and J.P. Newhouse (eds.), *Handbook of Health Economics*, Vol 1A, Amsterdam, Elsevier, pp 755-845, 2000.

VWS, A question of demand, Ministry of Welfare, Health Care and Sports, The Hague, available at www.minvws.nl, 2002.

Appendix

This appendix explains in more detail the retrospective compensation schemes introduced by the government during the sample period. Table A.1 shows the percentages as used for the retrospective equalisation schemes. In 1996, the government divided total expenditures into three new expenditure categories. Hospital expenditures were divided into three categories: one cannot be influenced by sickness funds (fixed hospital expenditures), the second (variable hospital expenditures) and third (other medical expenses, which includes prescription drugs, outpatient medical care and medical specialist care) are more impressionable by sickness funds. In 1999, the third category was divided, creating a new category: medical specialist care. Only for the latter category did retrospective equalisation take place in 2001.

Table A.1 Retrospective equalisation schemes

Year	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
Total acceptable expenditures	90%	95%	90%	90%	90%						
Hospital outlays (variable expenditures)						60%	30%	30%	30%		
Other medical expenses						60%	30%	30%	30%		
Medical specialist care											50%

High-risk equalisation was introduced in 1997. Table A.2 shows the thresholds and percentages for the high-risk equalisation schemes. 90% of the expenditures of all patients with health care expenditures above the threshold were transferred to a pool. The money in the pool was equally shared across sickness funds. The threshold has increased, which means that sickness funds bear more risk on expensive patients.

Table A.2 Retrospective equalisation schemes for high risks

	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
Threshold (euros)							2040	2040	3400	4540	4540
Percentage equalisation							90%	90%	90%	90%	90%

After equalisation, the government applies retrospective compensation.[30] Table A.3 shows the percentages for the retrospective compensation scheme. The table shows the change in policy during the period examined. In 1991, 90% of the expenditures were retrospectively reimbursed; this percentage declined to 75% during 1992-1995.[31] In 1996, the government introduced

[30] A complicating factor is, furthermore, that budgets are adjusted for several other factors such as the revenues that sickness funds obtain from third parties, expenditures that sickness funds incur for undertaking new health care activities, and the exact number of enrolees (CVZ, 2001).
[31] In 1991 and 1992, the government decided afterwards to reimburse sickness funds for the full 100%.

various expenditure categories with different declining compensation percentages. Note that the category ‘fixed hospital expenditures’ is not impressionable by sickness funds and therefore receives a retrospective compensation percentage of almost 100%.

Table A.3 Retrospective compensation schemes

	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001
Total acceptable expenditures	90%	75%	75%	75%	75%						
Hospital outlays (variable expenditures)						50%	25%	25%	25%	25%	25%
Hospital outlays (fixed expenditures)						95%	95%	95%	95%	95%	95%
Other medical expenses						50%	25%	15%			
Medical specialist care									95%	95%	40%