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Stefan Greß, Marcus Tamm, Harald Tauchmann and Jürgen Wasem*

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Abstract
In 1996 free choice of health insurers has been introduced in the German social health insurance scheme. Competition between insurers was supposed to increase efficiency. A crucial precondition for effective competition among health insurers is that consumers search for lower-priced health insurers. We test this hypothesis by estimating the price elasticities of insurers' market shares. We use unique panel data and specify a dynamic panel model to explain changes in market shares. Estimation results suggest that short-run price elasticities are smaller than previously found by other studies. In the long-run, however, estimation results suggest substantial price effects.

JEL-Classification: I18, C33

Keywords: Competition, generalized method of moments, health insurance

* Stefan Greß, University Duisburg-Essen; Marcus Tamm, RWI Essen and Ruhr-University Bochum; Harald Tauchmann, RWI Essen; Jürgen Wasem, University Duisburg-Essen. The authors are grateful to Boris Augurzky, Jörg Breitung, Christoph M. Schmidt, and to participants at the RAN (Risk Adjustment Network)-meeting for helpful comments and to Nadine Heck and Muhamed Kudic for research assistance. They also would like to thank Dostal & Partner (www.dostal-partner.de) for providing most of the data. All correspondence to Marcus Tamm, Rheinisch-Westfälisches Institut für Wirtschaftsforschung (RWI Essen), Hohenzollernstr. 1–3, 45128 Essen, Germany, Fax: +49-201-8149236, Email: tamm@rwi-essen.de.

"The exit option is widely held to be uniquely powerful: by inflicting revenue losses on delinquent management, exit is expected to induce that 'wonderful concentration of the mind' akin to the one Samuel Johnson attributed to the prospect of being hanged" (Hirschman 1970).

1. Introduction

In the mid 1990s Germany and several other European countries such as the Netherlands and Switzerland undertook major reforms of their social health insurance systems. In Germany consumers were given free choice of social health insurers starting in 1996. The concept of managed competition can be seen as blue print for reform in these countries (Enthoven 1988, 1993).[1] Managed competition implies that risk-bearing health insurers compete with each other based on price and quality. Furthermore, this concept presumes that insurers induce non-efficient providers to work more efficiently and provide good quality. Otherwise, they are not contracted and drop out of the market. Finally managed competition assumes that consumers have free choice between insurers and exercise their right to choose – at least to some degree.

This paper tests the last assumptions – for competition between health insurers to be effective, (i) consumers must have some choice between them. In Hirschman's terminology: consumers must have the exit option (Hirschman 1970; Schlesinger 1995). At the same time, a crucial precondition for the success of reforms providing a higher degree of choice is that (ii) consumers are inclined to search for lower-priced health insurers or for insurers with higher quality. Our analysis focuses on the estimation of the price sensitivity of consumers.

In the German social health insurance market risk-bearing health insurers compete for enrollees who are allowed to switch between companies on a regular basis. In contrast to other countries with open enrolment such as Switzerland or the Netherlands the enrolment period is not restricted to the end of the year. Clients may switch at the end of each month if they stayed at their insurance company for at least 18 months or if the company has raised the premium.

German social health insurers calculate income-dependent premiums that do not correspond to individual risks. Instead, they set contribution rates and clients' payments are equal to salaries times the contribution rate, up to an income ceiling. Half of the premium is paid by employers; the other half is paid by employees. Premiums are paid directly to the insurance companies. In principle, it is mandatory for all employed workers to acquire social health insurance. Yet, when the salary exceeds a threshold[2] the person can choose whether

[1] At least in the Netherlands, health care reforms were based explicitly on Enthoven's concept.

to remain in the social health insurance system or to opt out and acquire private health insurance. In this paper we do not address the choice between social and private health insurance but focus on the choice within the social health insurance system. I.e. our analysis does not explain the total number of clients in the social health insurance system, but is based on the market shares of insurance companies within this system.

The number of competing health insurance companies in Germany is quite high.[3] Overall, there were 282 companies in 2004. Some companies operate on a national level, whereas the majority of health insurers operate in some regions only. Furthermore, access to several company-based health insurers (BKK) and guild-based health insurers (IKK) is restricted.[4] Depending on the respective region (mostly states or *Bundesländer*), consumers can choose between 50 and 100 health insurance companies. Thus, one requirement for managed competition to work is clearly fulfilled: consumers have free choice between many insurers.

By legal restriction, competition among insurance companies is almost exclusively based on price (contribution rate), since more than 95 percent of the benefits package is standardized across companies. Only for some services, such as alternative medicine, it is up to the insurance company to decide whether or not and to what extent to include these services in the benefits package (Greß et al. 2005). Moreover, as a matter of principle, health insurance companies are obliged to contract collectively with all licensed health care providers. Legal opportunities to contract selectively are very limited. Hence, the quality of insurance is basically identical among companies and thus our analysis focuses on price differences, exclusively.

Consumers' price sensitivity has been analyzed before, for Germany as well as for other countries. Schwarze/Andersen (2001) use German micro-data and provide a descriptive analysis of the socioeconomic background of switchers and non-switchers. Another study found that member losses and member gains of health insurers are closely correlated with contribution rates (Greß et al. 2002). Moreover, price elasticities of market shares between 1996 and 2001 were found to be quite high (–2.90) and increasing over time (Schut et al. 2003). This study by Schut et al. (2003) is quite closely related to ours. However, they use data on health insurers aggregated by type of insurer and esti-

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2 This threshold generally is somewhat higher than the income ceiling used for the calculation of the individual premiums, e.g. in 2004 persons with a monthly income of 3 862.50 € were allowed to choose between private and social health insurance and the income ceiling was equal to 3 487.50 € per month.

3 In the remainder the term health insurer only refers to social health insurers.

4 For a more detailed description of the origins of the quite peculiar German health insurance market see for example Greß et al. (2004).

mate a static panel data model only. Studies on price elasticities in other social health insurance markets in Switzerland and in the Netherlands found much smaller price elasticities (Beck 2004; Schut, Hassink 2002; Schut et al. 2003).[5]

In our paper we apply a range of econometric techniques not applied to the research question before, in particular, we analyze dynamic models. Modeling a dynamic process seems to be more appropriate than a static one because only a small number of consumers is inclined to decide on their health insurer each period and therefore the market is likely to display persistence. We use a new data set which covers a quite recent period and is based on a complete panel of individual health insurers. So far, a complete panel has been available on an aggregated level only. Therefore our panel data on the level of individual health insurers is unique. Our findings support the notion that consumers display a distinct sensitivity to differences in contribution rates.

The rest of the paper is organized as follows. Section 2 describes the data set that has been collected for analyzing price sensitivity of consumers and section 3 specifies the econometric models used. Our estimation results are presented in section 4. Finally section 5 concludes.

2. Data

This study is based on a complete panel of all individual health insurers that were active in the German social health insurance market between January 2001 and April 2004.[6] For each health insurer the panel includes the contribution rate and the number of enrollees in each of seven waves. These seven waves are unequally spaced. The panel also contains information about mergers between health insurers. In our analysis, merged companies are considered to be new entrants into the market, i.e. we use an unbalanced panel.

Because health insurers are not obliged to publish information on the number of enrollees, data had to be collected by Dostal & Partner, a commercial market research company which is specialized in analyzing the German health insurance market. The data have been validated by comparing them to information which was provided by several branch organizations of health insurers and by the Federal Ministry of Health. It is the first time that a complete panel of individual health insurers has been constructed. As a consequence, our panel covers the complete German social health insurance market. Other studies on price elasticities in Germany were only able to base their analysis

5 There have been several studies on health plan choice in the US group health insurance market. Estimated out-of-pocket elasticities range from –0.2 (Feldman et al. 1989) to –1.8 (Royalty, Solomon 1999).

6 Only those health insurers are excluded that did not have open enrolment and whose members were not allowed to switch – e.g a health insurer for miners and one insurer for sailors.

Table 1

Contribution rates of health insumers
2001 – 2004; %

	1/2001	1/2002	7/2002	1/2003	7/2003	1/2004	4/2004
Mean	12.83	13.18	13.29	13.56	13.68	13.74	13.77
SD	0.82	0.80	0.80	0.87	0.85	0.77	0.75
Min	11.2	11.2	10.2	10.2	10.2	10.2	10.2
Max	15.3	14.9	14.9	15.7	15.7	15.7	15.7

Authors' calculations.

on aggregated data or on information about individual companies that covered only a small part of the market (Schut et al. 2003).

Table 1 shows descriptive statistics about the development of contribution rates of health insurers. The mean of the contribution rate increased over the whole period while the standard deviation decreased only after 2003. The spread between the lowest and the highest contribution rate is quite large. In 2004 switching from the company with the highest to the one with the lowest contribution rate generates a combined annual saving of 2300 € for an individual and his employer if salaries meet the maximum income ceiling. However, most of the health insurers operate on a regional level. Accordingly, the spread of contribution rates is much smaller in most regions. Thus, potential savings might be much smaller but are still substantial.

Table 2 describes the concentration of the market on a national level. Some of the biggest insurance companies lost market shares over time. Accordingly, the market share of the biggest 10 and the biggest 5 companies declined somewhat from 2001 to 2004. On a national level, the biggest company has a market share of 11 percent.

3. Empirical Framework

3.1 The Model

The focus of this study is the choice among health insurers within the German social health insurance system. Hence, this company-data based analysis tries

Table 2

Concentration of the social health insurance market
market shares of companies in %

	1/2001	1/2002	7/2002	1/2003	7/2003	1/2004	4/2004
Top 10	59.8	59.2	58.3	57.7	56.8	56.5	56.5
Top 5	43.8	43.2	42.5	42.0	41.0	40.7	40.7
Min	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01	< 0.01
Max	12.5	12.4	12.0	11.8	11.4	11.3	11.2

Authors' calculations.

to explain the market share of health insurer *i* in time *t*, s_{it}. The number of insurers in the market is *N*. To account for the fact that the endogenous variable is bounded between zero and one the model is specified as a conventional logistic function

$$s_{it} = \frac{\exp(\beta' x_{it} + \gamma_i + \varepsilon_{it})}{\sum_{l=1}^{N} \exp(\beta' x_{lt} + \gamma_l + \varepsilon_{lt})} \quad i = 1, \ldots, N. \tag{1}$$

Here, x_{it} represents a vector of explanatory variables, where *i* and *l* indicate insurance companies. In our basic specification x_{it} solely consists of the contribution rate p_{it}. The parameter γ_i captures unobserved heterogeneity across insurance companies, and ε_{it} is a random error term. Taking logarithms leads to a quite convenient linear representation of the model,

$$\log(s_{it}) = \beta' x_{it} + \delta_t + \gamma_i + \varepsilon_{it} \quad i = 1, \ldots, N. \tag{2}$$

δ_t captures the logged numerator of equation (1) which is a time specific constant. By estimating a time-fixed-effects model any variation that is purely across periods and that affects all companies alike is removed.

It is important to keep in mind that the market shares of health insurers are ultimately determined by consumers' decisions. Our specification (1), therefore, directly relates to a discrete choice model for individual consumers (for details see Scanlon at al. 2002). However, consumers are not forced to explicitly decide on their health insurer each period and many are likely to avoid transaction and information costs by simply staying with their current insurer without considering any alternative. Therefore, the model should account for potential persistence of market shares and, correspondingly, is augmented by including the lagged endogenous variable on the right hand side

$$\log(s_{it}) = \alpha \log(s_{it-1}) + \beta' x_{it} + \delta_t + \gamma_i + \varepsilon_{it} \quad i = 1, \ldots, N, \tag{3}$$

with $0 \le \alpha \le 1$ capturing the degree of persistence. The resulting dynamic model exhibits two quite interesting limiting cases.

If $\alpha = 0$ holds, the model coincides with the simple static one. This corresponds to a world where each individual decides on its health insurance each period. Since company specific effects γ_i reflect that the probability of being chosen systematically differs across companies, any persistence in market shares is – besides persistent deviations in premium rates – then solely due to these differences. Transitory variation in x_{it}, i.e. in the contribution rate, has only short-term and no long-term effect on market shares.

If in contrast $\alpha = 1$ holds, market shares are non-stationary, following a random walk. Hence, even if contribution rates remained stable no equilibrium of market shares exists and only changes in market shares but not their level can be explained. Consequently, transitory changes in the explanatory variables have permanent effects on market shares. A health insurer, for instance, which looses members because of a relatively high contribution rate, will not be able to attract them again by simply closing the premium gap. By closing the gap the insurer will only be able to prevent more clients from leaving. Finally, in the intermediate case ($0 < \alpha < 1$) transitory variation in explanatory variables has long-lasting effects on market shares. However, these effects fade out over time.

In correspondence to this discussion one might formulate different measures for consumers' price sensitivity. Our primary interest is on short-run premium elasticities of market shares η_{it}, which are directly comparable to the results obtained from earlier work based on static models. According to our model short-run premium elasticities of market shares are equal to

(4) $$\eta_{it} \equiv \frac{\partial s_{it}}{\partial p_{it}} \frac{p_{it}}{s_{it}} = \beta_p (1 - s_{it}) p_{it}.$$

β_p denotes the coefficient referring to the contribution rate. In addition we are interested in long-run effects, i.e. in dynamic multipliers $\partial \log(s_{iT}) / \partial \log(p_{it}) = \alpha^{T-1} \eta_{it}, T > t$ that capture the effect of a transitory change in the contribution rate on future market shares and ultimately the long-run effect of a permanent change

(5) $$\lim_{T \to \infty} \sum_{t=0}^{T} \frac{\partial \log(s_{iT})}{\partial \log(p_{it})} = \lim_{T \to \infty} \sum_{t=0}^{T} \alpha^{T-t} \eta_{it} \approx \frac{1}{1-\alpha} \eta_{it}.$$

Here the approximation is accurate for small s_{it}. Obviously, for the static model, (5) equals (4) since changes in the contribution rate only have instantaneous effects. In contrast, for $\alpha = 1$ the long-run effect of a permanent change in the contribution rate exceeds all limits, even if the short-term elasticity were small.

3.2 Estimation

While the limiting case specifications $\alpha = 0$ and $\alpha = 1$ can easily be estimated using conventional panel data techniques, this does not hold for the unrestricted dynamic specification. Therefore this subsection shortly discusses several methodological aspects for estimating this type of dynamic panel data model. An intuitive and more detailed survey on these models is given in Bond (2002).

By construction, in equation (3) the right hand side variable $\log(s_{it-1})$ is correlated with the composite error $(\gamma_i + \varepsilon_{it})$, leading to inconsistent estimates, even if individual heterogeneity is accounted for by either fixed- or random-effects (e.g. Baltagi 2001). When employing a simple instrumental variable estimator (Anderson, Hsiao 1982) consistency can be achieved under the assumption of serially uncorrelated errors ε_{it}. This estimator is based on first-differencing the regression equation and using twice lagged dependent variables as instruments for $\Delta \log(s_{it-1})$. If, however, the number of waves exceeds three and any explanatory variables x_{it} are taken into account, the IV-estimator's efficiency can substantially be improved by using higher-order lagged endogenous variables (i.e. $\log(s_{it-3})$, $\log(s_{it-4})$, ... instead of only $\log(s_{it-2})$) and – potentially – past, present and future values of x_{it} as additional instruments. This model is estimated within a GMM framework (Arellano, Bond 1991 for details). For GMM, the number of valid instruments varies whether the explanatory variables x_{it} are exogenous, predetermined or endogenous with respect to the error term ε_{it}. That is, whether $\mathrm{cov}(x_{it}, \varepsilon_{i\tau}) = 0$ holds for any t and τ, only for $\tau \geq t$, or just for $\tau > t$, respectively.

Fortunately, as long as the model is over-identified, Sargan tests are available for testing the validity of the underlying assumptions (the so-called moment conditions) and therefore serve as a basis for selecting an appropriate specification. The basic Sargan test (Sargan 1958; Hansen 1982) examines whether the whole set of moment conditions is valid, e.g. if the instruments used for estimation are valid and, hence, orthogonal to the error term. By contrast, the difference Sargan test focuses on the validity of the additional moment conditions and the additional instruments, only. The difference Sargan test is obtained by comparing the Sargan statistics of a restricted and an unrestricted model, the restricted also including the additional instruments and moment conditions.[7]

In this paper estimation of GMM is implemented using an efficient two-step procedure[8]. The conventional standard errors obtained from two-step GMM are likely to be underestimated in small samples. However, this can be adjusted for by using a corrected variance estimator proposed by Windmeijer (2005). This corrected variance estimator has been shown to exhibit at least equally good properties as the variance estimator obtained from one-step GMM (Bond, Windmeijer 2002). Henceforth, in the reminder we only present corrected standard errors and focus on two-step GMM.

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7 In the sequel we stick to the conventional notation and refer to (difference) Sargan tests, although we actually present Hansen's J-statistic based on robust estimates. Hansen's J-statistic is preferable due to bias of the original Sargan statistic in the presence of heteroskedasticity.

8 The second step is based on a weighting matrix which is estimated from the results of a consistent first step estimation.

All variants of GMM estimators discussed so far – likewise the simple Anderson-Hsiao IV-estimator – are based on moment conditions specified in terms of first-differenced equations (first-differenced GMM). Unfortunately, if α is close to unity, first-differenced GMM might suffer from small sample bias and imprecision, and often performs poorly. It might therefore be less suited if high persistence is prevailing, as is the case with our data. System GMM (Arellano, Bover 1997) exploits additional moment conditions specified in terms of levels rather than differences and is a potential solution to the problem. However, the additional moment conditions are valid only if certain assumptions on initial conditions are satisfied, i.e. the error term in the first period, ε_{i1}, and the first-differenced exogenous variables, Δx_{it}, have to be uncorrelated with the individual specific effect γ_i (Blundell, Bond 1998). Once again, these moment conditions can be tested on the basis of (difference) Sargan tests.

From the above discussion it becomes obvious that a rather large number of different GMM specifications can potentially be used for estimating the model. Therefore, the following section presents estimation results for several specifications and additionally discusses our strategy how to select the most appropriate one.

Unfortunately, our panel data is unequally spaced. This might lead to inconsistent estimates, both in the GMM case and in the case with $\alpha = 1$, as long as one insists that one period in the theoretical model has to coincide perfectly with a certain time span in the empirical data.[9] Under this assumption, McKenzie (2001) shows that the model is misspecified by a simple AR(1) process. First-differencing as well as including fixed-effects, therefore, fails to remove the individual heterogeneity. Since GMM is quite data consuming we nevertheless precede using all data available, i.e. the 7 unequally spaced waves, though keeping this problem in mind. Therefore we also check our results by comparing them to those obtained from a reduced sample which only includes waves 2 to 6 (Appendix B). These waves are equally spaced at semi-annual intervals.

4. Results

This section provides the estimation results of our empirical analysis of market shares of health insurers in Germany. We consider several regression models. First, we present a static panel data model where the market share of each individual company is determined by company specific individual effects and current contribution rates and never deviates from equilibrium. Then, we consider a dynamic model where the market share follows a stationary first-order

9 This assumption is likely to be regularly violated in survey data, since the date of interview will vary for practical reasons.

autoregressive process, i.e. market shares are persistent, yet differences in contribution rates can lead to long-lasting changes in the shares which fade out after some time. Finally, we provide results for a model where market shares are considered to follow a non-stationary unit-root process ($\alpha = 1$), i.e. differences in contribution rates lead to permanent changes in market shares and, hence, the model explains first differences of market shares. In order to discriminate between the stationary and the unit-root process we also provide results of panel unit-root tests. Subsequently, we test for differences in price elasticities for different types of insurance companies and for varying time periods.

4.1 Static model

First, we present the results of a static panel data model. The market share of each individual company is solely determined by its current contribution rate and by company specific individual effects. The company specific effects represent unobservable factors of health insurers which influence consumers in their choice between companies and which might be correlated with the contribution rate. Examples for such factors not included in the data are the number of branch offices, the quality of service the insurers provide, or any additional medical treatments not compulsorily covered by the standard benefit package which, however, are covered by some of the companies.

The situation can be modeled as in equation (2). Here, it is assumed that the individual specific effects are invariant over time. Depending on the assumptions on the correlation between individual specific effects and contribution rates the model can either be estimated using random-effects or using fixed-effects. The random-effects specification is clearly rejected by a Hausman test (Hausman 1978). Therefore we only present results of the fixed-effects panel data model. As can be seen in Table 3 the sign of the coefficient of the contribution rate is negative but the coefficient is clearly insignificant. I.e. after controlling for company specific effects the contribution rate has no influence on the market share.

4.2 Dynamic models

4.2.1 Generalized Method of Moments (GMM)

The dynamic model is equivalent to a world where only some consumers decide about staying with their health insurer or choosing a new one. Our estimation is based on equation (3). We compare several specifications based on different moment conditions or sets of instruments.

Before reporting the GMM estimates we shortly mention the results of a simple OLS regression and of a panel fixed-effects model (within-group-estimator) which both include the lagged market share as explanatory variable. The

Table 3

Panel fixed-effects estimates for static model

	Coefficient	Std. error
Contribution rate	–0.0045	0.0274
Within-R^2	0.1546	
F-Test	12.32***	
Observations	1960	
Hausman test (χ^2)	73.11***	

Authors' calculations. – Regression includes time dummies for each wave. Huber-White robust standard errors given. – ***Indicates significance at 1% level; **at 5% level; *at 10% level.

OLS estimates for the coefficient on the lagged market share are known to be biased upwards, whereas for the panel fixed-effects model they are known to be biased downwards. Hence, the two point estimates 1.0053 and 0.8122 (Table 8 in the Appendix) might serve as upper and lower bound for the true coefficient.[10]

The first two specifications in Table 4 are based on an Arellano-Bond type first-differenced GMM estimator. In column 1 we present the results for a specification where the contribution rate is assumed to be predetermined, all available instruments are used and the estimation is done by two-step GMM (GMM1). The contribution rate has a negative effect on the market share but is clearly insignificant and the lagged market share has a coefficient (α) close to one. Furthermore, the statistic of the Sargan test is highly significant, indicating that some of our over-identifying restrictions are not valid. A test where the matrix of possible instruments has been reduced to a minimum, i.e. $\Delta \log(s_{it-1})$ and Δx_{it} are instrumented by only one instrument each, indicates that the additional restrictions are not valid since the difference-Sargan test is significant ($\chi^2_{(25)} = 45.76$).

Yet, when the contribution rate is treated as endogenous (GMM2) the Sargan statistic becomes insignificant at the 5%-level. The difference-Sargan test between GMM1 and GMM2 is significant and clearly confirms these findings ($\chi^2_{(5)} = 17.38$). Therefore, we conclude that the contribution rate indeed is endogenous. In GMM2 the estimated coefficient for α is lower than before but still relatively close to unity and the contribution rate is insignificant. Since the first-differenced GMM model is only weakly identified if α is close to unity we might favor the system GMM estimator in this case. Therefore, we now go on considering an Arellano-Bover type system GMM estimator which includes additional moment conditions and therefore allows identification of the model even if α is close to unity.

10 This procedure has been proposed and applied in Bond (2002), at least for models without additional explanatory variables.

Table 4

GMM estimates for dynamic panel data model

	First-differenced GMM				System GMM			
	x_{it} predetermined GMM1		x_{it} endogenous GMM2		x_{it} predetermined GMM3		x_{it} endogenous GMM4	
	Coef.	Std. error	Coef.	Std. error	Coef.	Std. error	Coef.	Std. error
Market share in t–1	0.9798***	0.0751	0.9525***	0.0378	1.0123***	0.0191	1.0453***	0.0220
Contribution rate	–0.0034	0.0545	–0.0413	0.0451	–0.1187***	0.0387	–0.1715***	0.0307
Observations	1221		1221		1588		1588	
AR(1)	–3.32***		–3.87***		–4.16***		–4.41***	
AR(2)	0.12		0.16		–0.05		–0.12	
Sargan statistic	57.65***		40.27*		73.42***		57.34**	
Diff.-Sargan test (fewer instruments)	45.76*** (25)		21.83 (20)		32.12 (25)		24.69 (20)	
Diff.-Sargan test (system vs. first-dif. GMM)					15.77 (14)		17.07* (10)	

Authors' calculations. – Regression includes time dummies for each wave. Two-step GMM estimates with corrected standard errors (Windmeijer 2005). AR(1) and AR(2) are tests for first- and second-order serial correlation in the first-differenced residuals (Arrelano, Bond 1991). (Difference) Sargan statistics are χ^2 distributed; number in brackets behind difference Sargan test provides the number of restrictions/degrees of freedom. – ***Indicates significance at 1% level; **at 5% level; *at 10% level.

The system GMM estimates are also provided in Table 4. We see that in all of the system GMM models the coefficient for the contribution rate is significant and of much higher magnitude than in the first-differenced GMM models. As before, we once more start with a specification where the contribution rate is assumed to be predetermined (GMM3). These estimates show highly significant Sargan statistics. This Sargan statistic for invalid assumptions does not drop to an insignificant level if either (i) the matrix of possible instruments in reduced (not reported), nor if (ii) the contribution rate is considered to be endogenous (GMM4). For GMM4 a comparison with GMM2 indicates that the additional moment conditions in the system GMM are rejected to be valid at the 10%-level (difference-Sargan test: $\chi^2_{(10)} = 17.07$). This might indicate that the market shares observed in the first period systematically deviate from equilibrium shares conditional on contribution rates and individual effects. Taking into account that changes between insurance companies were heavily restricted if not impossible for consumers before 1996, it is quite plausible that these market constraints led to strong deviations from equilibrium under market conditions which had not been neutralized until our data sample starts in 2001. Hence, system GMM seems to rely on inappropriate assumptions in the case analyzed here.

Summing up, due to the Sargan tests these findings are in favor of the first-differenced GMM specification including the contribution rate as endogenous regressor (GMM2), although being close to a unit-root process and, hence, poor precision of the estimates.

4.2.2 Model in first differences

In this subsection we finally provide the results for a model as in equation (2), however in this case our dependent variable is not the market share but the first difference of market shares, $\Delta \log(s_{it})$. This refers to a world where transitory differences in contribution rates lead to permanent changes in market shares. Once a consumer changed his insurance company he will stay with the new company as long as no further differences in contribution rates prevail or any unsystematic effects occur. In the long-run there are no insurance companies which are big because they started big and there are no companies which are small because they started small, because as long as there are differences in contribution rates, market shares will change and these changes will be permanent.

In this model, which explains changes in market shares rather than levels, individual effects represent company specific drifts. Such drifts might, for instance, be due to death rates that vary across companies. In fact, an F-test on joint significance of the company specific fixed-effects is highly significant indicating that a fixed-effects model (UR2) is preferable to a simple OLS regression (UR1). Results for the fixed-effects as well as the OLS regression are provided in Table 5. In both models the contribution rate is highly significant and the magnitude is comparable.

However, taking into account the results from the GMM estimates in the preceding subsection one might doubt if the contribution rate is exogenous. Therefore we instrument the contribution rate by its own lagged value and provide the results of IV estimations in Table 5, too. In the case without fixed-effects (UR3) we only observe a very small change in the coefficient for the contribution rate. Yet, a test for endogeneity[11] indicates that the variable indeed is endogenous. In contrast, in the case with fixed-effects (UR4) the coefficient for the contribution rate almost doubles, there however a test for endogeneity is insignificant. Hence, we regard UR2 as our preferred specification in the class of the unit-root models and focus further discussion on UR2.

[11] The test for endogeneity proposed in Wooldridge (2002) is based on the estimated residual from a regression of the instrumented variable on all other exogenous variables. These estimated residuals are then included in the original regression. The instrumented variable is concluded to be endogenous if the estimated residuals have significant explanatory power in the original regression.

Table 5

Estimates for unit root case, i.e. dependent variable is first-differenced market share

	OLS model UR1		Fixed-effects UR2		IV UR3		IV fixed-effects UR4	
	Coef.	Std. error	Coef.	Std. error	Coef.	Std. error	Coef.	Std. error
Contribution rate	-0.0890***	0.0089	–0.0814***	0.0130	–0.0960***	0.0093	–0.1670***	0.0495
Observations	1589		1589		1589		1589	
R^2	0.1493				0.1484			
F-Test	20.36***		10.59***		21.33***			
Test for endogeneity (t-statistic)					2.09**		1.55	

Authors' calculations. – Regression includes time dummies for each wave. Huber-White robust standard errors given. – ***Indicates significance at 1% level; **at 5% level; *at 10% level.

4.2.3 Unit root tests

Whether the parameter α takes the limiting value of unity is a crucial question for both the theoretical model and the adequate estimation procedure. Our GMM estimates indicate that α is at least close to unity. Unfortunately, first-differenced GMM[12] is ill-suited for testing the null-hypothesis of a unit-root being present, since the model is not identified under the null. For this reason, additional panel unit-root test are carried out. Two distinct regression-based test procedures are considered which both are well suited for panels with a large number of cross-sectional units but a small number of waves. One test is based on a simple OLS regression of market shares on their lagged values. This leads to consistent estimates under the null. The second test (proposed by Breitung, Meyer 1994) specifies the regression in terms of deviations from initial conditions and therefore, is likely to be more powerful if the variance of the individual effects is high. Otherwise, it may lose power in comparison to simple OLS (Bond et al. 2002). Table 6 displays the test results for market shares and for contribution rates.

Results for the contribution rate are quite clear. The unit-root case is rejected by both tests. Hence, there is no need to formulate the model in terms of changes in contribution rates in order to avoid spurious regression. This result is quite plausible from the view-point of economic theory. By the use of time-fixed-effects the model is implicitly formulated in terms of deviations from cross-section means.[13] In fact, deviations form average prices cannot be non-stationary, since market forces will always impose pressure on price differences.

12 The highly significant Sargan statistics indicate, however, that system GMM is not consistent.

13 These issues are slightly more involved in the case of unbalanced panels (Wansbeek, Kaptyen 1989).

Table 6

Panel Unit-Root Tests

	Market share		Contribution rate	
	Simple OLS	Breitung-Meyer	Simple OLS	Breitung-Meyer
Coefficient lagged value	0.9844	1.2094	0.8743	0.9309
t-statistic	–6.2956	5.5407	–9.2923	–4.9029
p-value (one sided t-test)	0.000	1.000	0.000	0.000

Authors' calculations. – Regression includes time dummies for each wave. Regressions with/without uniform constant yield similar results. Heteroskedasticity-robust standard errors given.

Unfortunately, the different tests yield different results for market shares. While simple OLS clearly rejects the hypothesis of a unit-root being present in the series of market shares, the Breitung-Meyer test does not reject the null at any plausible level of significance. Therefore, these test results do not unambiguously answer the question whether the model has to be formulated in terms of first differences of market shares or as a more general dynamic one.

Taking together the information from these panel unit-root tests and the estimated results from our GMM specification (GMM2) and the specification of first-differenced market shares with fixed-effects (UR2), we conclude that both are quite similar. Although being found to be only weakly identified in comparable cases, the GMM2 estimates of α are not significantly different from one and therefore overlap with the unit-root case. Furthermore, the confidence interval of the estimate for β in GMM2 includes the point estimate of the other specification (UR2). Hence, these results are comparable. We prefer the results of UR2, mostly because of identification problems with GMM2 in small samples and its low precision in terms of large standard errors, but go on presenting further results for both specifications. The preference for UR2 is reinforced by the results based on waves 2 to 6, only, i.e. the equally spaced panel (Appendix B).

4.3 Model extensions

In order to check sensitivity of the price sensitivity of market shares with respect to (i) the chosen time period, (ii) the type of health insurance company, (iii) open vs. restricted enrolment in companies, and (iv) regional restrictions of insurance companies we estimate several models where we include interaction terms between contribution rate and time, type of insurance company, and other group indicators, respectively. These sensitivity checks are performed for our preferred models from the above specifications, i.e. GMM2 and UR2.

Sensitivity of price elasticities with respect to time is analyzed for two separate cases. First, we check for a time trend of the coefficient for the contribution rate and then we estimate the coefficient for two different time periods, i.e. the baseline coefficient represents waves 1–4 and an interaction term is included for period 2 which represents waves 5–7. For detailed results, see Table 9 in the Appendix. In both specifications, the GMM model seems to argue in favor of a recent increase in consumers' price sensitivity, while the unit-roots specification does not indicate any significant change. Nevertheless, the insights from the GMM models are weak. F-tests on the overall price effect in later periods do not indicate a significant price sensitivity. That is, in model GMM2a, the sum of the coefficient for the contribution rate and the interaction term between contribution rate and period 2 being different from zero is rejected ($H_0: \beta_p + \beta_{p\ period2} = 0$ and $F_{1,324} = 0.62$). In the specification with time trend (GMM2b), a comparable F-test ($H_0: \beta_p + t \cdot \beta_{p\ trend} = 0$) is also insignificant for $0 < t \leq 7$.

We distinguish five types of insurance companies which are large regional companies (*Allgemeine Ortskrankenkassen* AOK), two types of substitute companies (*Ersatzkassen für Angestellte* EAN and *Ersatzkassen für Arbeiter* EAR), guild-based insurance companies (*Innungskrankenkassen* IKK), and company-based insurance companies (*Betriebskrankenkassen* BKK). The latter is our reference group. In the specification for the unit root case as well as in the GMM case, the interaction-terms between contribution rate and type of insurance company are all insignificant both individually and jointly (Table 10 in the Appendix).

As already national, most companies operate only in one or few regions. Hence, the relevant market of companies with local restrictions is much smaller and might restrict their potential to attract new clients. Yet, our results show no significant differences in consumers' response to contribution rates between companies with and without regional restrictions (Table 11 in the Appendix).

Although being able to opt for open enrolment some of the BKK and IKK decide to have restricted enrolment. This is due to their roots as company- and guild-based insurance companies respectively. Hence, only those consumers working at the company or being member of a certain profession are eligible to enroll in these insurance companies, though are not forced to do so. Therefore, the relevant market of these companies is much smaller than of those with open enrolment. However, we do not see any differences between these companies either (Table 12 in the Appendix).

4.4 Elasticities

After having discussed several specifications for estimating the effect of price on market share, we now present the premium elasticities of market shares implied by the estimated coefficients. The point estimates of the elasticities and

Table 7

Estimates of the short run premium elasticity

	Static Model		GMM2		UR2	
Mean Premium elasticity	–0.06		–0.55		–1.09	
95%-confidence interval	–0.78	+0.66	–1.74	+0.64	–1.43	–0.75

Authors' calculations. – Elasticity estimated for sample mean. Estimation based on result from Tables 3, 4 and 5.

the corresponding 95%-confidence intervals for the static model, GMM2 and UR2 are given in Table 7. The short-run premium elasticity is based on equation (4) and calculated for the sample mean.

The point estimates of the elasticity for the static model and for GMM2 are both negative, however, due to insignificance of β_p, they are not significantly different from zero. Yet, our preferred estimate for UR2, i.e. the unit-root specification with fixed-effects, is significant. At the one hand, UR2 is preferred to the simple OLS unit-root specification. At the other, it displays more robust results and by far smaller standard errors than competing dynamic specifications estimated using GMM. Moreover, it is favored against the static model. Finally, estimates presented in the preceding subsection do not argue in favor of including interaction terms of the premium rate and other explanatory variables in the model.

At sample mean UR2 displays a short-term premium elasticity of about minus one. This indicates a distinct sensitivity of consumers to differences in contribution rates. Thus a crucial precondition for managed competition to work is fulfilled.

5. Conclusions

This paper provides two important elements that advance insights in the dynamics of the German social health insurance market. First, it is based on a unique panel data set which covers the social health insurance market completely on the level of individual insurance companies. Prior to this study, only data on an aggregated level with very few observations were available. Second, this paper uses an advanced econometric technique which takes into account the dynamics of the market. So far, studies on price elasticities in the German social health insurance market have been based on static models only.

The econometric analysis favors a dynamic model which explains changes in market shares by the level of premiums. For this specification we obtain a short-run premium elasticity of market shares of approximately minus one. This indicates a moderate short-run sensitivity of consumers to differences in contribution rates. Compared to earlier analyses dealing with the German case, e.g. Schut et al. (2003), our elasticity is smaller than the one estimated there. Interestingly, our results are much closer to those obtained for other countries like Switzerland (Beck 2004). However, from the point of view of economic theory, the estimated short-run price sensitivity appears to be rather small keeping in mind that consumers can choose between almost perfect substitutes; for perfect substitutes the price elasticity should approach infinity. Hence, one might hypothesize that different health insurers are not perceived as perfect substitutes by many consumers.

Yet, in contrast to earlier analyses our results are based on a dynamic specification. Our results indicate that market shares follow a unit-root process or are, at least, close to non-stationarity. That is, even if the price sensitivity might appear to be rather moderate in the short-term, in the long-run permanent relative changes in contribution rates will have dramatic effects on the market shares of health insurers. Insurers which permanently charge contribution rates that are higher than those of competitors and do not offset this by being attractive to consumers for other reasons than price, will ultimately drop out of the market. However, this process might take some time.

Clearly, we have been able to show that consumers exert their right to choose among social health insurers, that the choice is sensitive to price, and that therefore major conditions for managed competition to work are fulfilled. Furthermore, our results show that this will – at least in the long-run – impose substantial pressure on health insurers. In other words, "the prospect of being hanged" is real. Yet, it is less clear whether this will ultimately lead to enhanced efficiency as intended by the reform of 1996. Other – possibly more promising – strategies to reduce the premium are available, e.g. risk selection strategies (Jacobs et al. 2002; Behrend et al. 2004). Analyzing gains and losses in efficiency, therefore, remains a topic for future research.

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Appendix A – Tables

Table 8

OLS and within group estimates for dynamic panel data model

	OLS		Within group estimation	
	Coefficient	Std. error	Coefficient	Std. error
Market share in t-1	1.0053***	0.0027	0.8122***	0.0558
Contribution rate	–0.0952***	0.0101	–0.0675***	
R^2	0.9910			
F-Test	32562.03***		39.54***	
Observations	1588		1588	

Authors' calculations. – Regression includes time dummies for each wave. Huber-White robust standard errors given. – ***Indicates significance at 1% level; **at 5% level; *at 10% level.

Table 9

Estimation of differences in price sensitivity by time period

	Time trend				2 separate periods			
	GMM2a		UR2a		GMM2b		UR2b	
	Coef.	Std. error	Coef.	Std. error	Coef.	Std. error	Coef.	Std. error
Market share in t-1	0.8398***	0.0661	1	fixed	0.8557***	0.0544	1	fixed
Contrib. rate	–0.0062	0.0585	–0.0978***	0.0118	–0.0165	0.0507	–0.0785***	0.0120
rate · trend	–0.0094*	0.0059	0.0017	0.0027				
rate · period 2					–0.0313**	0.0157	0.0090	0.0095
Observations	1221		1589		1221		1589	
Sargan test	39.34* (27)				37.71* (27)			

Authors' calculations. – Regression includes time dummies for each wave. Period 2 includes waves 5–7, i.e. waves 1–4 are the omitted categories. Sargan statistics are χ^2 distributed; number in brackets provides the number of over-identifying restrictions/degrees of freedom. – ***Indicates significance at 1% level; **at 5% level; *at 10% level.

Table 10

Estimation of differences in price sensitivity by type of insurance company

	GMM2c		UR2c	
	Coef.	Std. error	Coef.	Std. error
Market share in t-1	0.9449***	0.0345	1	fixed
Contrib. rate	–0.0502**	0.0207	–0.0835***	0.0117
rate · AOK	–0.0052	0.0633	0.0600	0.0732
rate · EAN	0.0097	0.0213	0.0365	0.1054
rate · EAR	0.0192	0.0299	0.0172	0.0818
rate · IKK	0.1030	0.1545	0.0433	0.0503
Observations	1218		1585	
Sargan test	77.96 (74)			
F-test (all interactions = 0)	0.36		0.38	

Authors' calculations. – Regression includes time dummies for each wave. Omitted type of insurance company (baseline) is BKK. – ***Indicates significance at 1% level; **at 5% level; *at 10% level.

Table 11

Estimation of differences in price sensitivity by regional restriction of insurance companies

	GMM2d		UR2d	
	Coef.	Std. error	Coef.	Std. error
Market share in t–1	0.9766***	0.0690	1	fixed
Contrib. rate	–0.0496	0.0479	–0.0873***	0.0219
rate · regional market only	–0.0149	0.0310	–0.0349	0.0232
Observations	856		1096	
Sargan test	51.43 (42)			

Authors' calculations. – Regression includes time dummies for each wave. – ***Indicates significance at 1% level; **at 5% level; *at 10% level.

Table 12

Estimation of differences in price sensitivity between insurance companies with open and restricted enrolment

	GMM2e		UR2e	
	Coef.	Std. error	Coef.	Std. error
Market share in t–1	0.9455***	0.0440	1	fixed
Contrib. rate	–0.0318	0.0730	–0.0854***	0.0121
rate · no open enrolment	0.0020	0.0375	–0.0035	0.0033
Observations	1122		1435	
Sargan test	83.86*** (42)			

Authors' calculations. –Regression includes time dummies for each wave. – ***Indicates significance at 1% level; **at 5% level; *at 10% level.

Appendix B – Equally spaced time periods

As mentioned in subsection 3.2 estimation of dynamic panel data models might lead to inconsistent estimates if panel waves are unequally spaced. Therefore we reduce our data set to waves 2 to 6, i.e. to the waves collected semi-annually between January 2002 and January 2004, and compare the results with those reported above.

The two unit-root tests do not unambiguously discriminate between $\alpha = 1$ and $\alpha < 1$, either. In this setting, the GMM estimates are even weaker than those reported for the larger panel. The Sargan statistic is highly significant in all of the models, rejecting the underlying orthogonallity assumptions altogether. Yet, the point estimates of the coefficient are within the range of the results obtained from estimating the model using the whole sample. For the unit root case (corresponding to UR2) the point estimate for β_p is somewhat closer to zero (–0.0530) but still highly significant. If only five semi-annually spaced panel waves are considered, results do not qualitatively change with respect to time period, type of insurance company, regional restrictions or open vs. restricted enrolment. Estimation results based on the reduced sample consisting of regularly spaced waves, therefore, do not challenge the main findings of our analysis.