

ELASTICITIES OF MARKET SHARES AND SOCIAL HEALTH INSURANCE CHOICE IN GERMANY: A DYNAMIC PANEL DATA APPROACH

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SUMMARY

In 1996, free choice of health insurers was introduced to the German social health insurance system. One objective was to increase efficiency through competition. 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. Copyright © 2006 John Wiley & Sons, Ltd.

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INTRODUCTION

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).

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

This paper tests the last assumption – for competition between health insurers to be effective, (i) consumers must have some choice between them. In Hirschman's terminology: consumers must have

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the exit option (Hirschman, 1970). But in order for the reforms (which provide more choice) to work, (ii) consumers must also be inclined to search for lower-priced health insurers or insurers with higher quality. Thus, the main hypothesis to be tested in this paper is how sensitive consumers react to differences in prices of health insurance and whether health insurers with lower prices have bigger market shares or shares that grow over time.

Consumers' insurance choice behavior has been analyzed before, in Germany as well as in other countries. Schwarze and Andersen (2001) use German micro-data and provide a descriptive analysis of the socioeconomic background of 'switchers' and 'non-switchers'. Nuscheler and Knaus (2005) use the same data source to estimate a model of switching behavior that ultimately focuses on the issue of risk selection. However, their analysis has to rely on information on premiums aggregated at the level of insurer types (for an explanation of these types see the next section). 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 to increase over time (Schut *et al.*, 2003). The latter study is closely related to ours. However, they use data on health insurers aggregated by type of insurer and estimate 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 and Hassink, 2002; Schut *et al.*, 2003).¹ Because the institutional arrangements in these countries are not the same, different elasticities are hardly surprising.

Our analysis contributes to this literature by estimating the price sensitivity of consumers in Germany, using information on individual health insurers. The underlying model relates market shares with prices and is based on consumer choice. Moreover, we apply a range of econometric techniques that hitherto have not been applied to the research question; in particular, we analyze dynamic models. Modeling a dynamic process seems to be more appropriate than modeling a static one, because only a small number of consumers is inclined to decide on their health insurer each period, and the market is therefore likely to display persistence. We use a new data set that covers a recent period and is based on a panel of individual health insurers. Prior studies for Germany have had to rely on data aggregated over insurers only, making our panel data on individual health insurers 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. Next section describes those institutional features of social health insurance in Germany that are relevant for consumer choice. Third section describes the data set that has been collected for analyzing price sensitivity of consumers and fourth section specifies the econometric models used. Our estimation results are presented in the penultimate section. The conclusion can be found in the last section.

INSTITUTIONAL FEATURES OF SOCIAL HEALTH INSURANCE IN GERMANY

In the German social health insurance market, risk-bearing health insurers compete for enrollees. In principle, it is mandatory for all employees to acquire social health insurance. Yet, when the salary exceeds a threshold, individuals can choose whether to remain in the social health insurance system or to opt out and buy private health insurance; also self-employed and civil servants can opt out.²

¹ 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 and Solomon, 1999).

² In this paper, we do not address the choice between social and private health insurance, but focus instead on consumer choice within the social health insurance system. This means that our analysis does not explain the total number of clients in the social health insurance system, but is based on the market shares of health insurers within this system. In the remainder of this paper, therefore, the term health insurer only refers to social health insurers.

Consumers are allowed to switch between social health insurers on a regular basis. The only restriction is that they must have stayed at their previous insurance company for at least 18 months. However, if their health insurer raises its premium, they can switch immediately.

German social health insurers do not calculate individual risk-dependent premiums. Instead, they set contribution rates. Individual premiums are then equal to salaries times the contribution rate, up to an income ceiling.³ Half of the premium is paid by the employers; the other half is paid by the employees.

By legal restriction, competition between health insurers is almost exclusively based on price, i.e. the contribution rate that varies across insurers. More than 95% of the benefits package is standardized. There are only few services, such as spa treatment or the implementation of disease management programs, for which it is up to the insurance company to decide whether or not to include these services in the benefits package, and to what extent (Greß *et al.*, 2005). Contribution rates and differences in benefits packages between health insurers are being published by consumer associations more often and, thus, have become more transparent for consumers. Moreover, health insurance companies are obliged by law to contract collectively with all licensed health care providers. Legal opportunities to contract selectively are very limited, for instance for so-called 'integrated care contracts', combining inpatient and ambulatory care (Greß *et al.*, 2006). Hence, the quality of insurance is basically identical among companies. For this reason, one would expect the majority of consumers to choose low premium insurers. However, somewhat surprisingly, several insurers that charge rather high contribution rates are still among the largest companies in the market. This has prompted us to have a closer look at how consumers do react to price differences.

Another institutional aspect that might influence insurance choice is that insurance companies have historically been grouped by insurer types. There are large regional insurance companies (*Allgemeine Ortskrankenkassen* AOK), two types of so-called substitute companies (*Ersatzkassen für Angestellte* EAN and *Ersatzkassen für Arbeiter* EAR), guild-based companies (*Innungskrankenkassen* IKK), and company-based insurance companies (*Betriebskrankenkassen* BKK). Prior to 1996, specific groups of clients were directly assigned to certain types of insurer. Although there is no such assignment nowadays, some insurer types launch collective advertisement campaigns and, thus, companies within a type are often perceived as more homogeneous. Furthermore, although some health insurers operate on a national level, the majority of health insurers operate regionally. Access to some company-based and guild-based health insurers is still restricted.⁴

Overall, the number of competing health insurers in Germany has dropped dramatically from 642 in 1996 to 262 in 2006. This reduction in the number of insurance companies is mainly the result of several mergers between companies. Hardly any insurer dropped out of the market without being involved in a merger. Although this consolidation of the market seems to be continuing and choice is somewhat less ample on the regional level, consumers can still choose between many health insurers. Thus, one requirement for managed competition to work is clearly fulfilled: consumers have free choice.

³ The threshold for mandatory social insurance is generally somewhat higher than the income ceiling used for the calculation of the individual premiums, so in 2004, persons with a monthly income of 3862.50 € were allowed to choose between private and social health insurance and the income ceiling was equal to 3487.50 € per month.

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

Table I. Contribution rates of health insurers (in %-points)

	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
Std. dev.	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

Table II. 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

DATA

Our study is based on an almost complete panel of individual health insurers that were active in the German social health insurance market between January 2001 and April 2004.⁵ 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 that specializes in analyzing the German health insurance market. The data have been validated by comparing them to information provided by several branch organizations of health insurers and by the Federal Ministry of Health. This marks the first time that an almost complete panel of individual health insurers has been constructed. Prior studies on price elasticities in Germany were only able to base their analysis on data aggregated over types of insurance companies or on information about individual companies that covered only a small part of the market (Schut *et al.*, 2003).

Table I shows descriptive statistics about the development of contribution rates of health insurers. The mean of the contribution rate increased during the whole period. The range between the lowest and the highest contribution rate is quite large. In 2004, switching from the company with the highest contribution rate to the one with the lowest generated a combined annual saving of 2300 € for an individual and his employer if salaries met the maximum income ceiling.

Table II 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%.

⁵Only a few, very small, company-based health insurers who dropped out of the market during 2001 are excluded, along with those that did not have open enrollment and whose members were not allowed to switch, e.g. a health insurer for miners and one insurer for sailors.

EMPIRICAL FRAMEWORK

The model

This study focuses on estimating demand parameters of choice among health insurers within the German social health insurance system. We assume that consumers choose health insurer $i \in N$ in period t , which maximizes their utility. Indirect utility is specified as a linear function of insurance-specific attributes x_{it} , γ_i and ε_{it} , which are equal for all individuals. In our case, the vector x_{it} only consists of the contribution rate, i.e. the price, and is observed by consumers and the researcher. In contrast, γ_i and ε_{it} are not observed by the researcher but known to consumers. Assuming that consumers' individual specific heterogeneity is i.i.d. Type I extreme-value distributed, this leads to the well-known conditional logit model (McFadden, 1973). The distributional assumption also allows for the straightforward estimation of the model parameters on the basis of company data that are available to us, i.e. market shares s_{it} .⁶

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

Taking logarithms leads to a convenient linear representation of the model,

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

where δ_t captures the logged denominator of Equation (1), which is a time-specific constant. By estimating a time fixed-effects model, any variation that takes place across periods and that affects all companies alike is removed. In estimating the model, γ_i and ε_{it} are, respectively, dealt with as insurance-specific effects and random error terms.

Since insurance contracts are generally enduring, consumers are not forced to decide on their health insurer each period. Many are likely to avoid transaction and information costs by simply staying with their current insurer without considering any alternative; others might feel uncertain about the characteristics of alternatives. 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

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

with $0 \leq \alpha \leq 1$ 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. Here, each individual decides on his or her health insurance each period. Transitory variation in the contribution rate has only short-term effects on market shares, not long-term ones. If, in contrast, $\alpha = 1$ holds, market shares are non-stationary, following a random walk. Hence, even if contribution rates remain stable, no equilibrium of market shares exists, and only the changes in market shares can be explained, but not their level. Consequently, transitory changes in the explanatory variables have permanent effects on market shares. Finally, in the intermediate case ($0 < \alpha < 1$), transitory variation in explanatory variables has long-lasting effects on market shares. However, these effects will eventually fade out.

Different measures for consumers' price sensitivity are available. 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

⁶For a formal approach on how to derive the aggregate model relating market shares to prices from a complete system of demand and supply functions for differentiated products, see Berry (1994); also compare Scanlon *et al.* (2002).

shares are equal to

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

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

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

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

Elasticities and substitution patterns in (4) and (5) are quite restrictive, because insurer-specific heterogeneity in price elasticities only enters through differences in market shares. This is a result of the i.i.d. assumption leading to Equation (1), which also implies the independence of irrelevant alternatives. Several, less restrictive, specifications have been proposed to relax some of the assumptions of the basic logit model (cf. Nevo, 2000). For example, Berry (1994) uses a nested logit model, Berry *et al.* (1995) implements a random coefficients logit model, and Chintagunta (2001) specifies a probit demand model. While the probit model requires computing probabilities over multiple integrals (this is only manageable for a small number of alternatives), Berry *et al.* (1995) rely on additional information on the distribution of consumer characteristics not available in our data, such as the mean income of consumers of each alternative. In contrast, the nested logit model as specified in Berry (1994) can be estimated using our data and allows for testing whether insurance companies are grouped. Nesting by type of insurer (AOK, BKK, etc.), however, is rejected,⁷ so we stick to the basic logit case.

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. 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).

In dynamic models, such as Equation (3), the right-hand side variable $\log(s_{it-1})$ is correlated with the composite error $(\gamma_i + \varepsilon_{it})$, which leads to inconsistent estimates, even if individual heterogeneity is accounted for by either fixed- or random-effects (see, e.g. Baltagi, 2001). Yet, under the assumption of serially uncorrelated errors ε_{it} , a GMM estimator (see Arellano and Bond (1991) for details) allows for consistent estimation. This estimator is based on first-differencing the regression equation and uses lagged dependent variables and – potentially – past, present and future values of x_{it} as instruments for $\Delta \log(s_{it-1})$. For GMM, the number of valid instruments varies according to whether the explanatory variables x_{it} are exogenous, predetermined or endogenous with respect to the error term ε_{it} . That is, whether $\text{cov}(x_{it}, \varepsilon_{it}) = 0$ holds for any t and τ , only for $\tau \geq t$, or just for $\tau > t$, respectively.

⁷This general model was estimated including insurance type-specific time dummies. Nesting as specified in Berry (1994) would imply numerous restrictions on them. Yet, these restrictions are rejected in every specification of the model.

Fortunately, as long as the model is over-identified, Sargan tests (Sargan, 1958; Hansen, 1982) 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.⁸ In this analysis, estimation of GMM is implemented using an efficient two-step procedure. In order to obtain valid standard errors, a corrected variance estimator proposed by Windmeijer (2005) is employed.

All the variants of GMM estimators that we have discussed so far 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 (see Arellano and Bover, 1995) 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_{it} , and the first-differenced exogenous variables, Δx_{it} , have to be uncorrelated with the individual specific effect γ_i (see also Blundell and Bond, 1998). Once again, these moment conditions can be tested on the basis of (difference) Sargan tests.

The above discussion makes it obvious that a rather large number of different GMM specifications can potentially be used for estimating the model. The following section presents estimation results for several specifications and additionally discusses our strategy for selecting the most appropriate one.⁹

Unfortunately, our panel data is unequally spaced. This could lead to inconsistent estimates, both in the GMM case and in the case with $\alpha = 1$, if one insists that one period in the theoretical model has to coincide perfectly with a certain time span in the empirical data.¹⁰ 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. Because GMM is quite data consuming, we precede using all data available, i.e. the 7 unequally spaced waves. Yet, in order to account for the above problem, we also check our results by comparing them to those obtained from a reduced sample that only includes waves 2–6 (see Appendix A). These waves are equally spaced at semi-annual intervals.

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 in which the market share of each individual company is determined by company-specific individual effects and current contribution rates. Then, we consider a dynamic model in which the market share follows a stationary first-order autoregressive process, i.e. market shares are persistent, yet differences in contribution rates can lead to long-lasting changes in the shares that fade out after some time. Finally, we provide results for a model in which 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.

⁸In the rest of the paper, we will 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.

⁹Estimation was done using David Roodman's Stata command `xtabond2` (Roodman, 2005).

¹⁰This assumption is likely to be regularly violated in survey data, since the date of interview will vary for practical reasons.

Static model

In the 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 that influence consumers in their choice between companies and that 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, and any additional medical treatments that, although not compulsorily covered by the standard benefit package, are nevertheless covered by some of the companies.

Contrary to models analyzing individual consumer data, the contribution rate might not be exogenous at the level of company data, since price-setting insurers observe those insurance-specific effects that are unobservable to the researcher. Thus, we instrument contribution rates by their one-period lag and use a test for endogeneity as reported in Wooldridge (2002, pp.118–122). Equation (2) is estimated using the standard fixed-effects model. As can be seen in Table III, the contribution rate is insignificant if the contribution rate is not instrumented. Yet, in the instrumental variables model, it has a (nearly significant) negative effect on the market share of an insurer. Exogeneity of the contribution rate, however, is rejected at the 5% level but not at the 10% level.

Having said this, estimation results presented in subsequent sections strongly argue against the static model. It has to be regarded as misspecified rendering any results from the static model biased. Hence, any conclusions presented in the remainder of the paper rest on results obtained from dynamic model specifications rather than static ones.

Dynamic models

Generalized method of moments (GMM). The dynamic model is equivalent to a world in which 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.

The first two specifications in Table IV are based on an Arellano–Bond-type first-differenced GMM estimator. In column 1, we present the results for a specification in which 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 in which the matrix of possible instruments has been reduced to a minimum ($\Delta \log(s_{it-1})$ and Δx_{it} are instrumented by only one variable each), indicates that the additional restrictions are not valid, since the difference-Sargan test is significant ($\chi^2_{(25)} = 45.76$).

Table III. Fixed-effects estimates for static model

	Fixed-effects model		IV fixed-effects model	
	Coefficient	Std. error	Coefficient	Std. error
Contribution rate	−0.0045	0.0274	−0.1885*	0.0961
Within- R^2	0.1546			
F-/Wald Test	12.32***		2.01e+06***	
Observations	1960		1589	
Test for endogeneity (t -statistic)			1.77*	

Note: Regression includes time dummies for each wave. Huber–White robust standard errors given. *** indicates significance at the 1% level, ** at 5%, * at 10%.

Table IV. 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)	

Note: 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 and 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 the 1% level, ** at 5%, * at 10%.

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.

The Arellano–Bover-type system GMM estimator includes additional moment conditions, and therefore allows the identification of the model even if α is close to unity. The estimates are also provided in Table IV. We see that in all of the system GMM models, the coefficient of the contribution rate is significant and of a much higher magnitude than in the first-differenced GMM models. Once more, we start with a specification in which 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 is reduced (not reported), or 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 not valid (i.e. are rejected 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 prior to 1996, it is quite plausible that these market constraints led to strong deviations from equilibrium under market conditions that had not been neutralized until 2001, the beginning of our data sample. Hence, system GMM seems to rely on inappropriate assumptions in the case analyzed here.

Summing up, the Sargan tests tend to favor the first-differenced GMM specification that includes the contribution rate as an endogenous regressor (GMM2), although it is close to a unit-root process and, hence, poor precision of the estimates. Still, all variants of the GMM model strongly argue in favor of market shares being highly persistent, rendering any static specification inappropriate.

Model in first differences. In this subsection, we provide the results for a model that explains first differences of market shares $\Delta \log(s_{it})$, rather than levels; i.e. the restriction $\alpha = 1$ is imposed on

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

	OLS 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	−0.0953***	0.0092	−0.1642***	0.0497
Observations	1589		1589		1589		1589	
R ²	0.1493				0.1486			
F-Test	20.36***		10.59***		21.41***			
Test for endogeneity (<i>t</i> -statistic)					2.00**		1.61	

Note: Regression includes time dummies for each wave. Huber–White robust standard errors given. *** indicates significance at the 1% level, ** at 5%, * at 10%.

Equation (3). This refers to a world in which transitory differences in contribution rates lead to permanent changes in market shares. Once a consumer changes his/her insurance company, he/she will stay with the new company as long as no further differences in contribution rates prevail or any unsystematic effects occur. There are no insurance companies that are big just because they started big, and there are no companies that are small just 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, 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 V. The contribution rate is highly significant and the magnitude is comparable in both models.

When the model is estimated without fixed-effects, the coefficient of the contribution rate only marginally differs if the contribution rate is either assumed to be exogenous (UR1) or instrumented by its one-period lag (UR3). Yet, exogeneity of the contribution rate is rejected. In contrast, the coefficient of the contribution rate almost doubles in the case with fixed-effects when using the IV estimator (UR4). There, however, exogeneity is not rejected. Hence, we regard UR2 as our preferred specification in the class of the unit-root models and focus further discussion on UR2.

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 is ill suited for testing the null-hypothesis of a unit-root being present, since the model is not identified under the null hypothesis. For this reason, additional panel unit-root tests are carried out. Two distinct regression-based test procedures are considered that are both well suited for panels with a large number of cross-sectional units and 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 and Meyer, 1994) specifies the regression in terms of deviations from initial conditions and is therefore 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.*, 2005).

The two tests yield different results, as Table VI shows. 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

Table VI. Panel unit-root tests for market shares

	Simple OLS	Breitung–Meyer
Coefficient lagged value	0.9844	1.2094
<i>t</i> -statistic	−6.2956	5.5407
<i>p</i> -value (one sided <i>t</i> -test)	0.000	1.000

Note: Regression includes time dummies for each wave. Heteroskedasticity-robust standard errors given. Regressions with/without uniform constant yield similar results.

answer the question whether the model must be formulated in terms of first differences of market shares or as a more general dynamic one.¹¹

Taking together the information from these panel unit-root tests and the estimated results from both our GMM specification (GMM2) and the specification of first-differenced market shares with fixed-effects (UR2), we conclude that both are quite similar. Although 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, although GMM2 results are less easily interpreted because of substantially larger standard errors. Thus, we go on presenting further results for both specifications. There is a slight preference for UR2 if we take into account the results based on waves 2–6, only, i.e. the equally spaced panel (see Appendix A).

Model extensions

In order to gauge 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 enrollment in companies, and (iv) regional restrictions of insurance companies, we estimate several models in which 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 GMM2 and UR2 and are available from the authors upon request. Results show that none of these interactions are significant.

Elasticities

After having discussed several econometric specifications for estimating the effect of price on market shares, we now present the premium elasticities of market shares implied by the estimated coefficients. The point estimates of the elasticities and the corresponding 95% confidence intervals for GMM2 and UR2 are given in Table VII. The short-run premium elasticity is based on Equation (4) and calculated for the sample mean.

The elasticity is significant for the unit-root specification with fixed-effects (UR2). It displays more robust results and substantially smaller standard errors than competing dynamic specifications estimated using GMM. It is also favored over the static model. Finally, estimates presented in the preceding subsections do not argue in favor of including interaction terms of the premium rate and other explanatory variables in the model or any nesting by insurance type as specified in Berry (1994).

At sample mean, UR2 displays a short-term premium elasticity of about minus one; GMM2, of about minus one-half. The long-run elasticity of GMM2 is approximately minus twelve. In the unit-root case, it approaches infinity, by construction. This indicates a distinct sensitivity of consumers to differences in contribution rates. This fulfills one of the preconditions necessary for managed competition to work.

¹¹ In contrast, the series of contribution rates is unambiguously stationary.

Table VII. Estimates of short-run premium elasticity

	GMM2		UR2	
Mean premium elasticity		−0.55		−1.09
95% confidence interval	−1.74	+0.64	−1.43	−0.75

Note: Elasticity estimated for sample mean. Estimates based on results from Tables IV and V.

CONCLUSIONS

This paper analyzes an important issue that advances insights into the dynamics of the German social health insurance market. Results indicate that consumers are sensitive to price differences, which might have severe consequences for health insurers charging higher premiums than their competitors. The analysis is based on two novel elements. First, it is based on a unique panel data set that covers the social health insurance market on the level of individual insurance companies. Prior to this study, only data aggregated over insurers or with very few individual insurers were available. Second, this paper uses an advanced econometric technique that 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 that uses the level of premiums to explain changes in market shares or, if specified in levels, displays high persistence. For this specification, we obtain a short-run premium elasticity of market shares of minus one-half to 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. Interestingly, our results are much closer to those obtained for other countries like Switzerland (Beck, 2004). From the point of view of economic theory, the estimated short-run price sensitivity appears to be rather small. In theory, the price elasticity should approach infinity, because consumers can choose between products that are almost perfect substitutes from an objective perspective. But since the estimated elasticity is relatively small, one might hypothesize that most consumers do not treat health insurers as perfect substitutes.

Another reason for the small short-run price sensitivity estimate could be that 55% of the respondents in a recent survey stated that their health insurer gave them a feeling of security and reliability. Besides, the costs incurred by switching companies were considered to be very high, and information about the differences between health insurers was perceived to be poor (Höppner *et al.*, 2004). One instrument to enhance transparency for consumers and, thus, improve competition might for instance be a standardized reporting system.

In contrast to earlier analyses, our results are based on a dynamic specification. They 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, permanent relative changes in contribution rates will have dramatic effects on the market shares of health insurers in the long-run. Insurers who 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.

Overall, due to legal restrictions, health insurance companies within the German social health insurance system have only limited opportunities to seek for cheaper health care provision. The legislator therefore does not fully exploit the opportunity to realize efficiency gains that is offered by a considerable price sensitivity of consumers: If the legislator gave more room for managed competition, for instance by making selective contracting with health care providers easier for the insurers, efficiency within the system could potentially be improved, because price differences would then to a higher extent reflect differences in efficiency. International experience shows, however, that most additional instruments for managed competition can also be used for risk selection (van de Ven and Ellis, 2000). At present, risk adjustment between German health insurers is rather poor, leaving ample room for risk selection (Behrend *et al.*, 2004). Among others, better risk adjustment would therefore be necessary if more instruments of managed competition were implemented.

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APPENDIX A: EQUALLY SPACED TIME PERIODS

As mentioned under section 'Estimation', 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 2–6 (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 orthogonality 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 enrollment. Estimation results based on the reduced sample consisting of regularly spaced waves, therefore, do not challenge the main findings of our analysis.

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