HEDG

HEALTH, ECONOMETRICS AND DATA GROUP

WP 18/25

Do hospitals respond to decreasing prices by supplying more services?

Martin Salm and Ansgar Wübker

August 2018

http://www.york.ac.uk/economics/postgrad/herc/hedg/wps/

Do hospitals respond to decreasing prices by supplying more services?

Martin Salm[†] and Ansgar Wübker[‡]

Abstract

Regulated prices are common in markets for medical care. We estimate the effect of changes in regulated reimbursement prices on volume of hospital care based on a reform of hospital financing in Germany. Uniquely, this reform changed the overall level of reimbursement – with increasing prices for some hospitals and decreasing prices for others – without affecting the relative prices for different groups of patients or types of treatment. Based on administrative data, we find that hospitals react to decreasing prices by supplying more services. We interpret our findings as evidence for a negative income effect of lower prices on higher volume of care.

Keywords: government expenditures and health, hospital care, procurement

JEL Codes: I11, L10, L21

[*] The authors thank Boris Augurzky, Padmaja Ayyagari, Jan Boone, Mary Deely, Tobias Klein, Peter Kooreman, Ellen Meara, Robert Nuscheler, Jonas Schreyögg, and Jonathan Skinner for their valuable comments and suggestions as well as participants at the conference of the American Society of Health Economists in Los Angeles, the workshop on health economics in Heidelberg, the workshop on health econometrics in Wuppertal, the conference of the German Association for Health Economics in Bielefeld, the meeting of young micro-econometricians in Mannheim, the workshop on applied microeconomics in Odense, and seminar participants at the Universities of Bochum, Duisburg-Essen, Erlangen-Nuremberg, Groningen, and Tilburg. We thank the German Statistical Office for allowing access to data and Adam Pilny for making data on hospital mergers available to us. The financial support of the Fritz Thyssen Foundation is gratefully acknowledged.

[†] CentER, Tilburg University. e-mail: m.salm@tilburguniversity.edu
[‡] RWI Essen. e-mail: ansgar.wuebker@rwi-essen.de

1. Introduction

In many important markets prices are set not by the interaction of demand and supply, but by public or semi-public agencies. Regulated prices are especially common in markets for medical care, for example for the reimbursement of hospital services in Germany and for Medicare patients in the United States. This makes it an interesting and policy relevant question to examine how providers of medical care respond to changes in regulated prices. In this study we examine the effect of changes in reimbursement prices on the volume of hospital care in Germany.

The direction of this effect is not a priori clear, neither based on economic theory nor based on previous empirical evidence. In a seminal study, McGuire and Pauly (1991) develop a theoretical framework for how medical providers respond to changes in regulated prices. They show that the effect of changes in regulated prices on volume of care is a combination of two effects: 1) the response to a change in relative prices between different types of services or groups of patients, and 2) the response to a change in the overall price level. The second effect is typically referred to as "income effect" in the literature (McGuire 2000). While an increase in relative prices for a type of service is predicted to lead to an increase in the supply of this service, alternative economic models make different predictions about the sign of the income effect. According to standard economic theory, higher prices lead to more supply of services. In contrast, models of supplier-induced demand (Evans 1974, Gruber and Owings 1996) or of providers aiming at target incomes (Rizzo and Zeckhauser 2003) predict that higher prices can lead to the supply of fewer services, and lower prices can lead to the supply of more services. The underlying intuition is that providers compensate for the effect of lower prices on income by providing more services. In line with this reasoning the Federal budgeting process in the United States assumes that a one-percent decrease in Medicare reimbursement prices increases treatment volumes by around 0.3-0.5 percent (Congressional Budget Office 2007).

The empirical evidence on the effect of lower reimbursement prices on volume of care is mixed. Among recent studies, some find that lower prices decrease volumes of care (Clemens and Gottlieb 2014, He and Mellor 2012, Januleviciute et al. 2016), while other studies find that lower prices increase volumes of care (Heaton and Helland 2009, Shigeoka and Fushimi 2014). Most previous studies examine price changes that affect only a subset of patients or medical treatments, such as Medicare beneficiaries (Rice 1983, Yip 1998, He and Mellor 2012, Clemens and Gottlieb 2014), automobile accident victims (Heaton and Helland 2009), at-risk newborns (Shigeoka and Fushimi 2014), or patients with specific diagnoses (Januleviciute et al. 2016), while leaving prices for other patients and treatments unaffected. In these settings the effect of price changes is then a combination of the effect of a change in relative prices and an income effect, and it is very difficult to ascertain which part of the overall effect can be attributed to either of them.

In our study, we exploit a setting which makes it possible to estimate the income effect directly and without having to disentangle the income effect from the effect of a change in relative prices. Specifically, we look at changes in hospital-specific base rate factors (Basisfallwerte) in Germany between the years 2004 and 2009. Changes in base rate factors shift the overall level of reimbursement prices in hospitals without affecting relative prices for different groups of patients or types of services. In the year 2004, base rate factors varied widely between hospitals based on historical costs. Between 2004 and 2009, base rate factors gradually converged toward the average base rate factor at the state level. Thus, base rate factors increased for some hospitals and decreased for others.

In our empirical analysis we exploit this variation in reimbursement prices. We estimate the effect of changes in base rate factors between 2004 and 2009 on corresponding changes in volumes of care. We use a differences-in-differences estimation approach with the change in prices as continuous treatment variable.

Based on administrative data for a 70 percent random sample of all German hospitals provided by the German Statistical Office we find an elasticity of prices on the number of hospital admissions of -0.14 and an elasticity of prices on the case-mix index – a measure of the average payment per patient admitted – of -0.29 over a five-year period. Thus, hospitals respond to increasing prices by decreasing service supply, and to decreasing prices by increasing service supply.

Our empirical results are robust to controlling for changes in prices of competing hospitals as well as for regional demographic and economic trends. In robustness checks, we find that our results cannot be explained by pre-existing trends in volume growth or by heterogeneous effect of the introduction of DRG based payment. They can also not be explained by mergers or changes in ownership type, or by differences in initial capacity utilization. Furthermore, our results cannot be attributed to demand-side reactions to price changes, since prices faced by patients are not affected by changes in base rate factors. Moreover, we find no evidence for asymmetric effects of price increases and price decreases.

In addition, we examine whether the effects of price changes on volume of care differ among hospitals. We find a somewhat weaker effect for public hospitals than for private and not-for-profit hospitals, and for large hospitals than for small hospitals. However, these differences are not statistically significant. We also examine the effects of higher prices on lower volumes of care for types of treatment with a high degree of regional variation, and we find that effects are especially large for cataract surgery and for tonsillectomy.

We interpret our findings as evidence for a negative income effect of lower prices on higher volumes of hospital care. Our findings provide an interesting insight into the objective function of hospitals. Based on economic theory the sign of the income effect is ambiguous, and it is likely to be negative if the objective function of hospitals has a high degree of relative risk aversion with respect to lower revenues (see Section 3). During our study period the financial situation of many German hospitals was precarious. In the year 2008, 27% of hospitals faced an enhanced risk of bankruptcy (Augurzky et al. 2010). Many hospitals were struggling to cover their fixed costs which under German labor law cannot be adjusted quickly or cheaply. In order to avoid budget cuts, layoffs, or even closure, hospital management and staff might put in extra efforts to generate additional revenues. Thus, they responded to lower reimbursement prices by providing more services.

Evidence for a negative income effect is not the same as evidence for physician induced demand. Physician induced demand is defined as care that is in the financial interest of the provider, but not in the best interest of the patient (McGuire 2000). In our study, we do not know the optimal volume of care, and we are unable to say whether or not the additional care provided in response to lower reimbursement prices was in the best interest of patients.

Still, our findings have important policy implications. First, our results suggest that existing rules to limit volume of care for German hospitals do not prevent hospitals from increasing treatment volumes in response to lower prices. Second, the negative income effect has implications for the budgetary impact of changes in regulatory prices for hospitals. Following the example of the

Congressional Budget Office in the United States policymakers in Germany should take the income effect into account in their decisions about setting prices in the hospital sector.

Our study continues as follows. Section 2 describes the institutional setting of hospital financing in Germany. Section 3 presents a stylized model of hospitals' response to price changes. Section 4 discusses the empirical strategy. The data are described in Section 5, and our results are presented in Section 6. Section 7 concludes.

2. Institutional setting

Hospital financing in Germany comes from several sources. By far the most important sources are public and private health insurers, which cover around 88.5 percent of all hospital expenditures (Simon 2010).[1] Funding from these sources is largely used to cover hospitals' operating costs, including payments for physicians' services. In Germany, physicians are usually employees of the hospital where they work, and they receive a salary from the hospital. Importantly, payment rates for hospital care do not differ between publicly and privately insured patients.[2] The remaining hospital revenues are derived mainly from state governments, which are responsible for long-term infrastructure investments. Patient co-payments in Germany are small relative to hospital costs. Patients have to pay a fixed charge of €10 per night of their hospital stay as a contribution toward room and board, and there are surcharges for additional services such as a single room or treatment by the hospital director (For details of hospital financing in Germany see Quentin et al. 2010 and Simon 2010).

[1] The remaining 11.5 percent is covered by private households (2.3 percent), employers (3.4 percent), public accident insurance (1.2 percent) and the federal states (4.6 percent). All numbers refer to 2007 and are provided by Simon (2010).
[2] By contrast, payment rates for outpatient care differ between privately and publicly insured patients.

Before 2004, hospital payment for operating costs was based on a mixed payment system (see description by Busse and Riesberg 2004). 75% of cases were reimbursed by per-diem charges. Per-diem charges reflected the cost structure of hospitals, and they varied between hospitals.[3] Hospitals with higher expenditures in relation to the number of hospital beds received higher per-diem charges. The remaining 25% were already reimbursed according to patients' diagnosis-related groups (DRGs).[4] Before 2004, payment based on diagnosis related groups was restricted to certain diagnosis such as for example childbirth or coronary heart diseases. In 2004, payment based on DRGs was expanded to almost all diagnoses. The aim of this reform was to make hospital payment more transparent and promote efficiency and competition. The German DRG payment system is similar to hospital payment systems in other countries that have introduced DRG-type systems, starting from the early 1980s. A particular aspect of the German reform that sets it apart from DRG-payment introduction in other countries is that payment changes were introduced gradually. During a first "budget-neutral phase" in 2004, hospitals were reimbursed according to DRGs but prices were adjusted with hospital-specific base rate factors in such a way that hospitals could still achieve their historical budgets. During the "convergence phase," which lasted from 2005 until 2009, hospital-specific prices gradually converged toward average prices at the state level.

Under the German DRG system, payment for a hospital admission is based on the following formula: [5]

[3] Differences in per diem charges across hospitals to some extent also reflected historical differences in their reimbursement negotiations (Friedrich et al. 2008).
[4] This 25% also included some fee for service payments. However, they were restricted to a number of complex surgical procedures.
[5] This formula abstracts from adjustment factors for teaching hospitals etc. During our study period, DRG payment covered most but not all treatments with psychiatric treatments as the main exception.

$$payment_{i,j,t} = drg_{j,t} * baserate_{i,t} \quad (1)$$

Payment is the product of two factors: $drg_{j,t}$ is the cost-weight factor for DRG *j* in year *t*, while $baserate_{i,t}$ refers to a hospital-specific base rate factor for hospital *i* in year *t*. All discharged hospital patients are assigned to a DRG. This assignment is based mainly on diagnoses but in some instances it is also based on procedures and patient characteristics such as age, sex, and birthweight (for newborns). The German DRG system was modeled on the Australian DRG system and initially had 664 DRGs. DRG cost-weight factors are the same for all hospitals. They are set at the national level jointly by representatives of health insurers and hospitals, and they are adjusted annually based on detailed patient-level cost data from a sample of hospitals. The cost-weight factors are normalized such that the average cost-weight factor is set to one. Cost-weight factors are much higher than one for cost-intensive DRGs such as a liver transplant, and they are lower than one for less cost-intensive DRGs such as an ordinary hand fracture.

Hospital-specific base rate factors reflect historical budgets before the introduction of DRG payment. During the budget-neutral phase of the reform, hospital-specific base rate factors were computed by dividing pre-reform budgets by the sum of the cost-weight factors hospitals would have earned for their pre-reform services based on post-reform cost-weight factors. Using hospital-specific base rate factors ensured that hospitals could still achieve their historical budgets under DRG payments in the early stage of the reform as long as they continued to provide the same volume and type of services.

During the 2004–2009 convergence phase, hospital-specific base rate factors gradually converged toward state averages. Base rate factors gradually decreased for hospitals with above-average base rate factors, and they increased for hospitals with below-average base rate factors. The convergence process is illustrated in Figure 1. From 2009, hospitals in the same state received the same base rate factor. The convergence process put high-cost hospitals under significant pressure to lower costs. In order to protect hospitals from excessive budget cuts, annual reductions in total hospital budgets were limited, for example in 2008 to not more than 2.5 percent.

The distribution of hospital-specific base rate factors, i.e. the empirical version of Figure 1 for quantiles of initial prices is shown in Figure 2. The initial variation in base rate factors was substantial. In 2004, the difference between the 10th and 90th percentiles of base rate factors was around 36 percent. In 2009, base rate factors were equalized at the state level. Remaining differences at this stage reflected differences in base rate factors between states. The convergence of base rate factors implied substantial increases in across-the-board reimbursement prices for some hospitals and substantial reductions for others. Base rate factors at the 10th percentile increased by 15.4 percent in real terms between 2004 and 2009, while those at the 90th percentile decreased by 11.8 percent.

The German DRG rules make provisions to protect against induced demand. Hospitals may keep only 35 percent of additional revenues if they exceed the number of target admissions. Additional revenues that are generated by up-coding, i.e. charging a more expensive DRG for the same treatment, are meant to be reclaimed fully by health insurers (Tuschen et al. 2005). However, these provisions are not applied consistently in practice. Hospitals routinely delay budget negotiations until late in the year, and they then negotiate target numbers of admissions that are close to the actual number of admissions (Kumar and Schönstein 2013). Furthermore, increases in the case-mix index, which is the average cost-weight factor for patients in a hospital, can be reimbursed if the hospital can provide good medical reasons for more intensive treatment.

3. Hospitals' response to price changes

In the following we present and discuss a theoretical framework for how medical providers respond to price changes. This framework is a slightly modified version of a classical model developed by McGuire and Pauly (1991).[6] According to this framework, medical providers can influence the demand for their services due to their superior knowledge about patients' healthcare needs, and they can use this influence in their own financial interest. However, providers derive disutility from inducing demand, and they weight the disutility from inducing demand against the additional utility from higher revenues. The objective function of a hospital can then be characterized as:

$$\max U = U(Y, I) \tag{2}$$

where $Y = (P - MC)Q(I, X)$

A hospital maximizes the objective function U, which is an additively separable function of income Y and demand inducement I. We assume: $U_Y > 0; U_I < 0; U_{YY} < 0; U_{II} < 0$. Quantity of treatment Q is an additively separable function of demand inducement I and demand shifters X. We assume: $Q_I > 0; Q_{II} = 0$. P is the price the provider receives for a unit of treatment and MC is the marginal cost for a unit of treatment. We further assume that $P \geq MC$.[7]

The hospital chooses the level of demand inducement I to maximize the objective function. The first order condition is given by:

$$U_Y Q_I (P - MC) + U_I = 0 \tag{3}$$

[6] The model by McGuire and Pauly allows for two different prices for the same service for different groups of patients. Since prices in Germany do not vary between patients we consider the case of only one price.
[7] If $P < MC$ then the provider chooses $Q=0$.

Taking the derivative of (3) with respect to *P*, we can solve for the effect of a change in price on demand inducement *I*, and we obtain:

$$I_P = \frac{-U_{YY} Q Q_I (P - MC) - U_Y Q_I}{U_{YY} (Q_I (P - MC))^2 + U_{II}} \tag{4}$$

In general, the sign of (4) cannot be determined a priori, and it cannot be said whether a higher price leads to lower or higher demand inducement. The sign of the denominator is negative, but the numerator consists of a positive term $-U_{YY} Q Q_I (P - MC)$ and a negative term $-U_Y Q_I$.

Equation (4) determines the sign of the income effect and the slope of the supply curve. If equation (4) is positive then we obtain the standard result of a positive income effect and an upward sloping supply curve. If on the other hand equation (4) is negative then we receive the counterintuitive result that the supply curve is backward bending, and that there is a negative income effect. Equation (4) is negative if $-U_{YY} Y / U_Y > 1$, where $-U_{YY} Y / U_Y$ is the definition of the coefficient of relative risk aversion, a measure of the curvature of the income utility schedule. Thus, the income effect is negative if medical providers are very risk averse with respect to declining revenues and the income-utility schedule has a high curvature.

The above model was developed for the case of a single decision maker such as a self-employed physician in an individual practice. In contrast, hospitals are large organizations where quantity of treatment is the result of decisions taken by many individuals, including hospital management and (medical) staff. In Germany, physicians who work at hospitals are typically salaried employees who report to the hospital management. They share in the success of a hospital through bonuses and better working conditions. Performance related remuneration is an important part of physician payment in German hospitals, especially for medical directors (Thurn 2013).

In the model of McGuire and Pauly (1991) *I* measures demand inducement which reduces the utility of physicians. However, in our context *I* does not need to refer to treatment that is not in the best interest of patients. *I* can be seen as the utility cost of extra effort in order to increase the quantity of treatment. This leads to the question of how German hospitals can increase the demand for their services. One of the most important driving factors behind patients' hospital choices are recommendations from outpatient physicians (Salfeld et al. 2009). Thus, it is important for hospitals to cultivate good relationships with outpatient physicians who are able to refer patients to the hospital. For example, hospital directors can visit outpatient physicians and inform them about new treatment techniques available at the hospital. Reportedly, many hospitals also pay outpatient physicians for patient referrals (GKV-Spitzenverband 2012). In addition to increasing the number of patient admissions, hospitals can also aim to increase payments per patient admitted.[8]

The objective function can differ between hospitals. For example, risk aversion with respect to declining revenues could be lower for public hospitals than for private or not-for-profit hospitals. The reason is that a loss-making public hospital might have a higher chance of receiving a government bailout than a loss-making private or not-for-profit hospital. The size of the income effect could also be larger for small than for large hospitals. The relative impact of any individual decision-maker on the financial situation of a hospital will be larger in a small hospital than in a large hospital. Thus, collective efforts may be easier to organize in a small organization than in a large organization.

[8] For evidence of up-coding by hospitals see, for example, Dafny 2005 and Jürges and Köberlein 2013.

Demand inducement can also vary by type of patient admission. For example, there could be more room for discretionary decisions by physicians for treatments with a high degree of regional variation than for treatments with a low degree of regional variation. In equation (4) such differences can be reflected in different values of Q_I , the increase in quantity of care that can be generated by one unit of additional effort *I*.

4. Empirical approach

In our empirical strategy we examine how treatment volumes respond to changes in base rate factors. We use a differences-in-differences regression approach. However, instead of looking at a binary treatment variable, we examine the effect of a change in prices, which is a continuous treatment variable. Thus, we compare not just two groups with different treatments, i.e. one treatment group and one control group, but we look at a continuous range of treatments and compare different treatments with each other. We estimate linear regression models with two periods and hospital-specific fixed effects:

$$q_{it} = \beta \Delta price_{i,2004-2009} \mu_{2009} + \gamma \Delta price\_comp_{i,2004-2009} \mu_{2009} + X_{it}{}'\delta + \mu_{2009} + \alpha_i + \varepsilon_{it} \qquad (5)$$

where q_{it} is the treatment volume for hospital $i \in (1...N)$ in year $t \in (2004, 2009)$, $\Delta price_{2004-2009}$ is the change in the base rate factor of hospital *i* between the years 2004 and 2009, $\Delta price\_comp_{i,2004-2009}$ is the change in the average base rate factor for competing hospitals,[9] X_{it} includes regional demographic and economic characteristics,[10] μ_{2009} is a binary indicator for

[9] We define competing hospitals as hospitals that attract patients from the same geographical area. In Section 5 we describe how the variable for the average base rate factor for competing hospitals is constructed.
[10] Regional variables include the average age of men, the average age of women, population density, and the local unemployment rate.

the year 2009, α_i are unobserved hospital fixed effects, and ε_{it} represents unobserved time-varying hospital characteristics. β and γ are parameters, and δ is a vector of parameters. β is the parameter of interest, and it represents the effect of changes in reimbursement prices on changes in hospital volumes.

Our treatment variable is the change in base rate factors between the years 2004 and 2009. For the intermediate years 2005–2008 base rate factors did not always follow the adjustment schedule shown in Figure 1, but they were negotiated annually between sickness funds and hospitals. Thus, hospitals could have been able to influence base rate factors in those years. This concern does not apply to the year 2009, when base rate factors were equalized at the state level.

Regression equation (4) provides a consistent estimator of β if the exogeneity assumption below holds:

$$E[\varepsilon_{it} \mid \Delta price_{i,2004-2009}, \Delta price\_comp_{i,2004-2009}, \mathrm{X}_{it}, \mu_{2009}] = 0 \qquad (6)$$

Note that the equation above does not contain any assumptions about time-invariant unobserved characteristics α_i. Hospitals with different values of $\Delta price_{i,2004-2009}$ can differ in their observed and unobserved characteristics.[11] This does not violate the exogeneity assumption as long as changes in base rate factors are not correlated with changes in time-varying unobserved hospital characteristics ε_{it}.

The exogeneity assumption in equation (5) is akin to the common trend assumption in a differences-in-differences estimation framework. This assumption requires that in the absence of

[11] In Table A1 we show that hospitals with higher price changes are more likely to be small and rural.

price changes the number of admissions for hospitals with different values of $\Delta price_{i,2004-2009}$ should follow the same time trend. In the following paragraphs we discuss whether the exogeneity assumption is plausible in the context of our study. Specifically, we discuss potential violations of the exogeneity assumption and how we can test for these violations.

A first potential violation of the exogeneity assumption may arise if unobserved underlying trends in hospital volumes are correlated with $\Delta price_{i,2004-2009}$. We can test for this violation by examining whether trends in hospital volumes before the year 2004 are related to subsequent changes in base rate factors between the years 2004 and 2009.

A second potential violation could be related to the introduction of DRG based payments. Price changes coincide with the time period just after the introduction of DRG payment in the year 2004. In our empirical model we take time effects into account, but the effects of the introduction of DRG payment could be heterogeneous. For example, it could vary between small and large hospitals or between urban and rural hospitals. As a robustness check we estimate a model that allows for heterogeneous trends based on hospital and regional characteristics, and we test whether allowing for heterogeneous trends changes our estimation results.

A third potential violation of the exogeneity assumption could be related to non-linear effects of price changes on volume of care. For example, the effect of an increase in price and a decrease in price do not need to be symmetric. We can test for non-linear effects of price changes by computing the change in volume during the 2004-2009 period for different ranges of $\Delta price_{i,2004-2009}$, and assess whether these changes follow a linear pattern.

A fourth potential violation of the exogeneity assumption may be caused by changes in ownership type, the spread of private hospital chains, or by mergers of hospitals. As a robustness check we test whether our results change if we exclude hospitals from our sample who were affected by a change in the type of ownership between the years 2004 and 2009 or who were situated in the states of Hesse or Bavaria where private hospital chains were most active during our study period. We also test if being part of a merger is correlated with $\Delta price_{i,2004-2009}$.

One more alternative explanation for a negative β could be constraints on capacity utilization, meaning that hospitals with high initial capacity utilization are capacity constrained, and that they cannot increase volumes further. If high initial capacity utilization is correlated with increasing base rate factors, this could provide an alternative explanation for a negative β. We can test for this alternative explanation by examining the relationship between initial capacity utilization and initial base rate factors.

5. Data

Our main source of data is hospital statistics from the German Statistical Office for the period 2000–2009. These hospital statistics combine information about hospital characteristics such as ownership type and size with patient-level information on admissions, such as the main diagnosis and county of residence for each patient. These data are merged with county-level regional indicators from the German Statistical Office and with information on base rate factors provided by AOK, a group of health insurers.

Our study is based on a 70 percent random sample of all German hospitals. Our data include 1,159 hospitals with information on the number of admissions and base rate factors in the year 2004. Of those, 165 were excluded from the sample because they are not open year-round or they are day clinics or psychiatric hospitals which did not adopt the DRG system. A further 193 hospitals were excluded because they could not be tracked up to 2009. While hospital closures were very rare during our study period, mergers were quite common.[12] Our baseline estimation sample consists of 801 hospitals.

The outcome variable in the baseline specification is the natural logarithm of the total number of annual hospital admissions. In alternative specifications outcome variables are the natural logarithm of the total number of annual hospital admissions for specific diagnoses classified according to ICD 9 codes and the case-mix index. The main explanatory variable of interest is the percentage change in the base rate factor (*Basisfallwert*) of a hospital between the years 2004 and 2009.[13]

We compute a variable for the change in average base rate factors for competing hospitals that attract patients from the same geographical area. This calculation consists of two steps: 1) We first compute the average base rate factor for competing hospitals in each county. This calculation is based on hospital market shares for residents of each county. 2) We then compute the average base rate factors for competing hospitals for each hospital. This calculation is based on the county shares of patients for each hospital, e.g. what share of a hospital's patients comes from a specific county.[14,15]

[12] While over 2004–2009 only 19 hospitals were closed, about 20 percent of all German hospitals were involved in mergers. These numbers are based on the RWI Krankenhauspanel (see Pilny 2014), an alternative data source with detailed information on the full sample of German hospitals, but no information on volume of care. We use this data for robustness checks in section 6.

[13] We measure Basisfallwerte by "vereinbarte Basisfallwerte".

[14] Both hospital market shares for county residents and county shares for hospital patients are computed based on shares in 2004 and kept constant across years.

[15] Since our data come from a 70 percent random sample of German hospitals, this calculation leads to a slightly noisy but unbiased measure of average base rate factors for hospitals that attract patients from the same area.

We further compute variables on demographic and economic indicators for hospital catchment areas. For this calculation we weight county-level indicators for the average age of men, average age of women, population density, and unemployment rate based on the county shares of patients in each hospital.

Summary statistics for hospitals in our data set are shown in Table 1. Between 2004 and 2009 the average number of admissions per hospital increased from 10,940 to 11,878. The average values of the case-mix index were very close to one.[16] Between 2004 and 2009, public hospitals as a share of the total decreased slightly from 40.8 percent to 39.1 percent, while the share of not-for-profit hospitals fell from 45.1 percent to 44.2 percent. The remaining hospitals were private. The Herfindahl index for market concentration increased somewhat over the study period. Regional indicators showed a decline in unemployment rates and an increase in the average age of men and women. Average population density did not change much.

Figure 3 shows the distribution of our main explanatory variable, the changes in base rate factors between the years 2004 and 2009. Changes varied from substantial decreases to substantial increases in prices. The assignment of price changes was not random. Price changes between the years 2004 and 2009 are correlated with hospital characteristics and county-level regional characteristics in the year 2004. For example, hospitals with a large number of beds and hospitals in regions with a high population density were more likely to face decreasing prices while hospitals

[16] I.e. the average payment for a patient was similar to the value of the base rate factor and amounted to around €3,000 in 2009.

in regions with a high market concentration were more likely to face increasing prices (see Table A1 in the online Appendix).[17]

6. Results

Baseline specification

Table 2 shows estimation results for the effect of changes in base rate factors on the number of admissions. Our estimate of β in the baseline specification (equation 5) in Column 1 is -0,14. This coefficient is significantly different from zero at the five-percent level. This implies, that a one percent increase in prices causes a decrease in the number of hospital admissions by 0.14 percent. Correspondingly, a one percent decrease in prices causes the number of admissions to increase by 0.14 percent.

Robustness checks

In Section 4 we have discussed a number of potential violations of the exogeneity assumption in equation (6), and how we can test for these violations. The first potential violation arises if changes in base rate factors are correlated with unobserved underlying trends in hospital volumes. We can test for this violation by examining whether trends in hospital volumes before the year 2004 are related to subsequent changes in base rate factors between the years 2004 and 2009 based on the following linear regression model:

$$q_{it} = \sum_{t=2000}^{2009} \beta_t \Delta price_{i,2004-2009} \mu_t + \gamma_t \Delta price\_comp_{i,2004-2009} \mu_t + X_{it}'\delta + \mu_t + \alpha_i + \varepsilon_{it} \qquad (7)$$

[17] In a robustness check, we show that allowing for different time trends for small and large hospitals and for hospitals in regions with high and low population density and market concentration indices has very little effect on our estimation results.

Figure 4 shows estimation coefficients of β_t and their 95 percent confidence intervals for each year $t \in (2000,..,2009)$. The coefficients of β_t for the years $t \in (2000,..,2003)$ are close to zero, and they are not statistically significant, neither individually nor jointly. Thus, we conclude that there are no significant pre-trends in the years before 2004.[18]

A second potential violation of the exogeneity assumption can arise because of heterogeneous effects of the introduction of DRG payment. As a robustness check we estimate a model that allows for time trends that vary with hospital and regional characteristics based on the regression equation below:

$$q_{it} = \beta \Delta price_{i,2004-2009}\mu_{2009} + \gamma \Delta price\_comp_{i,2004-2009}\mu_{2009}$$

$$+\lambda X_{i2004}\mu_{2009} + \kappa H_{i2004}\mu_{2009} + X_{it}'\delta + \mu_{2009} + \alpha_i + \varepsilon_{it} \tag{8}$$

where $t \in (2004, 2009)$, X_{i2004} and H_{i2004} refer to regional and hospital characteristics in the year 2004 at the beginning of the payment reform, and μ_{2009} is a binary indicator for the year 2009. In Column 4 of Table 2 we show estimation results for the regression model in equation (8). The coefficient of β is essentially unchanged compared with the baseline specification in Column 1. Thus, our estimation results are robust to the inclusion of heterogeneous time trends with respect to observed hospital and regional characteristics.

A third potential violation of the exogeneity assumption could be related to non-linear effects of price changes on volume of care. For example, the effects of price increases and price decreases need not be symmetric. In Figure 5, we examine whether the effects of price changes on changes

[18] The coefficients of β_t in the years $t \in (2005,..,2008)$ show the effect of prices on volume in years when the adjustment in base-rate factors was only partially completed. We exclude the year 2004 as reference category.

in volumes of care are non-linear. We divide the sample into twenty bins of equal size according to the total change in base rate factors over 2004-2009. Subsequently we compute changes in the number of admissions between 2004 and 2009 for each of the bins. If we add a regression line to connect the points in the graph then the slope of the regression line is similar to the coefficient of $\Delta price_{i,2004-2009}$ in the baseline specification (-0.17 vs. -0.14). The observation points are roughly symmetrically distributed around the regression line. Thus, we find no evidence for non-linear effects of price changes and no evidence for asymmetric effects of price increases and price decreases.

A fourth potential violation of the exogeneity assumption may be caused by changes in ownership type, the spread of private hospital chains, or by mergers of hospitals. In Column 5 of Table 2 we exclude hospitals from our sample who are situated in the states of Hesse or Bavaria where private hospital chains were most active during our study period. In column 6 of Table 2 we exclude hospitals from our sample who were affected by a change in the type of ownership during the 2004-2009 period. For both regressions the estimation coefficient of β is very similar to the baseline specification. The coefficient of β is also essentially unchanged compared to the baseline specifications if we omit regional characteristics or the change in the average price of competitors from the regression equation (Columns 2 and 3 of Table 2). There is also no statistically significant correlation between $\Delta price_{i,2004-2009}$ and an indicator for whether or not a hospital was involved in a merger during the 2004-2009 period (see Table A2 in the online Appendix).

Furthermore, hospitals with decreases in base rate factors had significantly higher initial capacity utilization than hospitals with increases in base rate factors (see Table A3 in online Appendix). Thus, hospitals that started with higher capacity utilization before the reform also saw larger post-reform increases in the number of admissions. Subsequent increases in admission numbers cannot be explained by low initial capacity utilization.

Heterogeneous effects

In Table 3 we show how the effect of a change in base rate factors on the number of admissions varies according to hospital characteristics such as ownership status, size, and the competiveness of the local market environment. Column 1 shows results by type of ownership which could be public, not-for-profit, or private. The coefficient for not-for-profit hospitals (-0.20) is slightly more negative than for private (-0.14) and public (-0.08) hospitals. However, the differences between the coefficients are not statistically significant.

Size is measured by indicators that show whether a hospital's total number of admissions is above or below the median and alternatively whether a hospital's total number of beds is above or below the median. We find that for hospitals with a low volume (-0.19) or few beds (-0.19) the coefficient is more negative than for hospitals with a large volume (-0.05) and many beds (-0.01). However, the differences between the coefficients are not statistically significant.

The competitiveness of the environment is captured by indicators that show whether the Herfindahl index is above or below the median and alternatively whether the population density is above or below the median. We find slightly more negative coefficients for hospitals in areas with a high HHI (-0.18) or with a high population density (-0.15) compared to hospitals in areas with a low HHI (-0.09) or with a low population density (-0.12). These differences are however not statistically significant. In additional analyses we restrict the sample to hospitals that did not switch categories, e.g. their ownership status stayed the same over 2004–2009. The results are very similar.[19]

Effects of prices on the number of admissions for specific diagnoses

In Table 4 we show estimation results for the effect of prices on the number of admissions for specific diagnoses. We focus on diagnoses with large regional variation in treatment according to a report published by the Organization for Economic Cooperation and Development (OECD) (Kumar and Schönstein 2013). Diagnoses with a large amount of regional variation may be particularly susceptible to demand inducement. Specifically, we look at cataracts (three-digit ICD 9 code H25), chronic tonsillitis (three-digit ICD 9 code J35), cesarean sections (three-digit ICD 9 code O82), prostate cancer (three-digit ICD 9 code C61), and breast cancer (three-digit ICD 9 code C50). We restrict the sample to hospitals that had at least 30 admissions with the relevant diagnosis in a given year. We find that for the period 2004–2009, a one percent decrease in prices increased the number of admissions for cataracts by 1.07 percent and for tonsillitis by 0.6 percent. For cesarean sections, the coefficient for price is negative and similar in magnitude to the effect for all admissions in Table 2. However, the coefficient is not statistically significant. For prostate cancer and breast cancer the coefficients are positive and insignificant. One limitation of our data is that we only know patients' diagnosis codes and not their treatment codes. By using ICD 9 diagnosis codes we cannot distinguish whether, for example, a cancer patient underwent surgery.

[19] Results are available from the authors on request.

Effects of prices on the intensity of treatment

Table 5 shows estimation results for the effects of changes in base rate factors on the case-mix index, a measure of the average payment per patient admitted to the hospital. Our estimate for the baseline specification shown in Column 1 is -0.29. This implies, that a one percent decrease in base rate factors leads to an increase in the case-mix index by 0.29 percentage points. Correspondingly, an increase in base rate factors caused the case-mix index to decrease. This coefficient is statistically significant at the one-percent level. In the same table we also show that coefficients are robust to not including any covariates (Column 2) and to not controlling for changes in the average base rate factors of competing hospitals (Column 3). Estimation results are also not affected by regional demographic and economic time trends and by time trends in hospital characteristics (Column 4). Results are also robust to excluding the federal states with the highest share of private hospitals (Hesse and Bavaria) (Column 5) and to excluding the hospitals that change ownership status (Column 6).

These results suggest that lower prices lead to a substantial increase in the average payment for hospital patients. There are two possible explanations. First, it is possible that lower prices lead to an increase in up-coding, i.e. hospitals are more likely to classify services into higher-paying DRGs when the financial pressure is increasing (Dafny 2005, Jürges and Köberlein 2015). Second, it is also possible that hospitals respond to lower prices by adjusting the intensity of treatment (Cutler 1995).

DRG cost-weight factors for diagnoses were not constant during our study period. DRG weights are adjusted annually. Therefore, it is possible that the case-mix index for hospitals with declining prices increased not only because of up-coding or more intensive treatment but also because of higher DRG weights for the services these hospitals offered.

7. Conclusions

We examine the effect of changes in hospital reimbursement prices on volume of care based on a reform of hospital financing in Germany. We find that a one percent across-the-board increase in payment rates for the period 2004–2009 leads to a decrease in the number of hospital admissions by 0.14 percent and to a decrease in the case-mix index – a measure of the average payment per patient – of 0.29 percent. Our empirical results cannot be explained by pre-existing trends in volume growth, by heterogeneous effects of the introduction of DRG type payment, by the effects of mergers or changes in ownership type, or by differences in initial capacity utilization. We also find no evidence for asymmetric effects of price increases and price decreases.

A unique aspect of our study is that we can mostly separate income from substitution effects. Previous studies find that hospitals respond to decreases in prices for one group of patients or type of service by increasing the volume of care for other groups of patients or types of services. For example, He and Mellor (2012) find that hospitals respond to decreases in prices for Medicare outpatient services by providing more treatment to privately insured patients. In contrast, price changes in our study apply to essentially all patients and types of services hospitals provide. Specifically, price changes apply to both publicly and privately insured patients. Moreover, hospitals as well as physicians who work at hospitals are largely banned from providing outpatient services.[20] They cannot easily respond to reduced prices for inpatient care by providing more

[20] There is a strict institutional separation between inpatient and outpatient care. With some exceptions, hospitals are not allowed to provide outpatient care (Simon 2010).

outpatient services. Our results can also not be explained by demand responses to price changes. Thus, we interpret our findings as evidence for a negative income effect.[21]

One possible explanation for our findings is that a high share of costs for hospitals in Germany is fixed – at least in the short and medium run. The overwhelming part of hospital costs are personnel expenditures.[22] Hospital physicians are mostly salaried employees, and wages for both physicians and other staff are set in wage negotiations at the federal level. Layoffs are costly. Since a high share of costs is fixed the marginal cost of admitting an additional patient can be low. Low marginal costs make it attractive for hospitals to increase volume of care in response to lower prices.

Lack of possibilities for substitution and high fixed costs for inpatient care can also be possible explanations why our results differ from the results of some previous studies. For example, He and Mellor (2012) and Clemens and Gottlieb (2014) find that medical care providers respond to lower Medicare reimbursement prices for outpatient care by providing fewer services. They either increase services to patients not covered by Medicare, or they work less. Since outpatient physicians in the United States tend to be self-employed, their salaries are not part of the providers' fixed costs. Interestingly, He and Mellor (2012) find that the negative volume response to price decreases is smaller for hospitals with a higher share of Medicare patients and for hospitals where physicians are vertically integrated into the hospital. Thus, their results become more similar to our results if the institutional setting is more similar to the institutional setting in our study.

[21] Some substitution to other types of care is still possible in the German institutional setting, for example by treating more patients with residence outside the European Union. Thus, our estimates present upper bounds, and the true income effect might be even more negative than the income effect we estimate.

[22] In 2014, personnel expenditures accounted for 59.9% of hospital expenditures (German Statistical Office 2015)

Our findings have important policy conclusions. They suggest that existing rules to limit volume growth for German hospitals during our study period did not prevent hospitals from expanding volumes in response to lower prices. Recent law changes have strengthened rules which aim to limit growth in hospital volumes (KHSG 2015). For example, hospitals now have to inform patients about their right to obtain a second medical opinion for selected surgery procedures which are deemed to be sensitive to demand inducement. Also hospitals face reductions in payment if they exceed target volumes. Starting from the year 2017, these deductions are calculated based on the share of fixed costs that can be attributed to a given procedure in a given hospital (Fixkostendegressionsabschlag). In light of our findings it will be interesting to see how hospitals respond to these new regulations.

Better understanding the objective function of hospitals is of high interest for both policymakers and academics. In this study, we find a negative income effect for hospitals in Germany. It will be interesting to learn from future research whether negative income effects for medical providers can also be found in other countries and for other institutional settings.

References

Augurzky, Boris, Sebastian Krolop and Christoph M. Schmidt (2010). "Die wirtschaftliche Lage der Krankenhäuser." Chapter 2 in: Krankenhaus-Report 2010, edited by Jürgen Klauber, Max Geraedts and Jörg Friedrich, Schattauer.

Busse, Reinhard and Annette Riesberg (2004). "Health care systems in transition: Germany." WHO Regional Office for Europe on behalf of the European Observatory on Health Systems and Policies, Copenhagen.

Clemens, Jeffrey and Joshua D. Gottlieb (2014). "Do physicians' financial Incentives affect medical treatment and patient health?" *American Economic Review*, 104, 1320-1349.

Congressional Budget Office (2007). "Factors underlying the growth in Medicare's spending for physicians' services." The Congress of the United States, *Background Paper* 2597.

Cutler, David M. (1995). "The incidence of adverse medical outcomes under prospective payment." *Econometrica*, 63, 29-50.

Dafny, Leemore S. (2005). "How do hospitals respond to price changes?" *American Economic Review*, 95, 1525-1547.

Evans, Robert G. (1974). "Supplier-induced demand: some empirical evidence and implications." Chapter 10 in: The economics of health and medical care, edited by Mark Perlman, Macmillan.

Friedrich, Jörg, Gregor Leclerque and Kathrin Paschen (2008). "Die Krankenhausbudgets 2004 bis 2006 unter dem Einfluss der Konvergenz." Chapter 15 in: Krankenhaus-Report 2007, edited by Jürgen Klauber, Bernd Peter Robra and Henner Schellschmidt, Schattauer.

GKV-Spitzenverband (2012). "Studie belegt: Zuweisungen gegen Entgelt keine Einzelfälle – erhebliches Korruptionspotential." Press Release from May 22, 2012 by GKV-Spitzenverband (Association of public health insurers), Berlin.

German Statistical Office (2015). "Gesundheit: Kostennachweis der Krankenhäuser." Fachserie 12, Reihe 6.3, Wiesbaden.

Gruber, Jonathan and Maria Owings (1996). "Physician financial incentives and Caesarean section delivery." *RAND Journal of Economics*, 27, 99-123.

He, Daifeng and Jennifer Mellor (2012). "Hospital volume responses to Medicare's outpatient prospective payment system: evidence from Florida." *Journal of Health Economics*, 31, 730-743.

Heaton, Paul and Eric Helland (2009). "Does treatment respond to reimbursement rates? Evidence from Trauma Care." *RAND Working Paper* WR 648-ICJ.

Januleviciute, Jurgita, Jan Erik Askildsen, Oddvar Kaarboe, Luigi Siciliani and Mark Sutton (2016). “How do hospitals respond to price changes? Evidence from Norway.” *Health Economics*, 25, 620-636.

Jürges, Hendrik and Juliane Köberlein (2015). "What explains DRG upcoding in neonatology? The roles of financial incentives and infant health." *Journal of Health Economics*, 43, 13-26.

KHSG (2015). “Gesetz zur Reform der Strukturen der Krankenhausversorgung (Krankenhausstrukturgesetz - KHSG).” *Bundesgesetzblatt Teil I* 51, Bundesanzeiger Verlag.

Kumar, Ankit and Michael Schönstein (2013). “Managing hospital volumes, Germany and experiences from OECD countries.” *OECD Health Working Paper* 67.

McGuire, Thomas G. and Mark V. Pauly (1991). “Physician responses to fee changes with multiple payers” *Journal of Health Economics*, 10, 386-410.

McGuire, Thomas G. (2000). “Physician Agency.” Chapter 9 in: *Handbook of Health Economics* Volume 1, edited by Anthony J. Culyer and Joseph P. Newhouse, Elsevier.

Quentin, Wilm, Alexander Geissler, David Schneller-Kreinsen and Reinhard Busse (2010). “DRG-type hospital payment in Germany: The G-DRG system.” *Euro Observer*, 12, 4-7.

Pilny, Adam (2014). “Mergers and Acquisitions in the German Hospital Market – Who are the targets?” *Ruhr Economic Papers* 518.

Rice, Thomas H. (1983). “The impact of changing reimbursement rates on physician-induced demand.” *Medical Care*, 21, 803-815.

Rizzo, John A. and Richard J. Zeckhauser (2003). “Reference income, loss aversion, and physician behavior.” *Review of Economics and Statistics*, 85, 909-922.

Salfeld, Rainer, Steffen Hehner and Reinhard Wichels (2009). “Modernes Krankenhausmanagement: Konzepte und Lösungen.“, Springer.

Shigeoka, Hitoshi and Kiyohide Fushimi (2014). “Supply induced demand in newborn treatment: evidence from Japan.” *Journal of Health Economics*, 35, 162-178.

Simon, Michael (2010). “Das Gesundheitssystem in Deutschland: Eine Einführung in Struktur und Funktionsweise.“ 3rd edition, Huber.

Thurn, Thomas (2013). “Ausgewogene Anreize nötig.” 14. Auflage der Kienbaum-Vergütungsstudie zur Gehaltsentwicklung von Führungskräften in Krankenhäusern. *Arzt und Krankenhaus*, 4, 104-107.

Tuschen, Karl Heinz, Thomas Braun and Ferdinand Rau (2005). “Erlösausgleiche im Krankenhausberiech: Eine Orientierungshilfe.“ *Das Krankenhaus*, 11, 955-960.

Yip, Winnie C. (1998). “Physician response to Medicare fee reductions: change in the volume of coronary artery bypass graft (CABG) surgeries in the Medicare and private sectors.” *Journal of Health Economics*, 17, 675-699.

Table 1: Descriptive statistics

	Year 2004		Year 2009	
	Mean	Standard dev.	Mean	Standard dev.
Number of admissions	10940.590	10452.560	11878.680	11289.700
Case-mix index (CMI)	1.001	0.264	1.012	0.407
Public hospitals	0.408	0.492	0.391	0.488
Not-for-profit hospitals	0.451	0.498	0.442	0.497
Herfindahl index (HHI)	0.189	0.131	0.198	0.139
Unemployment rate	9.935	4.144	7.726	3.007
Average age men	36.950	1.006	37.889	0.889
Average age women	40.093	1.433	40.628	1.298
Population density	0.678	0.722	0.681	0.735
Number of hospitals	801		801	

Notes: Descriptive statistics refer to sample in baseline specification (Table 2, column 1).

Table 2: Effects of price changes on number of admissions

	Log volume					
	Baseline	Without covariates	Without price competitors	With additional trends	Without Hesse and Bavaria	No change of owner type
	(1)	(2)	(3)	(4)	(5)	(6)
Δ price (2004-2009)	-0.136** (0.055)	-0.133** (0.055)	-0.132** (0.055)	-0.130** (0.056)	-0.155*** (0.060)	-0.132** (0.058)
Year effects	Yes	Yes	Yes	Yes	Yes	Yes
Regional characteristics	Yes	No	Yes	Yes	Yes	Yes
Average price of competitors	Yes	No	No	Yes	Yes	Yes
Trends by regional and hospital characteristics	No	No	No	Yes	No	No
N (Hospitals)	801	801	801	801	689	755

Notes: The table shows estimation coefficients of a fixed-effects linear regression model. The sample includes observations for the years 2004 and 2009. Δ price (2004-2009) is defined as log(price 2009) – log(price 2004). Regional indicators include average age of men, average age of women, population density, and unemployment rate. All regional characteristics are measured for a hospital's catchment area. Trends by regional and hospital characteristics are captured by interaction terms of an indicator for the year 2009 and the following regional and hospital characteristics (measured in the year 2004): average age of men, average age of women, population density, unemployment rate, large number of beds, a high HHI index, public ownership, and not-for profit ownership. Parentheses show robust standard errors, clustered at the hospital level. * significant at 10%; ** significant at 5%; *** significant at 1%

Table 3: Heterogeneous effects of price changes on number of admissions

			Log volume		
	ownership	Number admissions	Number beds	HHI index	Population density
	(1)	(2)	(3)	(4)	(5)
Δ price (2004-2009)	-0.076				
* public	(0.071)				
Δ price (2004-2009)	-0.199***				
* not-for-profit	(0.066)				
Δ price (2004-2009)	-0.136				
* private	(0.119)				
Δ price (2004-2009)		-0.045			
* large volume		(0.073)			
Δ price (2004-2009)		-0.186***			
* low volume		(0.079)			
Δ price (2004-2009)			-0.010		
* many beds			(0.054)		
Δ price (2004-2009)			-0.191***		
* few beds			(0.079)		
Δ price (2004-2009)				-0.182**	
* high HHI				(0.065)	
Δ price (2004-2009)				-0.089	
* low HHI				(0.089)	
Δ price (2004-2009)					-0.154**
* high population density					(0.078)
Δ price (2004-2009)					-0.116
* low population density					(0.077)
Year effects	Yes	Yes	Yes	Yes	Yes
Regional characteristics	Yes	Yes	Yes	Yes	Yes
Average price of competitors	Yes	Yes	Yes	Yes	Yes
N (Hospitals)	801	801	801	801	801

Notes: The table shows estimation coefficients of a fixed-effects linear regression model. The sample includes observations for the years 2004 and 2009. Δ price (2004-2009) is defined as log(price 2009) – log(price 2004). Regional indicators include average age of men, average age of women, population density, and unemployment rate. All regional characteristics are measured for a hospital's catchment area. Parentheses show robust standard errors, clustered at the hospital level. * significant at 10%; ** significant at 5%; *** significant at 1%

Table 4: Effect of price changes on volume of treatment for specific diagnoses

	Log volume				
	Cataracts	Tonsillitis	C-section	Prostate cancer	Breast cancer
	(1)	(2)	(3)	(4)	(5)
Δ price (2004-2009)	-1.071** (0.463)	-0.592*** (0.181)	-0.267 (0.647)	0.045 (0.256)	0.103 (0.263)
Year effects	Yes	Yes	Yes	Yes	Yes
Regional characteristics	Yes	Yes	Yes	Yes	Yes
Average price of competitors	Yes	Yes	Yes	Yes	Yes
N (hospitals)	114	335	87	219	387

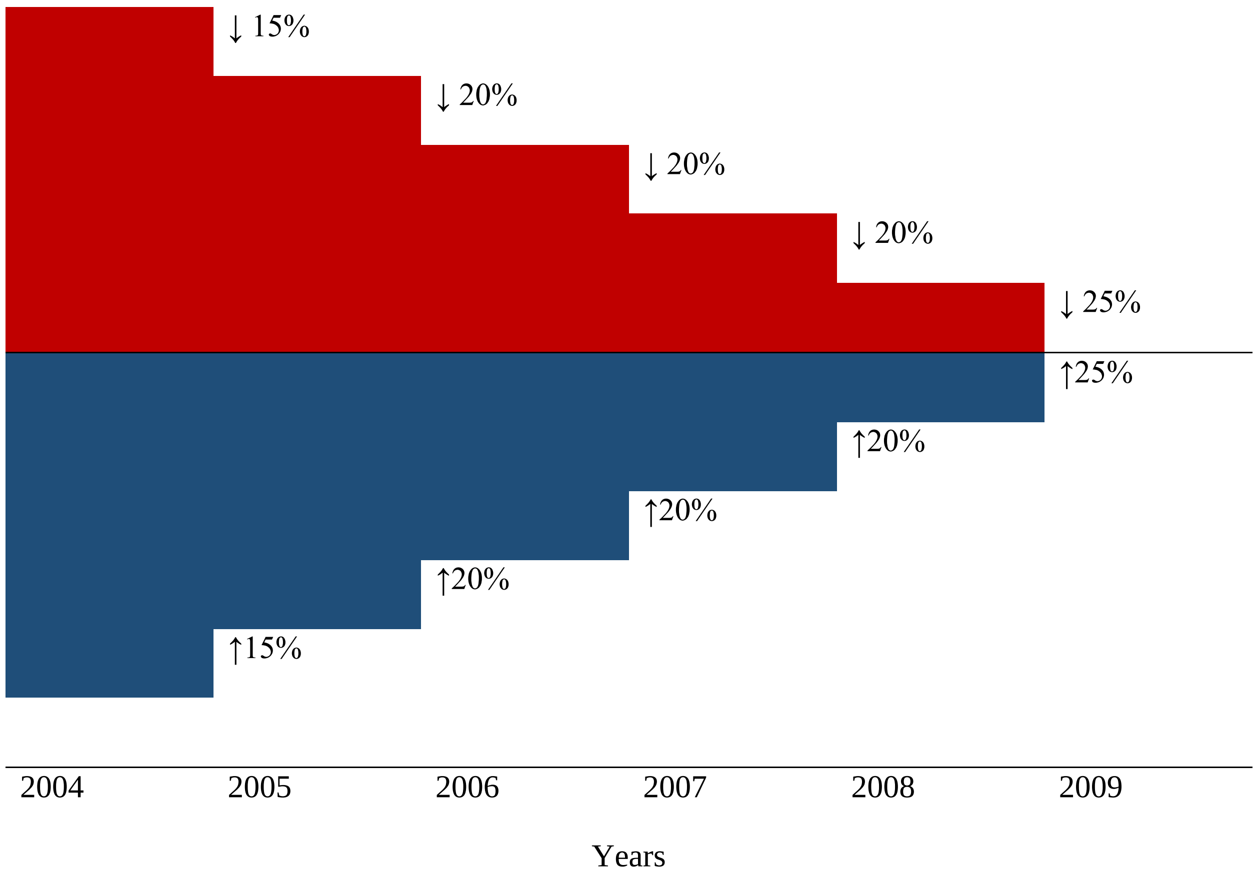
Notes: The table shows estimation coefficients of a fixed-effects linear regression model. The sample includes observations for the years 2004 and 2009. Δ price (2004-2009) is defined as log(price 2009) – log(price 2004). Regional indicators include average age of men, average age of women, population density, and unemployment rate. All regional characteristics are measured for a hospital's catchment area. Parentheses show robust standard errors, clustered at the hospital level. * significant at 10%; ** significant at 5%; *** significant at 1%

Table 5: Effects of price changes on case-mix index

	Case-mix index					
	Baseline	Without covariates	Without price competitors	With additional trends	Without Hesse and Bavaria	No change of owner type
	(1)	(2)	(3)	(4)	(5)	(6)
Δ price (2004-2009)	-0.285*** (0.082)	-0.292*** (0.081)	-0.287*** (0.085)	-0.273*** (0.105)	-0.295*** (0.088)	0.279*** (0.084)
Year effects	Yes	Yes	Yes	Yes	Yes	Yes
Regional characteristics	Yes	No	Yes	Yes	Yes	Yes
Average price of competitors	Yes	No	No	Yes	Yes	Yes
Trends by regional and hospital characteristics	No	No	No	Yes	No	No
N (Hospitals)	788	788	788	788	678	743

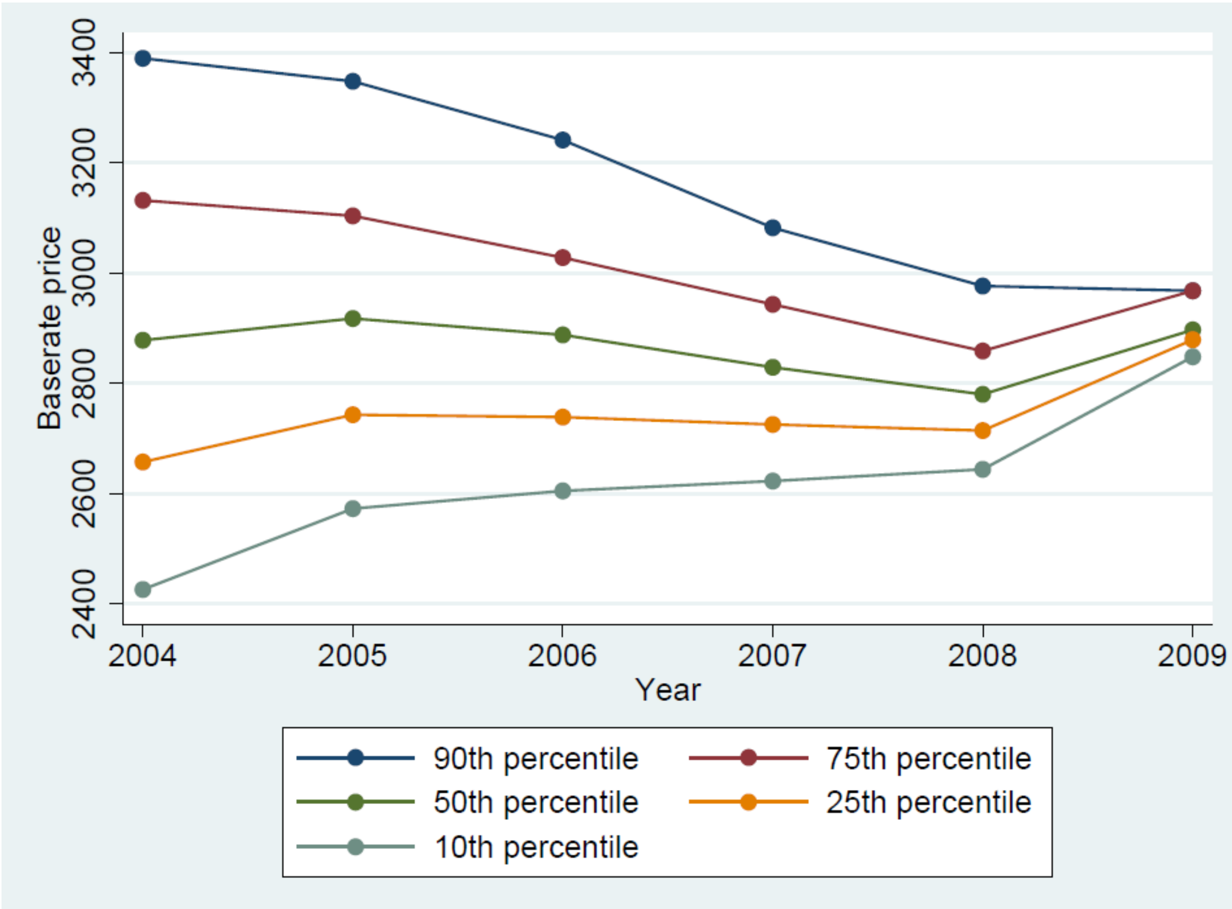
Notes: The table shows estimation coefficients of a fixed-effects linear regression model. The sample includes observations for the years 2004 and 2009. Δ price (2004-2009) is defined as log(price 2009) – log(price 2004). Regional indicators include average age of men, average age of women, population density, and unemployment rate. All regional characteristics are measured for a hospital's catchment area. Trends by regional and hospital characteristics are captured by interaction terms of an indicator for the year 2009 and the following regional and hospital characteristics (measured in the year 2004): average age of men, average age of women, population density, unemployment rate, large number of beds, a high HHI index, public ownership, and not-for profit ownership. Parentheses show robust standard errors, clustered at the hospital level. * significant at 10%; ** significant at 5%; *** significant at 1%

Figure 1: Convergence of base rate factors



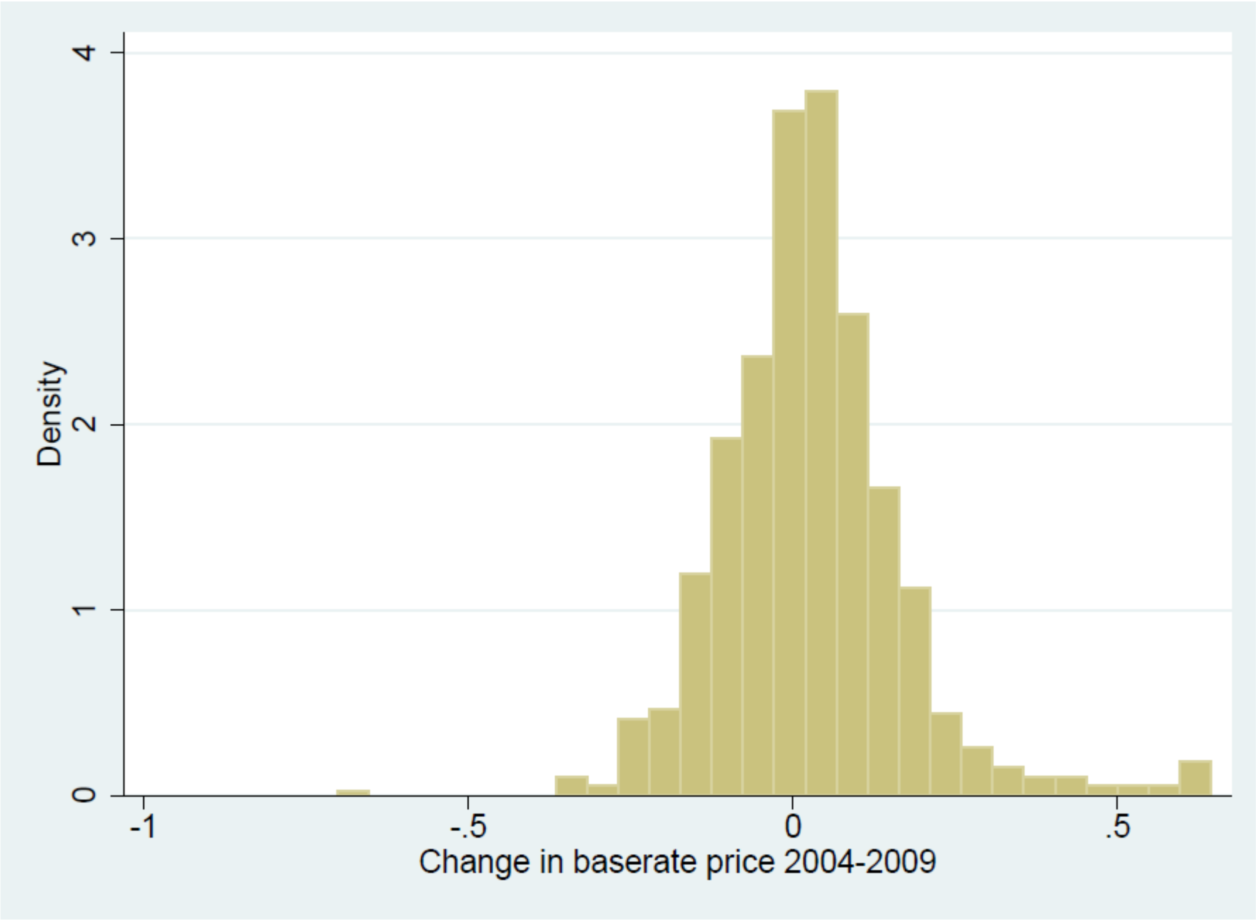
Notes: This figure shows the reduction of initial differences in base rate factors during the convergence period.

Figure 2: Convergence of base rate factors (empirical evidence)



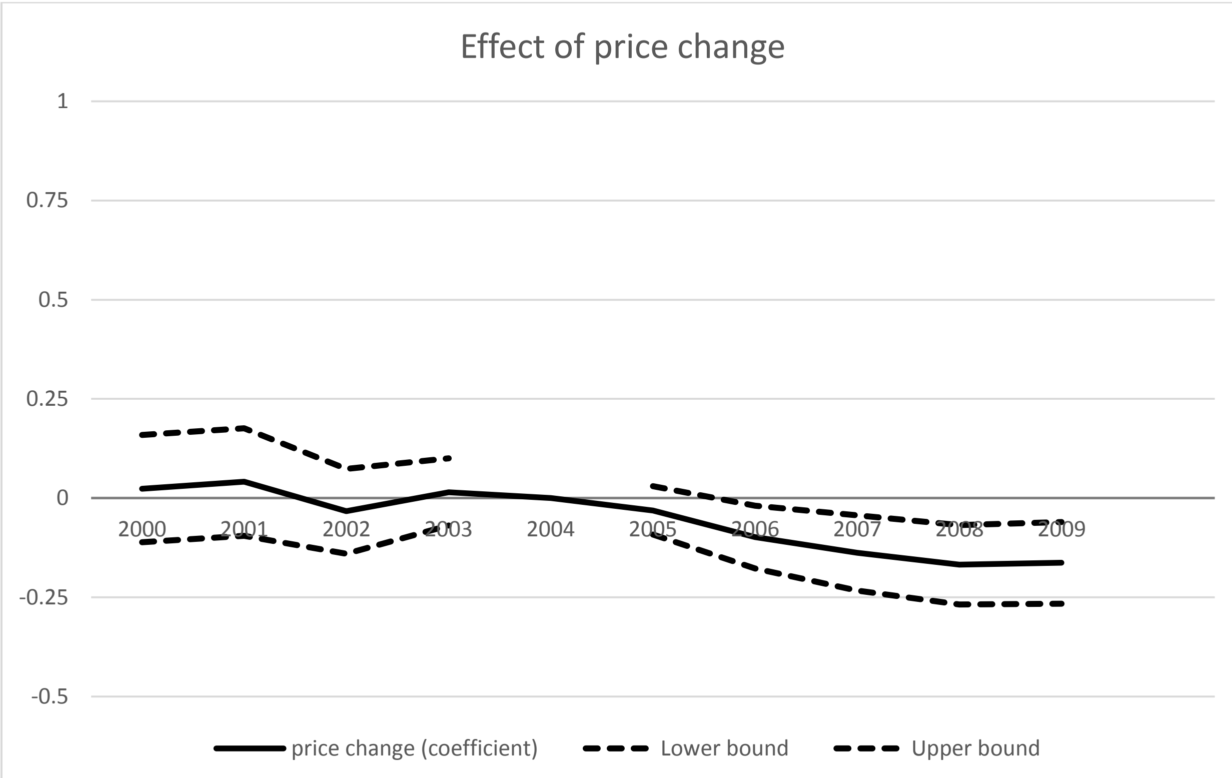
Notes: This figure shows the empirical distribution of hospital-specific base rate factors for quantiles of initial prices and the reduction of initial differences in base rate factors during the convergence period.

Figure 3: Distribution of price changes between 2004-2009



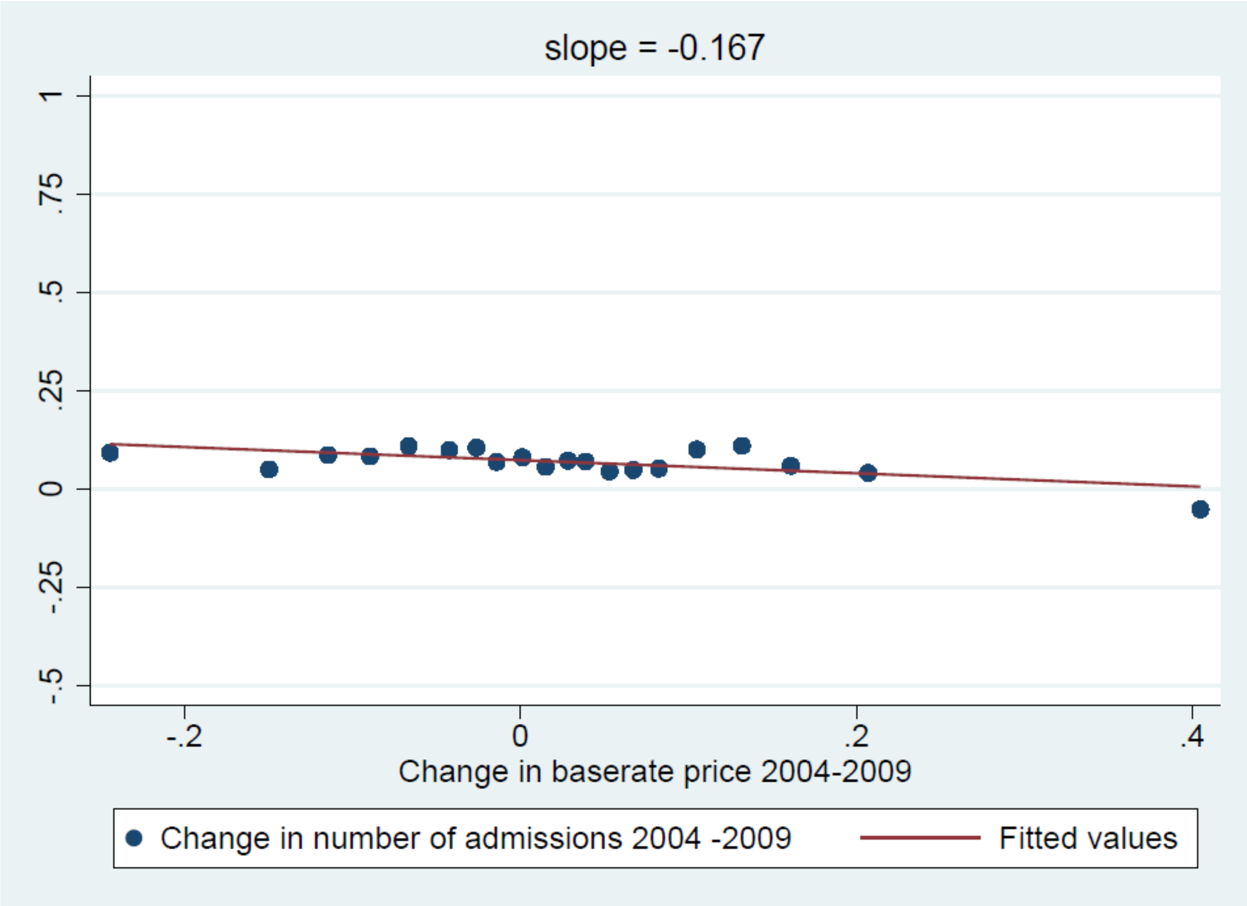
Notes: This figure shows the distribution of Δ price (2004-2009). Δ price (2004-2009) is defined as log(price 2009) – log(price 2004). The sample consists of N = 801 hospitals.

Figure 4: Pre-trends and post-trends for the effect of price changes on the number of admissions



Notes: The figure shows estimation coefficients and their 95 percent confidence intervals for the effect of changes in base rate factors (Δ price (2004-2009)) between the years 2004 and 2009 on the number of admissions in each year between 2000 and 2009. Δ price (2004-2009) is defined as log(price 2009) – log(price 2004).The coefficients are based on estimation equation (7). The sample includes all hospitals which are included in the baseline specification in Table 2 column 1. Standard errors are robust and clustered at the hospital level.

Figure 5: Non-linear effects of base rate factors on number of admissions



Notes: We group the sample of hospitals into 20 bins of equal size ("ventiles") according to their total change in base rate factors over 2004-2009. The x-axis displays the mean of price changes for hospitals in each ventile. The y-axis shows, for each ventile, the average difference in log number of hospital admissions between 2004 and 2009. We estimate a trend line from OLS regression using the 20 data points shown in the graph.

Online Appendix

Table A1: Determinants of price changes

	Price change 2004-2009	
	Coefficient	Standard error
Large number of beds in 2004	-0.048***	(0.012)
High HHI index in 2004	0.048***	(0.011)
Public ownership in 2004	-0.032	(0.022)
Not-for-profit ownership in 2004	-0.010	(0.022)
Unemployment rate in 2004	-0.001	(0.002)
Average age men in 2004	0.021	(0.018)
Average age women in 2004	0.024*	(0.013)
Population density in 2004	-0.017**	(0.007)
N (hospitals)	801	

Notes: The table shows OLS estimation coefficients. Parentheses show robust standard errors. * significant at 10%; ** significant at 5%; *** significant at 1%

Table A2: Robustness check – Relationship between initial base rate factor and subsequent mergers

	Mergers (2004–2009) (1)	Mergers (2004–2009) (2)
Δ price (2004-2009)	-0.038	-0.061
	(0.079)	(0.081)
Regional characteristics 2004	No	Yes
Average price competitors 2004	No	Yes
N (hospitals)	1226	1226

Notes: The table is based on data from the RWI hospital panel (a different data source compared to the other tables), and it shows OLS estimation coefficients. 259 hospital in the sample (= 21.1%) are involved in a merger between the years 2004 and 2009. Δ price (2004-2009) is defined as log(price 2009) – log(price 2004). Regional characteristics include average age of men, average age of women, population density, and unemployment rate in a hospital's catchment area. Parentheses show robust standard errors. * significant at 10%; ** significant at 5%; *** significant at 1%

Table A3: Alternative explanation – different capacity utilization before the reform

	Capacity utilization 2003 (1)	Capacity utilization 2003 (2)
Δ price (2004-2009)	-0.110***	-0.103***
	(0.029)	(0.029)
Regional characteristics 2004	No	Yes
N (hospitals)	800	800

Notes: The table shows OLS estimation coefficients. Regional indicators include log average price of competitors, average age of men, average age of women, population density, and unemployment rate in a hospital's catchment area. Parentheses show robust standard errors. * significant at 10%; ** significant at 5%; *** significant at 1%