THE CENTRE FOR MARKET AND PUBLIC ORGANISATION

Death by Market Power
Reform, Competition and Patient Outcomes in the National Health Service

Martin Gaynor, Rodrigo Moreno-Serra
and Carol Propper

July 2010 (updated August 2011)

Working Paper No. 10/242

Centre for Market and Public Organisation
Bristol Institute of Public Affairs
University of Bristol
2 Priory Road
Bristol BS8 1TX
http://www.bristol.ac.uk/cmpo/

Tel: (0117) 33 10799
Fax: (0117) 33 10705
E-mail: cmpo-admin@bristol.ac.uk

ISSN 1473-625X

CMPO Working Paper Series No. 10/242

Death by Market Power
Reform, Competition and Patient Outcomes in the National Health Service

Martin Gaynor[1]
Rodrigo Moreno-Serra[2]
and
Carol Propper[3]

[1] *Carnegie Mellon University and University of Bristol and NBER*
[2] *Imperial College*
[3] *Imperial College, University of Bristol and CEPR*

July 2010 (updated August 2011)

Abstract
The effect of competition on the quality of health care remains a contested issue. Most empirical estimates rely on inference from non-experimental data. In contrast, this paper exploits a pro-competitive policy reform to provide estimates of the impact of competition on hospital outcomes. The English government introduced a policy in 2006 to promote competition between hospitals. Using this policy to implement a difference-in-differences research design, we estimate the impact of the introduction of competition on not only clinical outcomes but also productivity and expenditure. We find that the effect of competition is to save lives without raising costs.

Keywords: competition, hospitals, quality
JEL Classification: I18, I11, L13, L32
Electronic version **www.bristol.ac.uk/cmpo/publications/papers/2010/wp242.pdf**

Acknowledgements
Many thanks are due to Carolyn Whitnall for data preparation and to George Leckie for initial data description. HES data were used with the permission of the Department of Health (The NHS Information Centre). Financial support was provided by the U.K. Department of Health under the HREP policy evaluation programme. We appreciate comments from Dave Balan, Simon Burgess, Dennis Epple, Vic Fuchs, Kate Ho, Jon Kolstad, Fiona Scott Morton, Bob Town, Ellerie Weber, Frank Windemeijer, participants in seminars at the NBER Summer Institute, the Free University of Amsterdam, Imperial College London, Carnegie Mellon, the Universities of Bergen, Bristol, Illinois at Urbana-Champaign, Illinois at Chicago, Indiana at Indianapolis, Minnesota, Pennsylvania, and Southern California, the Co-operation & Competition Panel, the Federal Reserve Bank of Minneapolis, American Enterprise Institute, and the American Society of Health Economists 2010 meeting. The authors are solely responsible for the analysis and the views expressed here.

Address for correspondence
CMPO, Bristol Institute of Public Affairs
University of Bristol
2 Priory Road
Bristol BS8 1TX
Carol.Propper@bristol.ac.uk
www.bristol.ac.uk/cmpo/

I. Introduction

Health care is one of the most important industries in developed countries, both because of its size and impact on well-being. Historically, health care has been provided through centralized, non-market means in most countries outside of the United States. However, recently market oriented reforms have been adopted or are being considered in many countries, including the UK, Netherlands, Belgium, Israel, and Australia, despite a lack of strong evidence on the effects of market reforms in health care. In the US markets have long been used for the delivery of health care. However, massive consolidation among hospitals has led to concerns about the functioning of these markets.[1] These developments raise questions as to whether pro-market reforms are an appropriate way of improving outcomes in health care and in particular, because of the importance of quality in health care, whether competition will deliver the socially optimal quality of care (Sage et al. 2003; Federal Trade Commission and US Department of Justice 2004).[2]

Analysing the relationship between competition and quality has been hampered by the endogeneity of market structure. Most existing studies come from the US Medicare program and use changes in cross sectional variation in levels of market structure over time to identify the impact of competition. But market structure may be determined by quality – for example, the existence of a high quality provider may deter entry by potential rivals.

In this paper we exploit an exogenous policy change to examine the effect of competition on health care quality. Our identification comes from a major policy shift designed to promote competition in the English National Health Service (NHS). In 2006 the NHS mandated that all patients requiring treatment be given the choice of five different hospitals and adopted a payment system in which hospitals were paid fixed, regulated prices for treating patients (similar to the Medicare hospital payment system in the US). Prior to this reform, the local public agencies responsible for purchasing health care on behalf of the population in their area engaged in selective contracting with hospitals, bargaining over price and quantity, referring primarily to local hospitals and providing hospitals with income that was based primarily on past activity.

[1] For example, Haas-Wilson (2003), Sage et al. (2003), Federal Trade Commission and US Department of Justice (2004), Cuellar and Gertler (2005) and Vogt and Town (2006). Critics of the use of competition include Schlesinger (2006), Rosenbaum (2006) and Jost et al. (2006).

[2] Quality is a major issue in health care, first because the effect of quality on an individual's well-being can be very great and second, because consumers in most systems are not exposed to the full expense associated with their health care decisions so that quality looms larger in consumer choice than price.

The reform gave patients more choice (via the mandated five alternatives and the end of selective contracting), increased the incentive for hospitals to win business and moved hospitals from a market determined price environment to a regulated price environment.

However, while the pro-competitive policy was national, the intensity of the competition induced by the reforms will vary according to pre-reform market structure, which is a function of the geographical configuration of patient location and hospital sites. In some places population density results in a market structure which permits a high degree of choice. In others, population density is low and hospitals are located in highly concentrated markets with few competitors. As a consequence, post-policy a hospital in an unconcentrated market faces more exposure to the policy change than does a hospital in a highly concentrated market. We exploit this variation in potential competition across time and space to identify the impact of competition.

The essence of our empirical approach is illustrated in Figures 1a and 1b. They show the unconditional relationship between hospital quality (as measured by mortality) and market structure (measured by the Herfindahl-Hirschman concentration index, HHI) before and after the reform.[3] The red lines show the smoothed nonparametric relationship between mortality and market structure. Figure 1a presents the mortality rate for acute myocardial infarction (AMI, commonly known as a heart attack), a commonly used measure of hospital quality.[4] The left hand panel shows that mortality is clearly higher in less concentrated markets before the reform. But post-reform the differences have disappeared due to a larger reduction in mortality in unconcentrated markets. The same pattern is observed in Figure 1b, which plots all-cause hospital mortality rates against the HHI. These figures suggest that the reform intensified competition and that intensified competition post-reform led to increased quality. The paper subjects this hypothesis to rigorous testing.

We first develop a simple theoretical model that allows us to understand how competition might operate in the setting of the reform we study. The limited theoretical literature on competition and quality in health care markets (see Gaynor 2006, for a review) suggests that competition will increase quality in markets with regulated prices provided that this price is above marginal cost. The intuition is as follows. Price is regulated, so firms compete for consumers on non-price dimensions, i.e. ‘quality.’ If the regulated price is set above marginal

[3] Adjusted for differences in severity between hospital populations.
[4] We discuss the use of mortality as an indicator of quality in section IV.A.

cost at some baseline level of quality, then firms will increase quality to try to gain market share. This will continue until profits are zero. But these models do not explicitly deal with the fact that hospitals are multi-product firms and supply services where little competition is possible (emergency care) and markets where competition is much more possible (elective care).[5] Our model explicitly addresses this issue. It shows that greater choice should increase the (quality) elasticity of demand for elective care faced by hospitals, which should intensify competition in that sector. In addition, the shift from selective contracting to fixed prices should focus competition on quality (conditional on fixed prices greater than marginal costs). Further, competition in the market for elective care will spillover and improve quality for emergency patients.

Second, we examine, as most of the rest of the literature (reviewed below), quality as measured by deaths following emergency admissions of AMI patients, but also a range of other quality measures and other outcomes, including hospital utilization and expenditure. Third, we are careful to show that the policy change did lead to changes in care seeking patterns that support a picture of greater responsiveness to quality post-reform. Finally, we undertake a large number of robustness tests to ensure that our identification strategy is not compromised by other factors such as changes in local demand, costs, or other policy interventions. Our results strongly indicate that the introduction of pro-competitive reforms led to an increase in quality without a commensurate increase in expenditure.

Our research contributes to the empirical literature on competition and quality in health care. Almost all of these use US data. The most prominent study of markets with fixed prices is Kessler and McClellan (2000), who examine the impact of market concentration on mortality for Medicare AMI patients. They find that mortality is substantially and significantly higher for patients in more concentrated markets. Kessler and Geppert (2005) find that high-risk Medicare patients' heart attack mortality is higher in highly concentrated markets, while there is no such effect for low-risk patients. Tay (2003) estimates a model of hospital choice for Medicare patients and finds that demand is responsive to quality, again measured by heart attack mortality, implying the potential for quality competition. Shen (2003) finds that the number of hospitals

[5] These models largely derive from analyses of industries subject to price regulation up until the 1970s and 1980s, such as airlines and taxis., though there are also some models specific to health care. See, for example, Douglas and Miller (1974), Schmalensee (1977), Vander Weide and Zalkind (1981) and White (1972) on airlines, and Frankena and Pautler (1984) on taxicabs. On health care see, for example, Allen and Gertler (1991), Held and Pauly (1983) and Pope (1989).

interacted with the Medicare payment leads to reduced Medicare patient heart attack mortality after 1990. In contrast, Gowrisankaran and Town (2003), using similar methods to Kessler and McClellan (2000), find that mortality is higher for Medicare heart attack and pneumonia patients receiving care in less concentrated markets in the Los Angeles area and Mukamel et al. (2001) find no effect of market concentration on mortality from all causes for Medicare patients. Cutler et al. (2010) examine the impact of entry into the market for heart bypass surgery. They find that entry led to improved quality, but that the welfare gains from increased quality are offset by fixed costs of entry.

There are a handful of papers which examine the impact of health care competition in the UK. Propper et al (2008) use a similar identification strategy to the present paper but study an earlier policy regime in which competition in the UK health care market was introduced in 1991 and abolished in 1997. In this regime, prices were negotiated and measures of quality very limited and not publicly available. The study found that competition reduced unmeasured quality, as measured by deaths following emergency heart attack admissions, but also decreased waiting times, which were observed. Cooper et al (2010) adopt the same identification strategy to examine the effect of the current pro-market reforms in the UK. They examine only one measure of quality, deaths following emergency heart attacks, and find the reforms reduced AMI mortality in less concentrated markets relative to more concentrated markets. Bloom et al (2010) examine the impact of competition on management quality and management quality on a range of outputs, including heart attack mortality and financial performance. Whilst they have only cross sectional data on management quality, they exploit variation in hospital closures that is driven by the political process to identify the impact of competition on management quality. They find that management quality is higher in more competitive markets and that higher management quality is associated with better outputs, including lower AMI mortality.

Our research also contributes to the growing evidence on the impact of competition in public services. There has been a great deal of interest in recent years in competition in education, both theoretically and empirically (e.g., Epple and Romano 1998; Hoxby 2000; Epple, Figlio, and Romano 2004). In this literature, as in health, the predictions from theoretical models are often ambiguous and the empirical evidence quite contested (e.g., Hoxby 2000; Rothstein 2007; Bayer and McMillan 2005; for a review see Burgess, Propper and Wilson 2005). Our results thus add to the evidence on the conditions under which gains from competition in the provision of public services may be realized.

In what follows, we provide background on the NHS and the market oriented reforms of 2006 (Section II), describe a simple theoretical model of hospital quality competition (Section (III), present our empirical strategy (Section IV), describe the data (Section V), present our results and report on an extensive set of robustness tests (Section VI), and provide a summary and conclusions (Section VII).

II. The reform program

A. The NHS reforms

In the UK health care is tax financed and free at the point of use. Almost all care is provided by the National Health Service (NHS). Primary care is provided in the community by publicly funded physicians known as General Practitioners (GPs), who also act as the 'gate keeper' for hospital-based care. Secondary care is provided in publicly funded public (NHS) hospitals.[6]

Prior to 1991 funding was allocated to public bodies at the local level (local health authorities), who were responsible for running hospitals. From 1991 the roles of buyer and seller of hospital-based health care were separated, with the intention of promoting competition between public hospitals. The local health authorities were given the task of buying hospital-based health care for their population.[7] Hospitals were turned into free standing public organizations, known as NHS Trusts, who competed for contracts from the buyers.[8] Both price and quality were negotiable, though information on quality was extremely limited (Propper et al. 2008). In 1997 the newly elected Labour administration retained the architecture of the buyer and seller split but changed policy to reduce competition and to implement instead longer term cooperative relationships between buyers and sellers. In this regime, which was similar to selective contracting in the US, buyers and sellers negotiated over price, and to some degree over

[6] There is a small private sector. Around 15 percent of the population have private health insurance which is a complement to, rather than a substitute for NHS care. Private medical insurance is not tax deductible and covers care in the small private sector, which specializes in the hospital based provision of non-acute services for which there are long NHS waiting lists.

[7] These buyers were known as District Health Authorities. Health authorities covered an administratively defined geographical area containing around 100,000 patients.

[8] Although purchasers were given the right to buy from whichever supplier of health care they wished, in practice, almost all care purchased by NHS purchasers was bought from NHS Trusts. Relatively little business went to the very small private sector.

quality (mainly waiting times, not clinical outcomes), and volume on an annual basis, with the majority of contracts taking the form of annual bulk-purchasing contracts.

In late 2002 the government signalled a shift in policy and initiated a reform package with a set of phased-in changes leading to the re-introduction of competition from 2006 onwards.[9] There were several elements to this policy (Farrar et al. 2007; Cooper et al. 2010), the most important of which were a policy designed to increase patient choice and a change in hospital payments from negotiated to ex ante fixed prices.

Patient choice was introduced in January 2006. Prior to this date, patients were referred by their GPs to the local hospital that provided the service they required and were not generally offered any choice over the location of their care.[10] After January 2006 patients had to be offered a choice of five providers for their hospital care (Department of Health 2004) and GPs were required (and paid) to ensure that patients were made aware of, and offered, choice.[11] Along with giving patients a formal choice of where they could receive secondary care, the government also introduced a new information system that enabled paperless referrals and appointment bookings and provided information on quality to help patients make more informed choices. This system, known as 'Choose and Book', allows patients to book hospital appointments online, with their GP, or by telephone. The booking interface gives the person booking the appointment the ability to search for hospitals based on geographic distance and see estimates of each hospital's waiting time.[12] Patient choice therefore signalled the end of selective contracting by encouraging movement of patients away from the local hospitals GPs had previously used.

---

[9] The NHS uses financial years which run April 1st to 31st March. In the description of the reforms we give precise dates of policy announcements. In the rest of the paper we refer to full years by the first calendar year which the financial year spans.

[10] The hospitals to which patients were sent were determined by the selective contracting arrangements made by the local NHS body responsible for purchasing health care (the Primary Care Trust, PCT).

[11] It was required that one of these choices was a private i.e. non-NHS provider. To facilitate this, the government placed contracts with a small number - around 15 initially, rising to around 30 by 2008 - of private sector providers of NHS care (known as Independent Sector Treatment Centres, ISTCs) that specialized in a limited range of elective procedures for which NHS waiting times were long (e.g., hip replacements and cataract removal). This policy turned out to be very limited. Neither the existing small private sector nor the new capacity were heavily used. In 2008 ISTCs accounted for less than 1% of all hospital activity. In April 2008, choice was extended so that patients could choose any hospital in England, as long as the hospital met NHS standards and was paid using the NHS ex ante fixed tariff (Department of Health 2007; Department of Health 2009).

[12] From 2007 the government also introduced a website designed to provide additional quality information to help patients' choices. This included information collected by the national hospital accreditation bodies, including risk-adjusted mortality rates and detailed information on waiting times, infection rates and hospital activity rates for particular procedures as well as information on hospital accessibility, general visiting hours and parking arrangements (Choose and Book website 2010).

The ex ante fixed prices are a case-based payment system known as 'Payment by Results' (PbR) (Department of Health 2002b). PbR is modelled on the diagnosis-related group (DRG) payment system used by the Medicare program and many private insurers in the US (Department of Health 2002a). A fixed price is set by the government for every procedure, with adjustments for whether a hospital is an academic centre, patient severity and local wage rates (Department of Health 2002a).[13] The price is exogenous to both the seller and buyer. The aims were that hospitals would only receive payment if they attracted patients (Le Grand 2007; Dixon 2004) and that fixed prices would mean that choice would depend on quality, and not price as in the previous system (Department of Health 2003).

In 2003 PbR was used for a very limited number of procedures (15 elective procedures) and only for purchases from a small group of hospitals (known as Foundation Trusts, FTs). In the following year it was extended to a wider set of elective and non-elective spells in FTs. In 2005 it was also applied to elective care (which accounts for approximately half of all hospital admissions) in non-FT NHS hospitals. In 2006 PbR was applied to almost all elective, non-elective and outpatient care (Farrar et al. 2007).[14]

In addition to Choose and Book and PbR, the government sought to give additional fiscal, clinical and managerial autonomy to NHS hospitals in order to further foster a competitive environment for hospitals. From April 2004 onwards high performing NHS hospitals could apply for FT status. This gave hospitals greater financial autonomy, allowing them to keep and reinvest surpluses across financial years. This represents considerable freedom over financial matters as non-FT NHS hospitals were required to break even on an annual basis and were heavily constrained in their access to capital. FTs were also given easier access to (primarily) private sources of capital. Hospitals earned this additional autonomy by performing well against key performance targets, the most important of which were good financial standing and the reduction in waiting times for elective care.[15]

[13] The payment unit is the HRG (Healthcare Resource Group), a group of diagnoses which utilize similar levels of resources. HRGs are very similar to DRGs used in the US.

[14] http://www.dh.gov.uk/en/Managingyourorganisation/Financeandplanning/NHSFinancialReforms/DH_077259 (accessed February 27, 2010).

[15] Granting of FT status is undertaken by an independent regulator, Monitor.

III. A Simple Model of Hospital Quality Competition

A. Introduction

We present a simple model of hospital quality that incorporates salient features of hospital markets in the UK that are relevant for the analysis of competition. These features are as follows. First, prices are set administratively (by regulators), rather than market determined. Second, hospitals are mostly public or not-for-profit firms. There are no for-profit hospitals in the NHS. Third, hospitals do not compete directly for emergency cases, such as emergency heart attack (AMI) patients, as such patients get taken to the nearest appropriate hospital.

Incorporating the last feature is a key contribution of our model. As noted above, AMI patients have been studied extensively in the empirical literature. Heart attack patients do not choose the hospital they are taken to, yet empirical studies find that heart attack mortality is lower in less concentrated (and presumably more competitive) hospital markets. How can the estimated effects be due to competition when heart attack patients who do not choose where they go? The prior literature is silent on this point.

Below, we build a model in which hospitals choose a level of quality for the hospital as a whole. They compete for elective patients, but not for emergency patients. Hospitals that are subject to tougher competition for elective patients choose higher quality, thus increasing the chances of survival of (emergency) heart attack patients, even though they are not competing for heart attack patients. This provides a coherent explanation for a causal relationship between hospital competition and AMI mortality rates.

B. The model

Let there be two types of patients: elective and emergency, $t = 1,2$. Elective (Type 1) patients choose their hospital based on the quality at that hospital, z_i . Elective patients can choose hospitals in advance. Examples include patients receiving cardiac bypass surgery, orthopedic procedures, or obstetrics. Emergency patients have no opportunity to choose, and are typically taken to the nearest hospital with adequate treatment facilities (and capacity). Thus Type 2 patients simply arrive at the hospital, irrespective of quality. Heart attack patients (AMI) are examples of Type 2 patients. We assume that the hospital chooses only one level of quality, which applies equally to both types of patients.[16]

Let quality have only a vertical dimension, i.e. 'more is better'. For simplicity in exposition, assume that the demand that any hospital i faces is separable in its market share, s_i, and the level of market demand, D. The demand for elective treatments is influenced by quality. Hospital i thus faces a demand from Type 1s of:

$$q_{i1} = s_{i1}(z_i, \mathbf{z}_{-\mathbf{i}}, N) D_1(\theta_1) \quad (1)$$

where s_{i1} is hospital i's market share in the elective market, z_i is i's quality, $\mathbf{z}_{-\mathbf{i}}$ is a vector of all other hospitals' qualities, N is the number of hospitals in the market (we assume all hospitals provide elective and emergency care), D_1 is market demand for elective treatments, and θ_1 is a vector of (exogenous) demand shifters for elective care (e.g., illness).[17] Assume that i's market share is increasing in own quality, decreasing in the number of firms, and that the responsiveness of market share to own quality is also increasing in the number of firms. We assume that market demand is determined by factors such as illness, but not by quality.[18]

The demand for emergency treatments is not affected by the quality of care the hospital chooses:

$$q_{i2} = s_{i2} \cdot D_2(\theta_2) \quad (2)$$

where s_{i2} is hospital i's market share in the emergency market, D_2 is market demand in the emergency market, and θ_2 is a vector of exogenous demand shifters in the emergency market.

Assume that hospitals all use the same technology and face the same input prices. Then they each have costs described by:

$$c_{it} = c\left(q_{it}, z_i, W\right) + F_t, \quad t = 1,2 \quad (3)$$

where $c(\cdot)$ is variable cost, W are exogenous cost shifters, and F_t is a fixed cost of entry.

[16] All that is necessary for our results is that there is a positive correlation between quality for Type 1s and Type 2s. We therefore assume a single quality level for simplicity in exposition. In this model hospitals directly choose the quality of service they provide. This is convenient for writing down a parsimonious model. In reality, hospitals may not directly choose the quality of care they provide. Instead they may choose effort, based on the incentives they face. Quality of care for a given patient is then determined by effort, among other factors. This distinction is immaterial for the purposes of our model, so we simply have hospitals choose quality directly.

[17] In the NHS consumers do not pay fees for health care, as a consequence demand does not depend on price. In a more general health care context consumers face a price less than the price received by the firm (e.g., in the US). We ignore this in order to keep this sketch of a model simple. It would not affect the conclusions in any event.

[18] This seems reasonable for the demand for hospital care. Most patients demand care because they are ill. If the quality of treatment were low enough that would cause some patients to choose to go untreated, but that seems unlikely in the context of the UK or other wealthy Western countries.

All hospitals in the NHS are public firms. Hospitals with Foundation Trust status are allowed to keep surpluses.[19] Hospitals that are not FTs do not retain surpluses, but are considered for FT status based on their performance, including financial measures. We therefore posit that NHS hospitals have an objective function that depends on profits and quality of care.[20] Quality of care because they care about their patients' wellbeing and for professional objectives (a high quality hospital confers prestige and status). Profits fund any other objectives the hospital may have.[21] For simplicity, let this function be additively separable in profits and quality and linear in profits:

$$U_i = u(\pi_i, z_i) = \pi_i + v(z_i). \qquad (4)$$

The hospital's profits are as follows:

$$\pi_i = \overline{p}_1 \cdot \left[s_{i1}(z_i, \mathbf{z}_{-\mathbf{i}}, N) D(\theta_1)\right] + \overline{p}_2 \cdot \left[s_{i2} \cdot D(\theta_2)\right] - c_1\left(q_{i1}, z_i\right) - c_2\left(q_{i2}, z_i\right) - F_1 - F_2 \qquad (5)$$

where $\overline{p}_1$ and $\overline{p}_2$ are the (fixed) prices the hospital receives for treating each patient type.

Substituting (5) into (4) and maximizing over quality gives us the following first-order condition:

$$\frac{\partial U_i}{\partial z_i} = [\overline{p_1} - \frac{\partial c_1}{\partial q_{i1}}]\left\{\frac{\partial s_{i1}}{\partial z_i} D_1(\cdot)\right\} - \frac{\partial c_1}{\partial z_i} - \frac{\partial c_2}{\partial z_i} + \frac{\partial v}{\partial z_i} = 0. \qquad (6)$$

Notice that the first-order condition for the choice of quality is affected by emergency patients (Type 2s) only through their marginal cost of quality.[22,23] The only difference with what the first-order condition would be for a profit-maximizing firm is the presence of the last term, $\partial v / \partial z_i$. Since this term is positive, the value that hospitals put on quality acts like a reduction in the marginal cost of producing quality, i.e., public hospitals will act like for-profit firms with a lower marginal cost of quality.[24] This implies that quality will be higher in equilibrium than if

[19] Approximately 50 percent of all NHS hospitals were FTs during this time period. The goal is ultimately for all hospitals to achieve FT status. As of 2010, 78% had achieved it.

[20] There have been many models proposed for hospitals (Newhouse, 1970; Lee, 1971; Pauly and Redisch, 1973; Lakdawalla and Philipson, 2006; Capps et al., 2010). While there is no agreement on a general model, most models posit an objective function which includes profits and some other argument, such as quantity or quality.

[21] This is general enough to include such divergent types of activities as research, community health activities, or perks for doctors or administrators. While we allow hospitals to have private objectives that may be socially wasteful (e.g., managerial perks), we do assume that they maximize whatever their objective function may be.

[22] If Type 1s were the only patient type, then quality would be higher, since Type 2s only affect marginal costs, not marginal revenues.

[23] We only consider an interior solution. It is possible that there is a corner solution in which the hospital chooses a minimum level of quality. In that case there will be no effect of competition.

[24] This is the same specification and result as for not-for-profit firms that care about quantity, as opposed to quality. See Lakdawalla and Philipson (2006) and Gaynor and Vogt (2003).

hospitals were for-profit. The comparative statics, however, are virtually the same as for a profit-maximizing firm.

The positive predictions of this model are clear. Conditional on price above marginal cost, quality is increasing in the number of firms in the market, that is, competition leads to more quality. This derives from the assumption that market share for elective patients becomes more responsive to quality as the number of firms increases (i.e. competition becomes tougher with the number of firms).[25] As competition gets tougher, hospitals choose higher quality in order to attract more elective (Type 1) patients. Since quality is hospital wide, this spills over into higher quality for emergency patients (Type 2s). Thus competition leads to higher quality for emergency patients, even though there is no direct competition for them. Further, increased demand responsiveness generally will lead to higher quality, for any number of firms in the market.

C. Expected hospital responses to the NHS reforms

The model above suggests that an increase in the elasticity of demand combined with a fixed price regime should lead to an improvement in hospital quality where hospitals face competition, and a larger increase in quality where hospitals face greater competition (conditional on prices above marginal costs). The question is whether the incentives facing the market participants in the English NHS are such that the reforms have this effect. There are reasons to expect this response.

First, NHS hospitals have incentives to respond to increased competition. While NHS hospitals are public organizations, the regime they operate under gives hospital managers strong incentives not to make losses. The government monitors the performance of hospitals on an annual basis and publishes summary assessments of their performance based on a range of indicators. These include measures of quality of care, access to care and financial performance. The weight given to financial performance in the summary assessments is high. Managers of hospitals which perform poorly in terms of the summary assessments may be replaced (and this does happen), while hospitals which perform well can get the greater autonomy awarded to

[25] This comparative static result is facilitated by assuming that market demand is unresponsive to quality. If market demand depends on quality, then quality is increasing in the number of firms if the responsiveness of market share to quality increases faster in the number of firms than market share itself decreases (with the number of firms).

Foundation Trust status. Further, hospitals with FT status can retain surpluses, and non-FT hospitals that perform well have the opportunity to earn FT status.

Second, PbR is a very highly geared reimbursement system. In 2006 over 60% of hospital income came from PbR payments (Department of Health 2007) and was projected to rise to about 90% in the following years (Street and Maynard 2007). The effect of PbR is to tighten the annual budget constraints for hospitals and increase the amount of uncertainty for hospital managers. Between 1997 and 2005, the use of annual contracts meant that annual revenues were known at the beginning of the year and costs were reasonably certain. Post-PbR revenues are more uncertain, as hospitals are no longer guaranteed volume at the start of each year. Supply has also been increased, as Choose and Book has opened hospitals up to competition from outside their local catchment area. Hospitals obtain revenues from patient volume (as prices are fixed) and, where rivals are present, have to compete for patients based on quality. Hospitals facing more competitors should have to compete more intensively for patient volume, and vice-versa.

Third, Choose and Book, by providing patients with greater choice and information, should increase the elasticity of demand facing hospitals. While increasing choice for patients might have little impact where patients have to make choices unassisted, the program is implemented by GPs. In addition to being mandatory, these physicians receive financial payments for the extra costs of implementing the system. Thus they have no reason not to offer their patients choice, other than their professional judgement. And while there is some evidence that not all primary care physicians thought that patients were able or wanted to make choices, a survey commissioned by the Department of Health found that 45% of patients recalled being offered a choice of hospital (Department of Health 2009). In addition, Dixon et al. (2010) found that the most important dimensions in patient choice of hospital were primarily measures of quality of care, such as hospital ‘superbugs’ (acquired infection rates) and cleanliness.

IV. Empirical strategy

Our goal is to test the hypothesis that the pro-competition policy improved hospital quality. To do this we exploit the variation in market structure across hospitals and examine whether quality is higher for hospitals in less concentrated markets after the introduction of competition. This is a difference-in-differences (DiD) approach. The simplest DiD strategy compares two groups over two time periods, where a treatment group is exposed in the second period and a control group is not exposed to the policy in either period. The NHS market-based reforms do not fit neatly within this simple DiD framework, as the reforms apply to all hospitals in England at the same time. However, the intensity of the competition induced by the reforms will vary according to market structure, which is a function of the geographical configuration of patient location and hospital sites. In some places population density results in a market structure which permits a high degree of choice. In others, population density is low and hospitals are located in highly concentrated markets with few competitors. As a consequence, post-policy a hospital in an unconcentrated market faces more exposure to the policy change than does a hospital in a highly concentrated market.

We therefore identify the impact of competition from the interaction of a treatment intensity variable, the degree of market concentration pre-policy, with a dummy for the post-reform year.[26] We use data from 2003 to capture the period before the policy change and data from 2007 for the period after the policy change. This gives the DiD regression specification:

$$z_{it} = \beta_0 + \beta_1 I(t=2007) + \beta_2 I(t=2007)*HHI_{i,\,2003} + \beta_4 X_{it} + \mu_i + \xi_{it} \qquad (7)$$

where z_{it} is the outcome variable, quality of care at hospital i at time t. $I(.)$ is an indicator function for the post-reform period, which takes the value 1 for financial year 2007 and 0 otherwise. $HHI_{i,2003}$ is the Herfindahl-Hirschman index pre-policy, our measure of market structure.[27] X_{it} is a vector of observed hospital characteristics which vary over time,[28] μ_i is an unobserved hospital fixed effect (which includes the level of the pre-policy market structure), ξ_{it} is random noise and t takes two values, financial year 2003 and financial year 2007. The DiD coefficient is β_2, which measures the change in the effect of market structure pre- and post-reform. Any common macro changes are picked up by the time dummy. This approach identifies a change in conduct due to the reform, the key identifying assumption being that without the

[26] Card (1992) was one of the first applications in economics to use a continuous treatment variable to estimate the impact of a policy. See also Angrist and Pischke (2008).

[27] The model from the previous section indicates that quality depends on market structure, but does not generate a specific functional form. We use the HHI as a convenient way of measuring market structure that is also consistent with the prior empirical literature in this area. We also note that we calculate HHIs using only elective care, not emergency care, as indicated by our model.

[28] In principle, X_i would contain the hospital's regulated price, marginal costs, etc. However, we do not have data on prices or marginal costs. Further, in practice we have only one year of data under the regulated price regime, which means we cannot estimate a parameter for price in the DiD framework. However, we do have a cost shifter, which we employ in regressions which augment the main specification. See Section VI.C.

policy intervention the trend in the outcome would have been the same whatever the market structure. Treatment induces a deviation from this parallel trend.

Endogeneity is a common concern in estimating regression models like (7) with the HHI on the right hand side (see, e.g., Bresnahan 1989). For example, if unobservably sicker patients go to better hospitals, which are in urban areas (and hence less concentrated markets), this would result in a negative correlation between the HHI and (poor) quality as measured, for instance, by mortality rates. On the other hand, if better hospitals have higher HHIs because of their higher quality, this would result in a positive correlation between the HHI and mortality. Our use of a short time series minimizes changes in populations or labor markets that may result in demand or supply changes. The use of a fixed effects estimator controls for the impact of any time invariant hospital-specific factors associated with quality, so that the levels of the outcomes may differ freely across hospitals in the pre-policy world. These hospital-specific factors include location, so we control for features that may be spatially associated with market concentration but are related to competition. In addition, we include in the X_{it} vector controls for observable time varying measures of the health of the patients admitted to the hospital and, in robustness tests, the health of the population in the catchment area of the hospital, as well as measures of local income, to control for patient health or income effects on demand.

However, there may remain concerns about unobserved heterogeneity. To deal with this, we instrument our measure of market structure with a measure of market structure based on factors unrelated to quality or unobserved patient heterogeneity. Following Kessler and McClellan (2000) and Gowrisankaran and Town (2003) we predict market structure on the basis of patient and hospital characteristics (patient distance from each hospital, patient demographics, patient illness severity, and size and teaching status of hospitals) and replace the actual 2003 HHI in (7) with the predicted HHI. This will depend only on these patient and hospital observables (in large part, patient distances from hospitals) and thereby eliminate possible correlation with the error in the quality equation. We discuss the construction of our predicted HHI in Section V.B and in Appendix B.

V. Data

We have assembled a rich database with hospital-level panel information on a variety of hospital quality and access to care indicators, financial performance, patient case mix and local area conditions. We use data on the universe of inpatient discharges from every hospital in the NHS in England for the financial years 2003 to 2007, comprising over 13 million admissions in around 240 hospitals per year.[29] We focus here on a (large) subset of these hospitals - short term general hospitals (called acute hospitals in the UK). These are the dominant suppliers of hospital-based services. Our data are derived from a large number of administrative data sources that we discuss briefly here and are presented in detail in Appendix Table A1.

Our sample selection criteria and the impact of selection on sample size are laid out in Table A2 in the Appendix. The population of acute hospitals falls slightly from 180 in 2003 to 175 in 2007 (due to hospital reorganization by the government to deal with longer term changes in population density). Our first selection rule is to select hospitals with at least 5,000 total admissions to ensure we are examining non-specialist hospitals. Second, we drop those hospitals for which mortality data or the data necessary to calculate HHIs are not available for both years. Our final sample contains 162 hospitals for each year of the analysis, totalling 324 hospital-year observations for the main analyses. For our analysis of AMI mortality we also exclude hospitals with fewer than 150 AMI admissions to avoid the problem of variability of rates from small denominators (see e.g., Kessler and McClellan 2000). This reduces the number of hospitals to 130 in 2003 and 121 in 2007, giving 251 hospital-year observations for the emergency AMI analyses.

A. Measures of hospital quality

We use mortality rates both within the hospital and including deaths post-hospital discharge as indicators of quality. These are derived from Hospital Episode Statistics (HES) data, which are administrative data on every NHS health episode such as an operation or physician consultation. We use deaths following emergency AMI and following all admissions. For emergency AMI, we use hospital-level annual deaths within 30 days in any location for patients

---

[29] Hospitals are called 'trusts' in the UK. An NHS 'trust' is a financial, managerial and administrative unit and may cover more than one physical hospital, all of which are located closely in geographical space as there are no hospitals chains in the NHS. We use the term 'hospital' rather than 'hospital trust' for expositional convenience.

aged 35-74 years old. These data are constructed by the organization which develops NHS quality indicators (National Centre for Health Outcomes Development, NCHOD). For all-cause mortality, which is not published by NCHOD, we construct annual data at the hospital level for 28 day in-hospital mortality rates for all admissions from HES.[30]

Deaths following emergency admission for AMI have been published by both the US and UK governments as indicators of hospital quality. There is not competition for emergency admissions of AMI patients, who are taken in the UK to the nearest hospital with capacity. But AMI mortality is treated as an indicator of overall quality in the hospital for a number of reasons. First, the infrastructure used to treat AMI is common to other hospital services, making it a good general marker of hospital quality (Gaynor 2006).[31] This is in accord with our theoretical model. Second, AMI admissions are reasonably high volume and mortality is a fairly common outcome so variability in the rates is less of an issue than for other treatments. Third, as all patients with a recognized AMI are admitted there is little scope for selection bias to affect the decision of who gets admitted. In addition, the use of within 30 day mortality in any location allows us to examine whether hospitals respond to competition by discharging patients in a poorer health state.[32]

We also use the all-cause in-hospital mortality rate, as studies have found falls in overall hospital mortality linked to clinical and managerial quality improvement programmes and to variability in hospitals' performance across a number of conditions (Wright and Shojania 2009; Jha et al. 2005).

In addition to these measures of quality, we also examine the impact of competition on other aspects of performance. These are the length of stay, the total number and mix of admissions, total expenditure and a simple measure of (lower) productivity, expenditure per admission.

Table 1 presents the means, standard deviations, minima and maxima for all the variables used in our main regressions and our robustness tests (Table A3 in the Appendix presents the within and between variation). The average hospital in our estimation sample admits just under 68,000 patients and has 412 emergency AMI admissions a year. 6.9% of those aged 35-74 die

[30] We also calculate deaths following emergency AMI admissions for all ages in order to allow us to examine all-cause in hospital mortality excluding AMI admissions.

[31] Many of the actions to reduce deaths from emergency admissions for AMI need to be taken soon after an attack, and so the performance of a hospital in terms of AMI reflects the performance of its A&E department (accident and emergency department, or emergency room).

[32] These data are constructed by linking information on deaths following discharge (from the UK Office for National Statistics, ONS) to the admitting hospital (from HES data, The NHS Information Centre).

within 30 days in either the hospital or the community. 1.6% of all patients admitted die in the hospital within the first 28 days after admission. However, there is wide variation in these rates between, and within, hospitals. Around 30-40% of the variation in the within 30 day AMI death rate and the all-cause mortality rate is within hospitals.

B. *Measures of hospital market structure*

We measure market structure at the hospital level using an HHI based on patient flows to each hospital. The HHI is built up from patient flows at neighborhood level and is calculated in two steps. In the first, the HHI in each geographically defined neighborhood in England is calculated as the sum of squared patient shares across all hospitals the neighborhood sends its residents to for all elective care.[33] The neighborhood definition we use (the MSOA) contains on average around 7,000 persons and so is similar, or smaller, in population, to a US zip-code.[34] We allow the market to be the whole of England (i.e. we include all hospitals used by each MSOA in the calculation of patient shares). In the second step, the HHI for each hospital is calculated as a weighted average of the HHIs for the neighborhoods it serves, where the weights are the shares of the hospital's patients that live in each neighborhood. Thus each hospital has its own market. Patient flows are from the information on admissions and patients' locations in the HES dataset. In what follows, we refer to HHIs based on actual patient flows to the hospital as *actual* HHIs. We calculate these for both 2003 and 2007 and use the 2003 HHI as our measure of market structure.

As noted above, there is concern in the industrial organization literature about the potential endogeneity of concentration indices when used as regressors for the purpose of making inference about competition. In the health care context, Kessler and McClellan (2000) have argued that measures of hospital competition based on actual patient flows are potentially endogenous in regressions with hospital quality as the outcome of interest, due to potential correlation of flows with the unobserved characteristics of patients, hospitals or geographic

[33] Elective admissions make up around half of all hospital admissions and these are the ones subject to Choose and Book.

[34] MSOAs (Middle Layer Super Output Areas) are defined to ensure maximum within MSOA homogeneity of population type. In England each of the 6,780 MSOAs has a minimum population of 5,000 residents and an average population of 7,200 residents. Data retrieved from http://neighborhood.statistics.gov.uk/dissemination/Info.do?page=userguide/moreaboutareas/furtherareas/further-areas.htm on 26/04/2010.

markets. To avoid this problem we follow Kessler and McClellan (2000) and use HHIs based on patient flows that are predicted using only exogenous patient and hospital level characteristics. To generate predicted patient flows we first estimate multinomial logit patient level hospital choice models and then derive the predicted probabilities that a given patient attends each hospital in their choice set (for details, see Appendix B). These predicted probabilities are used to calculate *predicted* HHIs for each hospital using the same method described above. The summary statistics in Table 1 show that predicted HHIs tend to be lower than actual HHIs, i.e. markets are less concentrated when HHIs based on predicted flows are used instead of actual patient flows.[35] This suggests that patient flows are likely to be influenced by potentially endogenous factors - such as unobserved hospital quality - leading hospital markets to appear to be more concentrated than they would otherwise be. So the use of predicted HHIs based on exogenous hospital and patient characteristics means our estimates of the impact of market structure on hospital quality are less likely to suffer bias arising from endogeneity between hospital quality and actual patient flows. In robustness tests we also use an HHI measure based on actual flows.

C. Other controls

As many potential control variables may be endogenous (for example, admissions, financial position, or staffing), in our primary estimates we use a very limited number of time-varying controls. In all specifications, to allow for differences in the health of hospitals' patient mix (often referred to as a hospital's 'case mix'), we include hospital fixed effects, which will pick up observed and unobserved time-invariant differences between hospitals, and the age-gender distribution of total admissions (cause-specific admissions for emergency AMI) through the proportions of admissions in five year age bands for men and women (36 variables). In the UK context this has been shown to do a good job of controlling for case mix (Propper and Van Reenen 2010). It is possible that there are omitted variables which account for cross-hospital heterogeneity and could be correlated with our included measure of market structure. To check for this we perform a large number of robustness tests in which we employ additional variables

[35] The correlation coefficients between predicted and actual HHIs in our estimation sample are 0.73 and 0.70 for the years 2003 and 2007 respectively (both statistically significant at the 1% level).

to control for any further heterogeneity across hospitals. The variables used in our robustness tests are described in Section VI.

D. Patterns in the data

We now examine some simple patterns in the data in order to see if they are consistent with responses implied if the reforms induced competition.

Did the reforms result in less concentration?

The policy introduced the potential for competition. If the policy did lead to competition, then patient admission patterns likely should have changed, leading to changed market concentration. To establish whether patient flows changed post-policy we examine measures of market concentration pre- and post-policy. Figure 2 presents the kernel density estimates of the distribution of the actual HHI at the hospital level for 2003 and 2007. The figure shows a clear leftward shift in the distribution of HHI levels over the time period so that in 2007 the level of concentration faced by hospitals had fallen at virtually all HHI levels, with the bulk of the change in the middle of the distribution.

To show the spatial distribution of market concentration, the left hand panel of Figure 3 plots the location of hospitals and their concentration levels in 2003, divided into quartiles of the actual HHI. Darker blue dots represent more concentrated markets. As expected, hospitals in the least concentrated markets pre-policy (those in the bottom quartile of the HHI distribution, light blue dots) were largely located in the more densely populated urban localities, particularly in the Greater London and Manchester areas, while hospitals in the most concentrated markets (those in the top quartile of the HHI distribution) tended to be located outside urban centres. However, *changes* in concentration were not confined to the cities. The second panel of Figure 3 shows that some hospitals located in the largest urban areas experienced large decreases in concentration (dark red dots indicate largest change, yellow indicates least change), but many of the hospitals which experienced the largest decrease after the implementation of the pro-competition policies of the 2000s are actually located around, rather than in, urban areas. Large decreases in the level of concentration faced by hospitals have therefore occurred both in the more densely populated areas, where the market structure was already relatively unconcentrated in the pre-policy period, and in more rural areas where market structure was more concentrated.[36]

Did patients respond to the reforms?

One of the intentions of the reforms was to change the patterns of care seeking by patients.[37] If the reforms were successful we would expect to see this reflected in the data. We examine this in Table 2. The top panel shows the change in patient care seeking post-reform by hospital quality. For our measure of quality we use 2003 AMI mortality rates to reduce the likelihood of simultaneous determination of mortality and patient volume. If patients became more responsive to quality post-policy we should see better hospitals (those in the bottom quartile of the mortality distribution) attracting more patients relative to worse hospitals (those in the top quartile). That is exactly what the data show. The total number of patients increased overall, but better hospitals experienced a larger increase in elective patients than did worse hospitals. Although the distances patients travelled for care rose similarly for worse and better hospitals, the share of patients bypassing their nearest hospital increased for better hospitals while it clearly *decreased* for worse hospitals. This provides reassurance that there is a patient response to quality and that it increased during the reform.

The lower panel of Table 2 shows changes in patient care seeking after the reform by levels of pre-reform market concentration. Our estimation strategy is based on potential exposure to policy change, as measured by market concentration. If the policy had an effect, we would expect hospitals with greater potential exposure - those located in markets which pre-policy were less concentrated - to experience more changes in the patterns of patient care seeking. Table 2 shows this is the case. Hospitals in markets in the bottom quartile of the HHI distribution had a greater increase in elective admissions, a greater increase in the distance travelled by their patients and a greater change in the share of patients bypassing their nearest hospital.

[36] The correlation between levels of HHI in 2003 and changes in HHI between 2003-2007 is -0.09 (p-value = 0.250) indicating that changes in competition levels after the reforms occurred for hospitals in both more and less concentrated markets pre-policy.

[37] We note that decisions about where to seek care are likely to be the product of patient and family preferences and doctor advice. The identity of the decision maker is not critical here. What matters is whether decisions about where to go respond to quality.

VI. Results

Figures 1a and 1b suggested that the introduction of pro-competition reforms reduced the AMI and all-cause death rates in markets where the policy could have more effect: for those hospitals operating in markets with lower levels of concentration. In this section we formally test this using equation (7) to estimate the effect of the policy. We begin by looking at the impact of market structure on quality (as measured by death rates). We then examine the impact of the policy on the volume and composition of patients treated and on simple measures of productivity, subject our results to a wide set of robustness checks, and present estimates of the financial magnitude of the effects.

A. The impact of market structure on quality

Table 3 reports our DiD estimates of the impact of market structure on hospital quality.[38] The estimates control for year effects (a 2007 year dummy), the age-gender composition of admissions, and hospital fixed effects. Column (1) presents estimates for the 30-day mortality rate following emergency AMI admission. Concentration has a statistically significant positive effect on mortality, i.e. higher market concentration (a larger HHI) leads to lower quality. A 10% increase in the HHI leads to an increase of 2.91% in the AMI death rate.[39]

Column (2) presents the DiD estimate for the all-cause mortality rate. The estimate again shows a significant relationship between quality and market concentration. The magnitude is smaller than that for AMI but precisely estimated. To test whether the estimated effect for all-cause mortality is driven only by AMI deaths, column (3) presents the DiD estimate for all-cause in-hospital mortality excluding deaths after AMI admissions. The coefficient is almost the same as when AMI deaths are included, indicating that there is an effect from the policy on both the AMI death rate and the death rate following all other admissions. The larger effect we find when quality of care is measured by the AMI mortality rate rather than all-cause mortality rate is likely to be due, at least in part, to the fact that many conditions that result in death in the hospital are

[38] Full estimation results are in Appendix Tables A4, A5 and A6.

[39] In principle, we should take account of the fact that the predicted HHI is estimated in calculating the standard errors. However, the predicted HHI is constructed from a patient choice model estimated on the population of hospital elective admissions, which number 6.5 million in 2003. As a consequence, there is little sampling variation to account for. To be conservative, we examined this empirically. We generated ten bootstrap samples of hospital elective admissions for each year, estimated the patient choice model on each sample, then constructed predicted hospital HHIs (as described in Section IV). The intra-hospital correlations between the bootstrapped predicted HHIs was very high (0.9977 for 2003) indicating that there is indeed very little sampling variation.

not responsive to better quality health care.[40] Therefore, the effect of improvements in hospital care quality (driven by the pro-competition reforms) is likely to be larger when quality is measured by AMI mortality rates than when quality is measured by the overall death rate, as the latter includes several conditions for which mortality is less (or not at all) affected by health care quality.[41]

The coefficients are statistically significant but the estimated magnitude of the response is relatively modest. A 10% fall in the HHI is associated with a fall in the 30 day death rate following AMI admissions by 2.91%. This amounts to 1/5th of a percentage point at the mean AMI death rate of 6.9%, implying a little over 8 fewer AMI deaths annually per hospital, or approximately 1,000 fewer total deaths per year over all 135 hospitals in our sample.

Our estimated magnitudes are similar to those from some other relevant studies. Kessler and McClellan (2000) estimate that a move from the top quartile to the bottom quartile of the HHI in their sample will lead to a 3.37 percentage point fall in the AMI death rate. The equivalent figure using our estimates and data is 2.26 percentage points.[42] Cooper et al. (2010, p.27) find that a one standard deviation increase in their measure of competition for English hospitals is associated with a 0.3 percentage point reduction in the 30 day in-hospital AMI mortality rate (per year) following the NHS pro-competition reforms. Our results imply a similar 0.21 percentage

[40] Nolte and McKee (2008) report that 'amenable mortality' - deaths from causes that should not occur in the presence of timely and effective health care - accounted for around 27% of total mortality for males aged under 75 years in the UK in 2002-03, and 33% for females. On the other hand, this study (and other studies based on systematic reviews of published clinical evidence, see Nolte and McKee (2004) for a review) includes ischemic heart disease - and its AMI component - among the causes of death which are amenable to better health care, with around half of such premature deaths considered to be avoidable by factors like better management of the condition within the hospital.

[41] We also examined another measure of quality - the MRSA rate - and two access measures - the share of patients waiting more than three months and share of attendances spending more than four hours waiting for care in the emergency room (A&E department). Using the same specifications as those presented in Table 3, the coefficients (standard errors) on these estimates are -0.027 (0.108), 0.028 (0.174) and -0.002 (0.011). Thus none of these outcomes are affected by the policy change. MRSA rates are highly influenced by changes in behavior in the community as well as hospital policy (see e.g., Ferry and Etienne 2007 and references therein) and so may not respond strongly to hospital-level attempts to reduce them. Waiting times had been the target of a major policy campaign between 2000 and 2005 and had fallen substantially by 2005 (Propper et al. 2010), perhaps leaving relatively little scope for further reductions in response to the competition policy.

[42] See Appendix 3 in Kessler-McClellan (2000), which presents the results for the specification in which HHI enters linearly. The measure of mortality in that paper is the one-year AMI mortality rate for US Medicare beneficiaries. The comparable estimated effect for our sample is derived as follows. A 43 unit decrease in HHI (=1% in our sample) leads to a 0.02 percentage point decrease in AMI deaths at the mean (= 0.291%*6.9%). So a unit decrease in HHI leads to 0.02/43 = 0.00046 percentage point decrease in AMI deaths. The difference in mean HHI between the top and bottom quartiles in our sample is 4,854.6. This equals a 2.26 percentage points (= 4854.6*0.00046) fall in the death rate.

point reduction for each year post-policy (2004 to 2007).[43] We discuss the economic significance of our estimates at the end of this section.

B. The impact of market structure on other aspects of performance

We also examine whether the reform had an impact on resource use. In Table 4 we examine the mean length-of-stay of admitted patients (LOS), the total number of admissions, the number and share of elective admissions, expenditure and a simple measure of (lower) productivity, expenditure per admission.

Column (1) indicates that increases in concentration are significantly linked to a rise in the length-of-stay. The estimated coefficient implies that a 10% fall in a hospital's HHI on average results in a 2.3% fall in length-of-stay. At the mean length-of-stay in the sample of 1.2 days this is just under an hour.[44] The policy does not seem to have affected total number of admissions or their composition (columns (2)-(4)). Nor did the policy result in any change in either total hospital operating expenditure or expenditure per admission, so we do not find evidence that resource utilization increased in less concentrated markets following the reforms.[45] Taken together, the findings for quality (Table 3) and resource utilization (Table 4) suggest that hospitals facing more competitive pressure were able to find ways to marshal resources more efficiently to produce better patient outcomes.

C. Robustness checks

To avoid inclusion of potentially endogenous variables our estimates above control only for time-invariant factors at the hospital level and a simple measure of case mix. It is possible that there are omitted variables potentially associated with market concentration that are driving our results. To examine this we undertake a wide set of robustness tests for our AMI, all-cause mortality and length-of-stay results. These are presented in Table 5. All cells report the DiD

[43] The effect in our sample is derived as follows. Using our estimated AMI coefficient, a one standard deviation reduction in HHI (=1,921 units, a 41% reduction) from the mean implies an 11.9% reduction in AMI deaths for the post-policy period, or 0.82 percentage point (our mean AMI death rate is 6.9%). This is equivalent to a 0.21 percentage point decrease per year in AMI mortality for the period 2004-2007 (=0.82/4).

[44] Farrar et al. (2007) find that the application of PbR was associated with reductions in the average length-of-stay ranging between 1.2%-2.3%, depending on the comparison group used.

[45] We also find no policy effect on a simple measure of labor productivity (number of admissions per clinical staff). The DiD coefficient (standard error) is -0.014 (0.026).

estimates from separate regressions. The first row presents the baseline results from Table 3, columns (1) and (2), and Table 4, column (1).

Placebo test and specification of the DiD estimator

If our results are being driven by pre-existing observable or unobservable differences between hospitals facing different levels of market concentration, we would expect to find significant estimates if we compared hospitals facing more and less concentrated market structure *before* the reforms were introduced. To test this we undertake placebo DiD tests by estimating the same specifications as in Tables 3 and 4 using data from before the reforms.[46] Row 2 in Table 5 presents the results using 1999 as the placebo 'policy-off' year and 2003 as the placebo 'policy-on' year. None of the estimated coefficients are even close to being statistically significant at conventional levels and their magnitudes are very small compared to the baseline estimates. This suggests that our results are driven by the reforms rather than due to pre-existing differences between hospitals.

In our model, identification comes from the interaction of pre-policy market structure with the policy change. This is in contrast to much of the previous literature, which estimates the impact of competition in health care from cross sectional as well as time series variation, using instrumented measures of market concentration. We therefore estimate a specification of the DiD in which we estimate the impact of the policy from both cross sectional and time series variation in exposure to the policy. With the inclusion of hospital fixed effects, this specification controls for the change in market concentration and thus for unobserved heterogeneity that is correlated with a change in market structure. We estimate:

$$z_{it} = \beta_0 + \beta_1 I(t=2007) + \beta_2 I(t=2007)*HHI_{it} + \beta_3 HHI_{it} + \beta_4 X_{it} + \mu_i + \xi_{it} \qquad (8)$$

Row 3, Table 5, presents the DiD estimates from equation (8). It is clear that these are very close to those from our base specification.

As a further specification test, we estimate equation (7) using actual HHIs rather than predicted. Row 4 of Table 5 indicates that our use of predicted HHIs provides more conservative estimates for all-cause mortality and LOS. The magnitude of the impact of HHI on AMI

[46] A few hospitals used in our main regressions were not active before 2003 so there is a small change in the sample size compared to the baseline estimates. Due to data limitations, the all-cause mortality variable for the placebo test is the in-hospital mortality rate at any point during a hospital stay and the corresponding test for that variable uses the shorter period 2001-2003.

mortality is similar using actual and predicted HHIs, but imprecisely estimated using actual HHI. As a final test of the market concentration variable, we replace the continuous measures of market concentration in 2003 in (7) with a dummy for a hospital being in the top quartile of the HHI distribution in 2003. Row 5 indicates that the results for AMI and all-cause mortality are driven by the poorer performance of hospitals operating in more concentrated markets.[47] We conclude that our results are generally robust to the exact specification of the DiD model.

Differences in costs, wages, and the financial position of the hospital

In our main specification the hospital fixed effects control for any constant differences across hospitals in costs and financial position (or anything else). They do not, however, control for any time varying cost differences that may exist. If those differences are uncorrelated with the HHI, this is unproblematic. However, it is possible that such a correlation exists. Specifically, hospitals in large urban areas tend to have higher costs and are also in less concentrated markets. Higher costs will lead hospitals to choose lower quality, so omitting costs could lead to a downward bias in the estimated effect of the HHI on quality.

While we do not have data on prices, including a variable for costs indirectly controls for prices. The reason is that costs have an impact only relative to price. If cost differences across hospitals are fully compensated by price differences then the cost variable in the regression will have no impact. Our measure of hospital cost differences is the Market Forces Factor (MFF). The MFF is an adjustment factor calculated by the NHS to capture geographic differences in hospital costs (specifically input prices: land values, staffing costs, etc.). Before the reform local health authorities' budgets for purchasing hospital care were adjusted based on the MFF. After the reform the MFF was applied to regulated prices, so that hospitals in high cost areas have their HRG prices adjusted upwards and vice-versa.

To test whether omitted hospital cost differences have biased our results we add the MFF as a control variable. Row 6 in Table 5 reports the estimates. As can be seen, there is very little impact on the estimated magnitudes of the DiD coefficients or their standard errors.

Propper and Van Reenen (2010) show that higher outside wages reduce the quality of labor a hospital can attract and thus reduce the quality of care. Therefore as a further control we

[47] The mean HHI in the top quartile is 7,155. One can calculate the number of equal size firms implied by a value of the HHI using the formula 1-log(HHI/10,000). This gives us an implied mean value of 1.3 hospitals for the top quartile. Therefore our test is very close to comparing monopoly vs. non-monopoly markets.

introduce the level of wages in the outside labour market, as measured by the average male wage in the area.[48] The results including the outside wage are reported in row 7 of Table 5. The results show our main estimates are robust to including the outside wage, although the all-cause death rate estimate is slightly reduced.

Our main models do not control for hospitals' financial surpluses or deficits, since these may be endogenous. Hospitals with higher quality may admit higher volumes of patients, resulting in higher surpluses, and vice-versa. However, the switch to PbR may give hospitals an exogenous large income shock. Under PbR hospitals get a price per procedure equal to the average cost of all hospitals (adjusted for local cost conditions, as described above). This price may be far from the hospital's own cost. As PbR payment accounted for a large proportion of hospital revenue, some hospitals may have experienced large income shocks (positive and negative) when full PbR was rolled out in 2006. It is possible that our results may be driven by this income shock rather than changes in market concentration if the income shock is positively correlated with concentration (that is, hospitals in concentrated markets received positive income shocks, and those shocks were larger than those received by hospitals in unconcentrated markets). To test this, we include an additional control for the hospital's financial position as measured by the end of year surplus/deficit. Row 8 of Table 5 presents the estimates including this control. The DiD point estimates change very little. We therefore conclude that the response we see is not due to differential price-cost mark ups or gearing associated with PbR.

Controls for patients heterogeneity and number of admissions

In our main regressions we control for differences in case mix between hospitals using the shares of admissions within 5 year age-gender bands and a full set of hospital dummies. To alleviate the concern that our estimates are biased by different patterns of case severity across hospitals facing different market structure (for instance, by more severely ill patients systematically choosing high-quality urban hospitals post-reform) we add, separately, two further controls for patient composition. The first is the health of the local population as measured by the

[48] We use the average of the median full-time gross wages for male workers (all occupations) in local authorities (units of local government) within a 30 kilometers radius of the hospital. Male wages are highly correlated with female wages over time within local authorities.

standardized mortality rates of the population in the catchment area of the hospital.[49] This is a measure of the potential rather than actual patient mix so should suffer from less endogeneity bias than actual case mix. The results in row 9 show that again our estimates are little changed by this control, though it does bring down the estimates on within hospital all-cause mortality. This is perhaps to be expected as the mortality rate of the population in a hospital's catchment area will be affected by the all-cause in-hospital mortality rate.

The second control is a direct measure of the severity of patients treated in the hospital, the Charlson index, widely used as a marker of patient severity.[50] We do not use the index in our main regressions because of concerns over endogeneity and that it may pick up-coding responses to the PbR system.[51] The results in row 10 of Table 5 show that the DiD coefficients change little after inclusion of this measure. We conclude that our estimates are robust to observed changes in case mix controls and so reducing concerns that changes in unobserved heterogeneity in patient severity may be driving our results.

We have shown that there is no differential change in total volume by market concentration. There is evidence, however, that shows that patient health outcomes are better at hospitals that treat a large volume of patients (see Gowrisankaran et al., 2006; Gaynor et al., 2005). While it is not likely that a volume-outcome relationship drives our results, we nonetheless check this by adding a control for the number of cause-specific admissions. Row 11 indicates that, as expected given the results in Table 4, our results are little affected by this control.

Local area economic conditions

Between 2003 and 2007 areas with higher and lower hospital concentration may have experienced different economic growth rates. Recent research has suggested that economic growth can adversely affect AMI outcomes, with a greater number of heart attacks being observed during upturns in the business cycle (Ruhm 2006). It could therefore be argued that the impact of market concentration that we find is not due to the result of poor hospital quality but to

[49] This is constructed from data on 353 Local Authorities (LA) and standardized for age and gender. The hospital-specific area standardized mortality rate is an inverse distance-weighted average of the figures for all LAs. Data sources for this and all other covariates are listed in the Appendix (Table A1).
[50] The Charlson index is an index of severity of illness based upon a patient's diagnoses and procedures, including 19 co-morbidities. These are aggregated using weights derived from estimates of the co-morbidities' contribution to predicting mortality.
[51] DRG systems have been argued to give incentives for greater coding of patient severity in order to move patients into categories which have higher reimbursement (see e.g., Ellis and McGuire 1986; Dafny 2005).

smaller falls in fatalities due to higher rates of economic growth in more concentrated markets. A test of this is to control for a time varying measure of economic activity at the level of the hospital's catchment area. This is essentially the test in row 7 of Table 5 which controls for male wages, which we have shown slightly decreases the point estimate on AMI deaths. It may therefore be the case that some of the AMI death rate effect comes from a rise in heart attacks following an upturn in the cycle.

In addition to the potential effect of the economic cycle on health status, economic growth might also directly affect hospital emergency outcomes. This is particularly relevant for deaths following AMI admissions. Stronger economic activity is likely to generate increased traffic flows and road congestion, thus potentially increasing the time elapsed between the heart attack and hospital arrival ('floor to door' time), so decreasing the chances of patient survival. To address this we add a control for ambulance speeds to our estimates for AMI. The government monitors the proportion of urgent and life-threatening ambulance calls which arrive at the scene of the incident within eight minutes and we use this independent assessment. The results in row 12 show that our market concentration impacts on AMI mortality are robust to this additional control. These tests provide reassurance that our market concentration effects are not attributable to differential economic growth rates across localities.

Possible contamination from other policy changes

Our estimation strategy exploits a policy change. We therefore need to check that the change we observe is due to the pro-competition policy rather than other policies that might have been running at the same time. One potential policy candidate, which was part of the pro-competition reforms, was the attempt to increase the supply of non-NHS facilities through the offer of NHS guaranteed payments (regardless of actual volume) to non-NHS entrants of specific types of (elective) care for which there were long NHS waiting lists. These entrants were known as Independent Sector Treatment Centres (ISTCs). This policy began in 2003 but did not achieve the entry that was initially hoped for. Even by 2008 the volume of admissions in ISTCs was only equal to 1% of NHS admissions. However, it is possible that these centres might have sharpened the competitive pressure on NHS hospitals. To test this we control for the number of ISTCs in each NHS hospital's markets.[52] The results are presented in row 13 of Table 5 and show that this also has little effect on the estimates.

There is a second possible confounding candidate, the establishment of cardiac networks to improve the treatment of cardiac patients (Minap 2010). This policy started in 2001 and was in operation whilst regulated prices and mandated patient choice were introduced; while having no element of promotion of competition for patients, the policy was geographically defined. These networks sought to improve treatment of patients through the promotion of specialist units that undertook angioplasty (PCI) as the first treatment for AMI patients (generally replacing thrombolytics), faster use of thrombolytics and ambulance response times, and greater prescription of beta blockers, aspirins and statins on discharge. Performance in these dimensions was published at hospital level annually.

The first two activities are geographically differentiated. PCI as the first type of treatment is only possible in urban areas and the use of thrombolytics in ambulances therefore increased more in rural areas. It is therefore possible that the better AMI survival rates in less concentrated markets are due to the operation of these networks rather than the competition policy.

To test this we first examine the correlation of market concentration in 2003 with the change in use of PCI, thrombolytics, speed of ambulance arrival and use of appropriate drugs on discharge between 2003 and 2007 (at hospital level).[53] We find that there is indeed a significant association: the increase in use of thrombolytics is positively associated with high market concentration and in use of primary PCI with low levels of concentration.[54] However, these associations do not drive our estimates of better outcomes in less concentrated areas. Row 14 presents our AMI results controlling for all the cardiac treatment measures. The estimate shows that our results are actually stronger when we control for differences in treatment. Analysis of the coefficient estimates on the treatment variables (not shown) shows this is mainly driven by

[52] We do not have accurate estimates of the actual volumes of patients treated in ISTCs and therefore measure the competition from ISTCs for NHS hospital *i* by the number of ISTCs located within 30km of *i*.

[53] We measure the cardiac treatment variables by the annual performance against the standards published by the body which promoted this policy (Minap). The treatment variables used in our analyses are the shares of AMI patients: (1) having thrombolytic treatment within 30 minutes of arrival at hospital; (2) having thrombolytic treatment within 60 minutes of calling for help; (3) having primary angioplasty (PCI); (4) discharged on aspirin; (5) discharged on beta blockers; and (6) discharged on statins. For variables (4)-(6) we use data for year 2004 as this is the earliest available year.

[54] The estimated OLS coefficient (standard error) on HHI for the change in the percentage of patients having thrombolytic treatment within 60 minutes of calling for help is 4.770 (0.871). For the change in the percentage of patients having PCI the corresponding estimates are -3.707 (1.401).

response rates for thrombolytic treatment. The share of patients who have thrombolytic treatment within 60 minutes of calling for help increased more between 2003 and 2007 in more concentrated markets (rural areas). Controlling for this, hospitals in more concentrated markets have a higher increase in AMI death rates, i.e. a lower increase in quality.

Weighting and functional form

We exclude hospitals with low volumes of admissions to ensure our results are not driven by the variability in rates induced by low denominators. To test that our results are robust to this exclusion we re-estimated the models including all observations, but weighted our regressions by either the total number of admissions (for the all-cause mortality rate and length-of-stay) or the cause-specific admissions (for the AMI mortality rate). Row 15 of Table 5 presents these estimates. The coefficients for all-cause mortality and length-of-stay are virtually unchanged, while that for AMI falls by around 35% but remains statistically significant. The results suggest that hospitals with small volumes of AMI admissions are different from those with higher volumes, but we still find a statistically significant effect even after inclusion of these hospitals.

Finally, in row 16, we test the robustness of our results to our functional form assumptions and estimate the models with both the dependent variable and HHI in levels instead of logs. We report the implied elasticity estimates from this specification to enable comparison with the baseline results. The table shows that our estimates are very similar in levels.

Overall, we conclude that our results for the impact of the reforms are robust to a wide range of checks.

D. Did the policy matter?

To provide a better sense of the economic significance of the reforms we undertake some simple back of the envelope calculations. The first is to calculate the benefits in monetary terms from the observed change in market structure following the reforms. We calculate the value of the gain in life years that would arise from a change in the HHI equal to the observed (average) decrease in HHI between 2003 and 2007 (this equals 118 units). Using the estimated coefficient from Table 3, column (2), the average hospital would experience a 0.3% fall in its overall mortality rate from this decrease in the HHI.[55] The average age of death of patients in hospital is 77 years. A 77 year old male in Britain has an additional life expectancy of 9.5 years and a female has an additional life expectancy of 11 years. Using our estimate that 0.3% of these deaths are averted and combining it with these extra years of life leads to an estimated 4,791 life years saved.[56] If we adopt the $100,000 benchmark of Cutler and McClellan (2001) for the value of a year of life, the beneficial effects of the pro-competition reforms are on the order of $479.1 million, or approximately £298 million.[57]

This calculation gives the average benefit associated with the policy change. A second calculation is to derive the cost of being in a concentrated compared to an unconcentrated market. We therefore compare the difference in life years saved for a hospital located in a market at the average concentration versus one with low concentration, defining low concentration to be an HHI one standard deviation below the mean. This is a difference in HHI of just under 2,000. Using the estimated coefficient from Table 3, column (2), a hospital in the lower HHI market would have 4.4% fewer deaths per year. Using the same methods and numbers as above this translates into 78,318 more life years saved, with a monetary value of $7.8 billion, or £4.8 billion.[58]

As a basis for comparison, the annual NHS budget is of the order of £100 billion. The estimate of the immediate impact of the policy (£298 million) is approximately 3-10ths of one percent of the NHS budget. While this is small, it is not trivial, as it represents the short run impact of the policy immediately after implementation. Moreover, we only enumerate the gains

---

[55] A one percentage change in HHI is equal to 43.5 units, so a fall of 118 would lead to a change of 0.099*118/43.5 = 0.3%.

[56] The calculations of lives saved were made separately for each hospital and then aggregated up, weighting by the hospital's number of admissions in 2007. The average age of death of 77 years is a weighted average for our sample of hospitals using as weights the observed death rates within six age bands (under 15, 15-34, 35-54, 55-64, 65-74 and 75+). Male and female life expectancy is from "Interim Life Tables, England, 2006-2008", Office for National Statistics.

[57] A US dollar exchanged for a pound sterling at a rate of 0.621 on May 23, 2011 (http//:www.xe.com).

[58] We find no change in either operating expenditure or operating expenditure per admission following the implementation of the policy. So these life year gains did not increase the cost paid by the tax payer. However, the policy did lead to a fall in length-of-stay, which should have reduced costs. A conservative assumption is that the cost of achieving the extra quality is equal to the value of the reduction in length-of-stay. Using the coefficient from Table 4, column (1), and assuming a cost per day in hospital of £250 (based on personal communication with the UK Department of Health), this gives values of £22 million ($35 million) for the average change in HHI 2003-2007, and £0.36 billion ($0.58 billion) for the difference between the mean HHI and one standard deviation below the mean for the value of the reduction in length-of-stay. In both cases this is considerably less than the value of the lives saved.

in quality arising from decreases in death rates and not any other aspects of quality that are important, but not readily measured (e.g., quality of life).

It is likely that we are recovering short run effects of the policy, since we only have one year of data following implementation. If so, the long run effects could be larger once hospitals and patients adjust fully to the new system. In addition, the estimate of the gains from a change in market structure from high to low concentration (£4.8 billion) is substantially greater than the short run effect. This suggests that there could be large positive effects of policies that result in substantial decreases in concentration.[59]

VII. Summary and conclusions

We have examined the impact of the introduction of a pro-competition policy on hospital outcomes in England. Our results constitute some of the first evidence on the impact of a market-based reform in the health care sector. We find strong evidence that under the regulated price regime hospitals engaged in quality competition and that the 2006 NHS reforms were successful. Within two years of implementation, the NHS reforms resulted in significant improvements in mortality and reductions in length-of-stay without changes in total expenditure or increases in expenditure per patient. Our back of the envelope estimates suggest that the immediate net benefit of this policy is around £276 million. While this is small compared to the annual cost of the NHS of £100 billion, we have only calculated the value from decreases in death rates. Allowing for improvements in other less well measured aspects of quality will increase the benefit, as will any further falls in market concentration which may occur as the policy continues in operation. If the UK were to pursue policies that lead to de-concentration of hospital markets, the gains could be substantially larger.

These results suggest that competition can be an important mechanism for enhancing the quality of care patients receive. The adoption of pro-market policies in health care, as well as policies directed at increasing or maintaining competition such as antitrust enforcement, appears to have an important role to play in the functioning of the health sector and assuring patients' well being.

[59] This does not mean that quality was optimal following the reform. Further, this is not a precise welfare calculation, merely a simple illustrative exercise.

References

Allen, Robin, and Paul J. Gertler. 1991. "Regulation and the Provision of Quality to Heterogeneous Consumers: The Case of Prospective Pricing of Medical Services." *Journal of Regulatory Economics*, 3(4): 361-375.

Angrist, Joshua D., and Jörn-Steffen Pischke. 2008. *Mostly Harmless Econometrics: An Empiricist's Companion*. New Jersey: Princeton University Press.

Bayer, Patrick, and Robert McMillan. 2005. "Choice and Competition in Local Education Markets." NBER Working Paper 11802.

Bloom, Nicholas, Carol Propper, Stephan Seiler, and John Van Reenen. 2010. "The Impact of Competition on Management Quality: Evidence from Public Hospitals." NBER Working Paper 16032.

Bresnahan, Timothy F. 1989. "Empirical Studies of Industries with Market Power." In *Handbook of Industrial Organization*, ed. Richard Schmalensee and Robert Willig, 1011-1057. Vol. 2, Chapter 17. Amsterdam and New York: Elsevier Science/North-Holland.

Burgess, Simon, Carol Propper, and Deborah Wilson. 2005. "Will More Choice Improve Outcomes in Education and Health Care: The Evidence from Economic Research." CMPO, University of Bristol. http://www.bris.ac.uk/cmpo/choice/index.html (accessed May 04, 2010).

Capps, C., Carlton, D., and David, G. 2010. "Antitrust Treatment of Nonprofits: Should Hospitals Receive Special Care?" unpublished manuscript, Bates White LLC, University ofChicago, University of Pennsylvania.

Card, David. 1992. "Using Regional Variation in Wages to Measure the Effects of the Federal Minimum Wage." *Industrial and Labor Relations Review*, 46(1): 22-37.

Choose and Book website. 2010. Department of Health. http://www.chooseandbook.nhs.uk/patients (accessed February 27, 2010).

Cooper, Zack, Stephen Gibbons, Simon Jones, and Alistair McGuire. 2010. "Does Hospital Competition Save Lives? Evidence from the English NHS Patient Choice Reforms." LSE Health Working Paper 16/2010.

Cuellar, Alison E., and Paul J. Gertler. 2005. "How the Expansion of Hospital Systems Has Affected Consumers." *Health Affairs*, 24(1): 213-219.

Cutler, David, and Mark McClellan. 2001. "Is Technological Change in Medicine Worth It?" *Health Affairs*, 20(5): 11-29.

Cutler, David M., Huckman, Robert S., and Kolstad, Jonathan T. 2010. "Input Constraints and the Efficiency of Entry: Lessons from Cardiac Surgery." *American Economic Journal: Economic Policy*, 2(1):51-76.

Dafny, Leemore. 2005. "How Do Hospitals Respond to Price Changes?" *American Economic Review*, 95(5): 1525-1541.

Department of Health. 2002a. "Reforming NHS Financial Flows: Introducing Payment by Results." http://www.dh.gov.uk/prod_consum_dh/groups/dh_digitalassets/@dh/@en/documents/digitalasset/dh_4060476.pdf (accessed February 27, 2010).

Department of Health. 2002b. "Reforming NHS Financial Flows: Payment by Results." http://www.dh.gov.uk/prod_consum_dh/groups/dh_digitalassets/@dh/@en/documents/digitalasset/dh_4018704.pdf (accessed February 27, 2010).

Department of Health. 2003. “Response to Reforming NHS Financial Flows.” http://www.dh.gov.uk/en/Consultations/Responsestoconsultations/DH_4017035 (accessed February 27, 2010).
Department of Health. 2004. “Choose and Book.” http://www.dh.gov.uk/prod_consum_dh/groups/dh_digitalassets/@dh/@en/documents/digitalasset/dh_4088352.pdf (accessed February 27, 2010).
Department of Health. 2007. “Options for the Future of Payment by Results 2008/9-2010/11.” http://www.dh.gov.uk/prod_consum_dh/groups/dh_digitalassets/documents/digitalasset/dh_073106.pdf (accessed February 28, 2010).
Department of Health. 2009. “Report on the National Patient Choice Survey - March 2009, England.” http://www.dh.gov.uk/prod_consum_dh/groups/dh_digitalassets/documents/digitalasset/dh_103681.pdf (accessed April 27, 2010).
Dixon, Anna, Ruth Robertson, John Appleby, Peter Burge, Nancy Devlin, and Helen Magee. 2010. “How Patients Choose and How Providers Respond.” http://www.kingsfund.org.uk/publications/patient_choice.html (accessed June 04, 2010).
Dixon, Jennifer. 2004. “Editorial: Payment by Results - New Financial Flows in the NHS.” *British Medical Journal*, 328(7446): 969-970.
Douglas, George W., and James C. Miller. 1974. “Quality Competition, Industry Equilibrium, and Efficiency in the Price-Constrained Airline Market.” *American Economic Review*, 64(4): 657-669.
Ellis, Randall, and Thomas McGuire. 1986. “Provider Behavior under Prospective Reimbursement: Cost Sharing and Supply.” *Journal of Health Economics*, 5(2): 129-151.
Epple, Dennis, and Richard E. Romano. 1998. “Competition between Private and Public Schools, Vouchers, and Peer-Group Effects.” *American Economic Review*, 88(1): 33-62.
Epple, Dennis, David Figlio, and Richard Romano. 2004. “Competition between Private and Public Schools: Testing Stratification and Pricing Predictions.” *Journal of Public Economics*, 88(7-8): 1215-45.
Farrar, Shelley, Jon Sussex, Deokhee Yi, Matt Sutton, Martin Chalkley, Tony Scott, and Ada Ma. 2007. “National Evaluation of Payment by Results - Report to the Department of Health.” http://www.abdn.ac.uk/heru/documents/pbr_report_dec07.pdf (accessed April 27, 2010).
Federal Trade Commission, and U.S. Department of Justice. 2004. “Improving Health Care: A Dose of Competition.” http://www.ftc.gov/reports/healthcare/040723healthcarerpt.pdf (accessed April 28, 2010).
Ferry, Tristan, and Jerome Etienne. 2007. “Editorial: Community Acquired MRSA in Europe.” *British Medical Journal*, 335(7627): 947-948.
Frankena, Mark W., and Paul A. Pautler. 1984. “An Economic Analysis of Taxicab Regulation.” Federal Trade Commission - Bureau of Economics staff report. http://www.ftc.gov/be/econrpt/233832.pdf (accessed May 04, 2010).
Gaynor, Martin. 2006. “Competition and Quality in Health Care Markets.” *Foundations and Trends in Microeconomics*, 2(6): 441-508.
Gaynor, Martin and Vogt, William B. 2003. “Competition Among Hospitals.” *Rand Journal of Economics*, 34(4):764-785.

Gaynor, Martin, Harald Seider, and William B. Vogt. 2005. “The Volume-Outcome Effect, Scale Economies, and Learning-by-Doing.” *American Economic Review, Papers and Proceedings*, 95(2): 243-247.

Gowrisankaran, Gautam, and Robert J. Town. 2003. “Competition, Payers, and Hospital Quality.” *Health Services Research*, 38(6 Pt 1): 1403-1422.

Gowrisankaran, Gautam, Vivian Ho, and Robert J. Town. 2006. “Causality, Learning and Forgetting in Surgery.” http://www.u.arizona.edu/~gowrisan/pdf_papers/learn_forget_surgery.pdf (accessed May 6, 2010).

Haas-Wilson, Deborah. 2003. *Managed Care and Monopoly Power: The Antitrust Challenge*. Cambridge, MA: Harvard University Press.

Held, Philip J., and Mark V. Pauly. 1983. “Competition and Efficiency in the End Stage Renal Disease Program.” *Journal of Health Economics*, 2(2): 95-118.

Hoxby, Caroline M. 2000. “Does Competition among Public Schools Benefit Students and Taxpayers?” *American Economic Review*, 90(5): 1209-1238.

Jha, Ashish K., Zhonge Li, John Orav, and Arnold M. Epstein. 2005. “Care in U.S. Hospitals - The Hospital Quality Alliance Program.” *New England Journal of Medicine*, 353(3): 265-274.

Jost, Timothy S., Diane Dawson, and André den Exter. 2006. “The Role of Competition in Health Care: A Western European Perspective.” *Journal of Health Politics, Policy and Law*, 31(3): 687-703.

Kessler, Daniel P., and Mark B. McClellan. 2000. “Is Hospital Competition Socially Wasteful?” *Quarterly Journal of Economics*, 115(2): 577-615.

Kessler, Daniel P., and Jeffrey J. Geppert. 2005. “The Effects of Competition on Variation in the Quality and Cost of Medical Care.” *Journal of Economics and Management Strategy*, 14(3): 575-589.

Lakdawalla, Darius and Philipson, Tomas. 2006. “The Nonprofit Sector and Industry Performance.” *Journal of Public Economics*, 90:1681-1698.

Lee, Maw-Lin. 1971. “A Conspicuous Production Theory of Hospital Production.” *Southern Economic Journal*, 38(1):48-58.

Le Grand, Julian. 2007. *The Other Invisible Hand: Delivering Public Services through Choice and Competition*. New Jersey and Woodstock: Princeton University Press.

Minap. 2010. How the NHS Cares for Patients with Heart Attack: 9th public report. London: Royal College of Physicians.

Minap. 2008. How the NHS Manages Heart Attacks: 7th Public report. London: Royal College of Physicians.

Minap. 2005. How the NHS Manages Heart Attacks: 5th Public report. London: Royal College of Physicians.

Mukamel, Dana B., Jack Zwanziger, and Kenneth J. Tomaszewski. 2001. “HMO Penetration, Competition, and Risk-Adjusted Hospital Mortality.” *Health Services Research*, 36(6): 1019-1035.

Newhouse, Joseph P. 1970. “Toward a Theory of Nonprofit Institutions: An Economic Model of a Hospital.” *American Economic Review*, 60(1): 64-74.

Nolte, Ellen, and C. Martin McKee. 2008. “Measuring the Health of Nations: Updating and Earlier Analysis.” *Health Affairs*, 27(1): 58-71.

Nolte, Ellen, and C. Martin McKee. 2004. *Does Healthcare Save Lives? Avoidable Mortality Revisited*. London: Nuffield Trust.
Pauly, Mark V. and Redisch, Michael. 1973. "The Not-for-Profit Hospital as a Physicians' Cooperative." *American Economic Review*, 63(1):87-99.
Pope, Gregory C. 1989. "Hospital Nonprice Competition and Medicare Reimbursement Policy." *Journal of Health Economics*, 8(2): 147-172.
Propper, Carol, Simon Burgess, and Denise Gossage. 2008. "Competition and Quality: Evidence from the NHS Internal Market 1991-9." *The Economic Journal*, 118(525): 138-170.
Propper, Carol, Matt Sutton, Carolyn Whitnall, and Frank Windmeijer. 2010. "Incentives and Targets in Hospital Care: Evidence from a Natural Experiment." *Journal of Public Economics*, 94(3-4): 318-335.
Propper, Carol, and John Van Reenen. 2010. "Can Pay Regulation Kill? Panel Data Evidence on the Effect of Labor Markets on Hospital Performance." *Journal of Political Economy*, 118(2): 222-273.
Rosenbaum, Sara. 2006. "A Dose of Reality: Assessing the Federal Trade Commission/Department of Justice Report in an Uninsured, Underserved, and Vulnerable Population Context." *Journal of Health Politics, Policy and Law*, 31(3): 657-670.
Rothstein, Jesse. 2007. "Does Competition among Public Schools Benefit Students and Taxpayers? Comment." *American Economic Review*, 97(5): 2026-2037.
Ruhm, Christopher. 2006. "A Healthy Economy Can Break Your Heart." NBER Working Paper 12102.
Sage, William M., David A. Hyman, and Warren Greenberg. 2003. "Why Competition Law Matters to Health Care Quality." *Health Affairs*, 22(2): 31-44.
Schlesinger, Mark. 2006. "Markets as Belief Systems and Those Who Keep the Faith." *Journal of Health Politics, Policy and Law*, 31(3): 417-421.
Schmalensee, Richard. 1977. "Comparative Static Properties of Regulated Airline Oligopolies." *Bell Journal of Economics*, 8(2): 565-576.
Shen, Yu-Chu. 2003. "The Effect of Financial Pressure on the Quality of Care in Hospitals." *Journal of Health Economics*, 22(2): 243-269.
Street, Andrew, and Alan Maynard. 2007. "Activity Based Financing in England: the Need for Continual Refinement of Payment by Results." *Health Economics, Policy and Law*, 2: 419-427.
Tay, Abigail. 2003. "Assessing Competition in Hospital Care Markets: The Importance of Accounting for Quality Differentiation." *RAND Journal of Economics*, 34(4): 786-814.
Vander Weide, James H., and Julie H. Zalkind. 1981. "Deregulation and Oligopolistic Price-Quality Rivalry." *American Economic Review*, 71(1): 144-154.
Vogt, William B., and Robert J. Town. 2006. "How Has Hospital Consolidation Affected the Price and Quality of Hospital Care?" Robert Wood Johnson Foundation Research Synthesis Report 9. http://www.rwjf.org/files/research/no9researchreport.pdf (accessed April 28, 2010).
White, Lawrence J. 1972. "Quality Variation When Prices Are Regulated." *Bell Journal of Economics and Management Science*, 3(2): 425-436.
Wright, John, and Kaveh G. Shojania. 2009. "Editorial: Measuring the Quality of Hospital Care." *British Medical Journal*, 338(569): 783-784.

Table 1: Descriptive statistics

Variable	Mean	Standard deviation	Minimum	Maximum	Observations
Death rates					
30 day AMI mortality rate (on or after discharge, ages 35-74, %)	6.9	2.6	1.8	22.8	N = 251 n = 133
28 day all-cause mortality rate (in-hospital, all ages, %)	1.6	0.6	0.0	3.3	N = 324 n = 162
28 day all-cause mortality rate (in-hospital, excluding AMI all ages, %)	1.6	0.6	0.0	3.2	N = 324 n = 162
Market concentration					
Herfindahl-Hirschman index (actual patient flows)	5,543	1,410	2,674	9,050	N = 324 n = 162
Herfindahl-Hirschman index (predicted patient flows)	4,308	1,931	1,878	9,550	N = 324 n = 162
Length of stay and admissions					
Mean length-of-stay (days)	1.2	0.8	0.3	7.1	N = 324 n = 162
Total admissions	67,896	35,929	8,792	206,633	N = 324 n = 162
Total AMI admissions (all ages)	412	198	153	1,275	N = 262 n = 135
Elective admissions (number)	35,135	20,109	3,882	116,471	N = 324 n = 162
Elective admissions (% of total)	52.4	12.2	24.4	98.4	N = 324 n = 162
Finances and prices					
Operating expenditure (£1,000)	197,082	125,368	18,881	766,137	N = 319 n = 162
Total expenditure per admission (£1,000)	3.0	1.3	0.2	9.9	N = 319 n = 162
Market forces factor	1.00	0.07	0.89	1.28	N = 324 n = 162
Retained surplus (£1,000)	0.2	6.5	-40.3	56.0	N = 303 n = 162
Average male full time wage in area (£)	24,955	3,774	18,985	34,551	N = 320 n = 160
Area health, case mix and economic conditions					
Standardized mortality rate (per 100,000, normalized)	100.0	10.0	77.6	129.5	N = 324 n = 162
Charlson index (average for all admissions)	0.48	0.23	0.03	1.85	N = 324 n = 162
Urgent ambulance calls responded within eight minutes (%)	76.4	3.3	55.7	86.6	N = 306 n = 162
Controls for other policy changes					
Number of ISTCs within 30 kilometers	0.9	1.9	0.0	11.0	N = 324 n = 162
AMI patients having thrombolytic treatment within 30 minutes of arrival at hospital (%)	83.2	9.6	47.0	100.0	N = 238 n = 141
AMI patients having thrombolytic treatment within 60 minutes of calling for help (%)	58.9	19.6	8.0	95.0	N = 239 n = 136
AMI patients having primary angioplasty (PCI) (%)	7.1	22.6	0.0	100.0	N = 266 n = 144
AMI patients discharged on aspirin (%)	97.8	2.7	83.0	100.0	N = 281 n = 144
AMI patients discharged on beta blockers (%)	92.6	6.6	68.0	100.0	N = 281 n = 144

(cont.)

Table 1: Descriptive statistics *(continued)*

Variable	Mean	Standard deviation	Minimum	Maximum	Observations
AMI patients discharged on statins (%)	96.0	3.5	81.0	100.0	N = 281
					n = 144

Notes: Summary statistics refer to fiscal years 2003 and 2007. N = Total number of hospital-year observations for the whole sample; n = Total number of hospitals in the sample. The samples for the AMI mortality rates include only hospitals with at least 150 AMI admissions. Herfindahl-Hirschman indices computed using all elective services. Market forces factor used by the Department of Health to adjust hospital reimbursement tariffs. Male full time wage is the average of the median full-time gross wages for male workers (all occupations) in the local area districts within a radius of 30 kilometers from the hospital. Age-gender area standardized mortality rate (normalized) is an inverse distance weighted average rate specific to the hospital. Average Charlson index for admissions at the hospital. Share of urgent and life-threatening (category A) ambulance calls receiving an emergency response at the scene of the incident within eight minutes. We also use, as measures of case mix, 36 variables corresponding to shares of cause-specific admissions within 5 year age-gender bands.

Table 2: Descriptive Changes in Patient Care Seeking by Hospital Mortality Rate and Market Concentration

	AMI mortality rate (2003)					
	Bottom quartile			Top quartile		
	2003	2007	% change (2003-07)	2003	2007	% change (2003-07)
Number of elective admissions	27,787	31,882	14.7%	37,212	42,417	14.0%
Average distance travelled by patients	12.9	13.1	1.6%	10.9	11.1	1.8%
Share of patients bypassing nearest hospital	0.43	0.44	2.3%	0.43	0.40	-7.0%
Number of hospitals	38	38		37	37	
	HHI level (2003)					
	Bottom quartile			Top quartile		
	2003	2007	% change (2003-07)	2003	2007	% change (2003-07)
Number of elective admissions	21,757	26,924	23.8%	55,253	61,049	10.5%
Average distance travelled by patients	8.1	8.3	2.3%	15.5	15.5	0.5%
Share of patients bypassing nearest hospital	0.45	0.46	2.2%	0.47	0.47	0.0%
Number of hospitals	41	41		40	40	

Notes: Time period is years 2003 and 2007. Elective admissions only. 30 day AMI mortality rate on or after discharge (ages 35-74) measured in 2003 for hospitals with at least 150 AMI admissions. HHI for all elective services calculated using actual patient flows measured in 2003. Sample means of variables in the rows for quartiles of AMI mortality and HHI (bottom 25% hospitals and top 25% hospitals). Average distance travelled by patients who attended the hospital in kilometers.

Table 3: Difference-in-differences estimates of the impact of pre-reform market structure on outcomes

Dependent variable	(1) 30 day AMI mortality rate (on or after discharge, ages 35-74)	(2) 28 day all-cause mortality rate (in-hospital)	(3) 28 day all-cause mortality rate (in-hospital, excluding AMI all ages)
DiD coefficient	0.291**	0.099***	0.098***
	(0.115)	(0.031)	(0.031)
Case mix controls (36) (p-value)	0.000	0.000	0.000
Adjusted R-squared	0.483	0.977	0.977
Number of hospitals	133	162	162
Observations	251	324	324

Notes: Time period is 2003 and 2007. Models estimated by OLS with standard errors (in parentheses under coefficients) robust to arbitrary heteroskedasticity. HHI measured in year 2003 for all elective services, calculated using predicted patient flows. In addition to HHI in 2003 and the year 2007 dummy, controls are 36 case mix variables corresponding to shares of cause-specific admissions within 5 year age-gender bands. The estimation sample for the AMI mortality rate includes only hospitals with at least 150 AMI admissions. Due to data limitations, in column (3) AMI admissions for all ages (not only 35-74) are excluded. Dependent and independent variables (except case mix) are in logs. All models also include a constant and a full set of hospital dummies. P-values refer to two-tailed joint Wald tests of significance of the group of variables. * Significant at 10%; ** significant at 5%; *** significant at 1%.

Table 4: Difference-in-differences estimates of the impact of pre-reform market structure on length-of-stay, admissions and expenditure

	(1)	(2)	(3)	(4)	(5)	(6)
	Length-of-stay and admissions				Expenditure and spending per admission	
	Mean length-of-stay (days)	Total admissions (number)	Elective admissions (number)	Elective admissions (share of total)	Operating expenditure (£1,000)	Operating expenditure (£1,000) per admission
DiD coefficient	0.230***	-0.019	-0.012	0.006	-0.043	-0.029
	(0.057)	(0.030)	(0.036)	(0.016)	(0.068)	(0.071)
Case mix controls (36) (p-value)	0.000	0.000	0.000	0.000	0.879	0.011
Adjusted R-squared	0.871	0.976	0.966	0.942	0.898	0.715
Number of hospitals	162	162	162	162	162	162
Observations	324	324	324	324	319	319

Notes: Time period is 2003 and 2007. Models estimated by OLS with standard errors (in parentheses under coefficients) robust to arbitrary heteroskedasticity. HHI measured in year 2003 for all elective services, calculated using predicted patient flows. In addition to HHI in 2003 and the year 2007 dummy, controls are 36 case mix variables corresponding to shares of admissions within 5 year age-gender bands. Expenditure in columns (5)-(6) excludes capital expenditure. Dependent and independent variables (except case mix) are in logs. All models also include a constant and a full set of hospital dummies. P-values refer to two-tailed joint Wald tests of significance of the group of variables. * Significant at 10%; ** significant at 5%; *** significant at 1%.

Table 5: Robustness tests: difference-in-difference estimates

Robustness test	(1) 30 day AMI mortality rate (on or after discharge, ages 35-74)	(2) 28 day all-cause mortality rate (in-hospital)	(3) Mean length-of-stay (days)
1. Baseline	0.291**	0.099***	0.230***
	(0.115)	(0.031)	(0.057)
Observations	251	324	324
Placebo test and specification of the DID estimator			
2. Placebo DiD test for 1999-2003	0.029	0.025	-0.158
	(0.150)	(0.030)	(0.099)
Observations	238	310	295
3. Using time-varying HHI (2003 & 2007)	0.301**	0.102***	0.237***
	(0.117)	(0.030)	(0.057)
Observations	251	324	324
4. Using actual HHI (2003)	0.258	0.154***	0.489***
	(0.192)	(0.046)	(0.091)
Observations	251	324	324
5. Using indicator for top quartile of HHI (2003)	0.258***	0.079***	0.103
	(0.091)	(0.028)	(0.063)
Observations	251	324	324
Differences in costs, wages and the financial position of the hospital			
6. Controlling for the market forces factor	0.282**	0.104***	0.227***
	(0.116)	(0.030)	(0.056)
Observations	251	324	324
7. Controlling for average male wage in area	0.285**	0.075**	0.233***
	(0.123)	(0.031)	(0.060)
Observations	249	320	320
8. Controlling for surpluses/deficits	0.335***	0.103***	0.205***
	(0.121)	(0.034)	(0.060)
Observations	237	303	303
Patients' heterogeneity and number of admissions			
9. Controlling for the age-gender standardized	0.309**	0.074**	0.238***
mortality rate in area	(0.122)	(0.030)	(0.061)
Observations	251	324	324
10. Controlling for the Charlson index	0.314***	0.098***	0.219***
	(0.114)	(0.031)	(0.057)
Observations	251	324	324
11. Controlling for the number of cause-specific	0.279**	0.092***	0.223***
Admissions	(0.112)	(0.029)	(0.055)
Observations	251	324	324
Local area economic conditions			
12. Controlling for the share of urgent ambulance	0.374***		
calls responded within eight minutes	(0.131)		
Observations	234		

(*cont.*)

Table 5: Robustness tests: difference-in-difference estimates *(continued)*

Robustness test	(1) 30 day AMI mortality rate (on or after discharge, ages 35-74)	(2) 28 day all-cause mortality rate (in-hospital)	(3) Mean length-of-stay (days)
Possible contamination from other policy changes			
13. Controlling for the number of ISTCs within 30 kilometers of the hospital	0.333***	0.096***	0.215***
	(0.124)	(0.031)	(0.062)
Observations	251	324	324
14. Controlling for cardiac treatment measures	0.472**		
	(0.188)		
Observations	210		
Weighting and functional form			
15. All hospitals (weighted by number of admissions)	0.190*	0.100***	0.262***
	(0.102)	(0.026)	(0.064)
Observations	294	324	324
16. Using levels of the dependent variable and HHI (implied elasticity)	0.241**	0.084***	0.206***
Observations	251	324	324

Notes: Models estimated by OLS with standard errors (in parentheses under coefficients) robust to arbitrary heteroskedasticity. Time period is 2003 and 2007, except for the placebo DiD models in test 2 (1999 and 2003). HHI measured in year 2003 (except in test 3) for all elective services, calculated using predicted patient flows (actual patient flows in test 4). Controls are year 2007 dummy (or year 2003 dummy in test 2) and 36 case mix variables corresponding to shares of cause-specific admissions within 5 year age-gender bands. Due to data limitations, in row 2 the dependent variable in column (2) is the all-cause in-hospital mortality rate at any point during a hospital stay and the test uses the shorter period 2001-2003. Row 3 uses the time-varying HHI at the hospital level for years 2003 and 2007. Row 5 replaces the continuous HHI measure with an indicator for whether the hospital belongs to the 25% of hospitals with the largest HHIs measured in 2003. Row 6 adds as a control the market forces factor used by the Department of Health to adjust hospital reimbursement tariffs. Row 7 adds as a control the average of the median full-time gross wages for male workers (all occupations) in the local area districts within a radius of 30 kilometers from the hospital. Row 8 adds the hospital retained surplus or deficit as a control. Row 9 adds as a control the area age-gender standardized mortality rate (an inverse distance-weighted average rate specific to the hospital) normalized with mean 100 and standard deviation 10. Row 10 adds the average Charlson index for admissions to the hospital as a control. Row 11 adds as a control the number of AMI admissions (column (1)) or total number of admissions (columns (2) and (3)). Row 12 adds the share of category A ambulance calls (defined as urgent and life-threatening) receiving an emergency response at the scene of the incident within eight minutes. Row 13 adds as a control the number of ISTCs within a radius of 30 kilometers from the hospital. Row 14 adds controls for the shares of patients (*i*) having thrombolytic treatment within 30 minutes of arrival at hospital, (*ii*) having thrombolytic treatment within 60 minutes of calling for help, (*iii*) having primary angioplasty (PCI), (*iv*) discharged on aspirin, (*v*) discharged on beta blockers, and (*vi*) discharged on statins. For the variables in (*iv*)-(*vi*) we use data for year 2004 as this is the earliest available year. In row 15 regressions are weighted by AMI admissions (column (1)) or total admissions (columns (2) and (3)), and use all hospitals regardless of the number of AMI admissions in column (1) (full sample). Row 16 displays the elasticity implied by the estimated coefficient (calculated at mean values) and its significance level. In all rows except 16 the dependent and independent variables (except age-gender controls, surplus, number of ISTCs, share of patients having primary angioplasty and the indicator for HHI top quartile) are in logs. All models also include a constant and a full set of hospital dummies. * Significant at 10%; ** significant at 5%; *** significant at 1%.

Figure 1a: AMI mortality pre- and post-reform vs. market structure in 2003

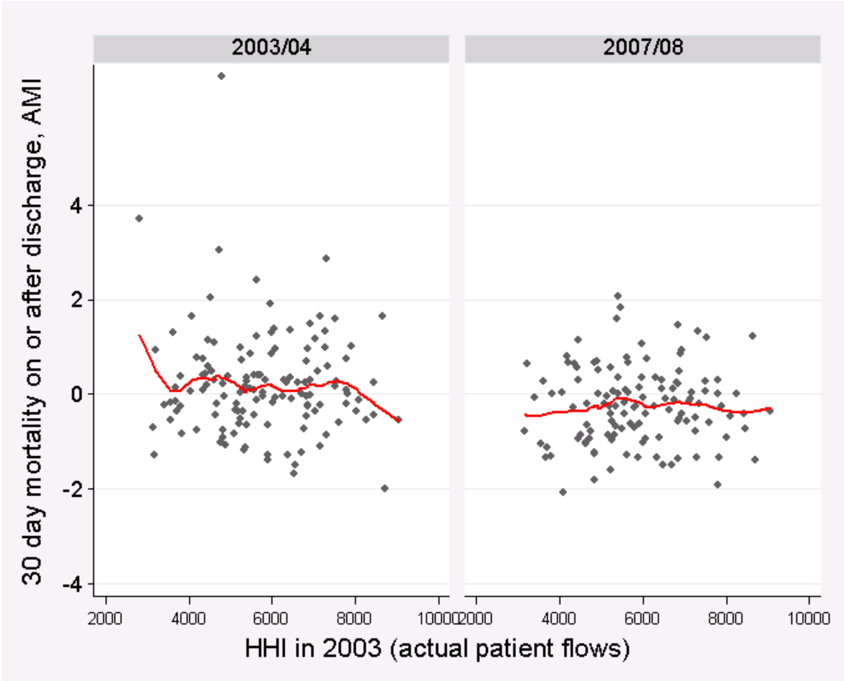
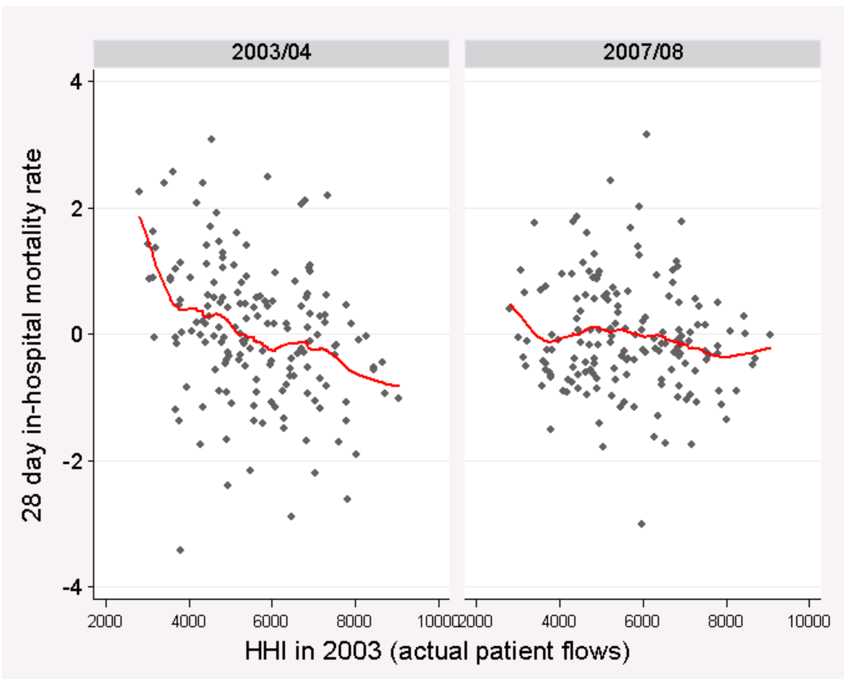


Figure 1b: All-cause mortality pre- and post-reform vs. market structure in 2003



Notes: Each point in the figure represents a hospital. HHI for year 2003 is for all elective services calculated using actual patient flows. AMI mortality rate refers to deaths in the hospital or after discharge within 30 days of emergency admission for patients aged 35-74. The line is the prediction from a locally weighted regression of the normalized mortality rate adjusted for case mix (using the shares of admissions within 5 year age-gender bands) on HHI.

Figure 2: Kernel density estimates for the distribution of HHI (all elective services)

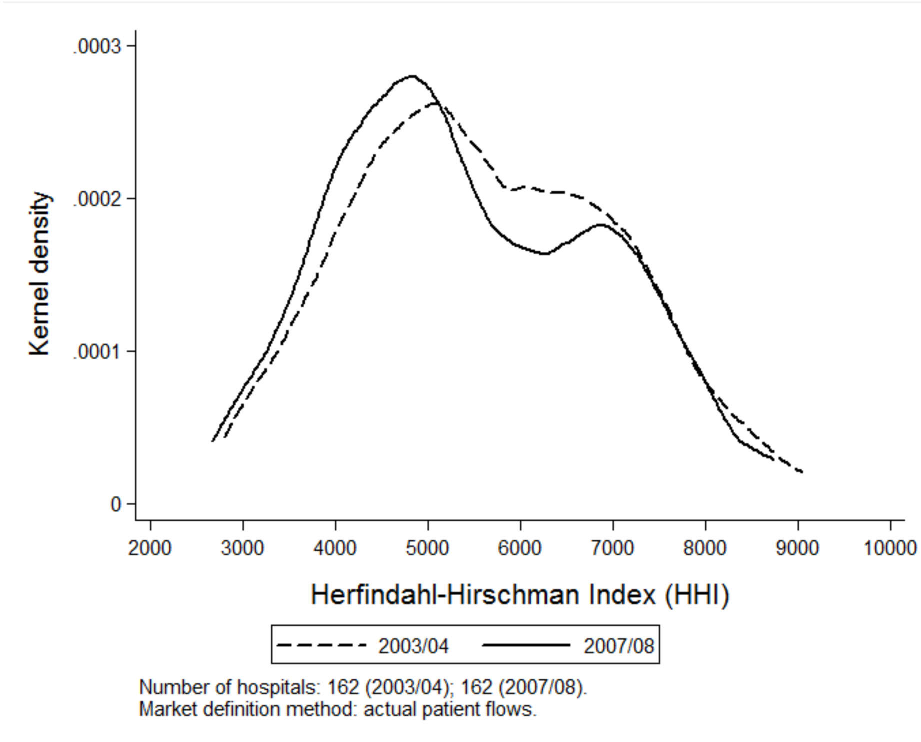
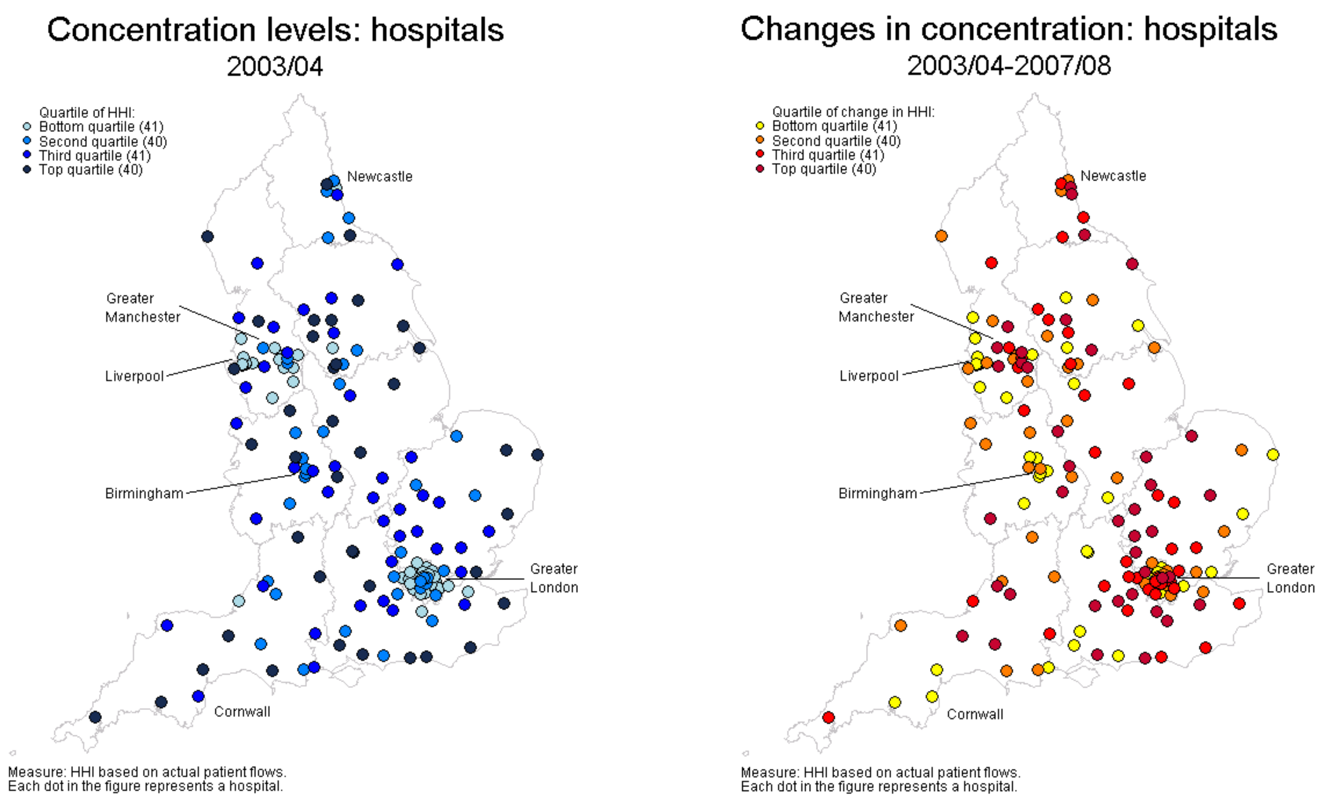


Figure 3: The location of, and changes in, market concentration (2003/04-2007/08)



Not for Publication

Appendix A

Table A1: Data sources

Variable	Source
30 day AMI mortality rate (on or after discharge, ages 35-74)	Office for National Statistics (ONS) and Hospital Episode Statistics (HES)
28 day all-cause mortality rate (in-hospital, all ages)	Hospital Episode Statistics (HES)
28 day all-cause mortality rate excluding AMI admissions (in-hospital, all ages)	Hospital Episode Statistics (HES)
Herfindahl-Hirschman index (all elective services)	Authors' own calculations using admissions data from Hospital Episode Statistics (HES)
Mean length-of-stay	Hospital Episode Statistics (HES)
Total, elective and AMI admissions	Hospital Episode Statistics (HES)
Age-gender distribution of admissions within 5 year age-gender bands	Hospital Episode Statistics (HES)
Operating expenditure	Department of Health: Trust financial returns; Published Income and Expenditure Accounts for Foundation Trusts
Market forces factor	Department of Health: Reference costs
Retained surplus or deficit	Department of Health: Trust financial returns; Published Income and Expenditure Accounts for Foundation Trusts
Full time male wages at the local authority level	Office for National Statistics: Annual Survey of Hours and Earnings (ASHE)
Age-gender standardized mortality rate at the local authority level	NHS: National Centre for Health Outcomes Development (NCHOD)
Charlson index	Authors' own calculations based on data from Hospital Episode Statistics (HES)
Urgent and life-threatening (category A) ambulance calls responded within eight minutes	Care Quality Commission (CQC)
Number of ISTCs within 30 kilometers	Authors' own calculations using Hospital Episode Statistics (HES) and Department of Health data
AMI patients having thrombolytic treatment	Myocardial Infarction Audit Project - Minap (Minap 2005, 2008, 2010)
AMI patients having primary angioplasty (PCI)	Myocardial Infarction Audit Project - Minap (Minap 2005, 2008, 2010)
AMI patients discharged on secondary prevention medication (aspirin, beta blockers, statins)	Myocardial Infarction Audit Project - Minap (Minap 2005, 2008, 2010)

Table A2: Sample selection

Year	(1) Active acute hospitals	(2) Hospitals with at least 5,000 total admissions	(3) Hospitals with non-missing HHI and mortality data	(4) Hospitals with at least 150 AMI admissions per year
2003	180	170	162	130
2007	175	167	162	121

Notes: The table reports the number of hospitals in our sample in each year (2003 and 2007) under different restrictions on the set of all English NHS hospitals. Each column puts a further restriction on the sample compared to the column before it, so column (4) is a strict sub-sample of (3) and so on. Column (1) presents the total number of active acute hospitals in each year. Column (2) refers to the number of hospitals with at least 5,000 admissions in each year. Column (3) reports the number of hospitals for which the Herfindahl-Hirschman (HHI) concentration indices could be calculated, and for which mortality data (all causes) were available. Thus column (3) corresponds to the full sample used in our main difference-in-differences model estimations. Column (4) presents the sub-sample used in the AMI mortality rate regression models.

Table A3: Descriptive statistics

Variable		Mean	Standard deviation	Minimum	Maximum	Observations
Death rates						
30 day AMI mortality rate	overall	6.9	2.6	1.8	22.8	N = 251
(on or after discharge, ages 35-74, %)	between		2.4	3.3	22.8	n = 133
	within		1.6	0.8	13.1	
28 day all-cause mortality rate	overall	1.6	0.6	0.0	3.3	N = 324
(in-hospital, all ages, %)	between		0.5	0.1	2.7	n = 162
	within		0.2	1.0	2.2	
28 day all-cause mortality rate	overall	1.6	0.6	0.0	3.2	N = 324
(in-hospital, excluding AMI all ages, %)	between		0.5	0.1	2.6	n = 162
	within		0.2	1.0	2.2	
Market concentration						
Herfindahl-Hirschman index	overall	5,543	1,410	2,674	9,050	N = 324
(actual patient flows)	between		1,395	2,742	8,896	n = 162
	within		221	4,663	6,423	
Herfindahl-Hirschman index	overall	4,308	1,931	1,878	9,550	N = 324
(predicted patient flows)	between		1,929	1,966	9,548	n = 162
	within		139	3,218	5,398	
Length of stay and admissions						
Mean length-of-stay (days)	overall	1.2	0.8	0.3	7.1	N = 324
	between		0.8	0.5	6.8	n = 162
	within		0.3	0.1	2.2	
Total admissions	overall	67,896	35,929	8,792	206,633	N = 324
	between		35,331	9,079	201,744	n = 162
	within		6,817	25,471	110,321	
Total AMI admissions (all ages)	overall	412	198	153	1,275	N = 262
	between		182	154	1,142	n = 135
	within		79	180	643	
Elective admissions (number)	overall	35,135	20,109	3,882	116,471	N = 324
	between		19,715	4,024	111,307	n = 162
	within		4,111	18,350	51,921	
Elective admissions (% of total)	overall	52.4	12.2	24.4	98.4	N = 324
	between		12.0	26.7	98.2	n = 162
	within		2.3	42.6	62.2	
Finances and prices						
Operating expenditure (£1,000)	overall	197,082	125,368	18,881	766,137	N = 319
	between		120,012	37,764	691,830	n = 162
	within		35,841	85,736	308,428	
Total expenditure per admission (£1,000)	overall	3.0	1.3	0.2	9.9	N = 319
	between		1.2	1.1	9.4	n = 162
	within		0.4	1.4	4.7	
Market forces factor	overall	1.00	0.07	0.89	1.28	N = 324
	between		0.07	0.91	1.23	n = 162
	within		0.01	0.96	1.05	
Retained surplus (£1,000)	overall	0.2	6.5	-40.3	56.0	N = 303
	between		4.7	-22.2	28.0	n = 162
	within		4.3	-27.7	28.2	
Average male full time wage in area (£)	overall	24,955	3,774	18,985	34,551	N = 320
	between		3,391	19,691	32,362	n = 160
	within		1,668	22,111	27,799	
Area health, case mix and economic conditions						
Standardized mortality rate	overall	100.0	10.0	77.6	129.5	N = 324
(per 100,000, normalized)	between		8.4	83.5	123.0	n = 162
	within		5.4	91.5	108.5	
Charlson index (average for all admissions)	overall	0.48	0.23	0.03	1.85	N = 324
	between		0.22	0.04	1.83	n = 162
	within		0.05	0.28	0.69	

(cont.)

Table A3: Descriptive statistics *(continued)*

Variable		Mean	Standard deviation	Minimum	Maximum	Observations
Area health, case mix and economic conditions						
Urgent ambulance calls responded within eight minutes (%)	overall	76.4	3.3	55.7	86.6	N = 306
	between		2.6	63.9	83.8	n = 162
	within		2.0	68.1	84.6	
Controls for other policy changes						
Number of ISTCs within 30 kilometers	overall	0.9	1.9	0.0	11.0	N = 324
	between		1.2	0.0	5.5	n = 162
	within		1.5	0.0	6.4	
AMI patients having thrombolytic treatment within 30 minutes of arrival at hospital (%)	overall	83.2	9.6	47.0	100.0	N = 238
	between		8.7	57.0	100.0	n = 141
	within		5.2	59.2	100.0	
AMI patients having thrombolytic treatment within 60 minutes of calling for help (%)	overall	58.9	19.6	8.0	95.0	N = 239
	between		14.8	8.0	90.0	n = 136
	within		13.9	21.4	96.4	
AMI patients having primary angioplasty (PCI) (%)	overall	7.1	22.6	0.0	100.0	N = 266
	between		16.3	0.0	100.0	n = 144
	within		16.2	42.9	57.1	
AMI patients discharged on aspirin (%)	overall	97.8	2.7	83.0	100.0	N = 281
	between		2.1	91.0	100.0	n = 144
	within		1.8	89.3	100.0	
AMI patients discharged on beta blockers (%)	overall	92.6	6.6	68.0	100.0	N = 281
	between		5.4	77.0	100.0	n = 144
	within		3.8	77.6	100.0	
AMI patients discharged on statins (%)	overall	96.0	3.5	81.0	100.0	N = 281
	between		2.8	85.5	100.0	n = 144
	within		2.0	89.5	100.0	

Notes: Summary statistics refer to fiscal years 2003 and 2007. N = Total number of hospital-year observations for the whole sample; n = Total number of hospitals in the sample. The samples for the AMI mortality rates include only hospitals with at least 150 AMI admissions. Herfindahl-Hirschman indices computed using all elective services. Market forces factor used by the Department of Health to adjust hospital reimbursement tariffs. Male full time wage is the average of the median full-time gross wages for male workers (all occupations) in the local area districts within a radius of 30 kilometers from the hospital. Age-gender area standardized mortality rate (normalized) is an inverse distance weighted average rate specific to the hospital. Average Charlson index for admissions at the hospital. Share of urgent and life-threatening (category A) ambulance calls receiving an emergency response at the scene of the incident within eight minutes. We also use, as measures of case mix, 36 variables corresponding to shares of cause-specific admissions within 5 year age-gender bands.

Table A4: Full difference-in-differences estimates of the impact of market structure on the 30 day AMI mortality rate (on or after discharge, ages 35-74)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
DiD coefficient	0.291**	0.029	0.301**	0.258	0.258***	0.282**	0.285**	0.335***	0.309**	0.314***	0.279**	0.374***	0.333***	0.472**	0.190*	0.00035**
	(0.115)	(0.150)	(0.117)	(0.192)	(0.091)	(0.116)	(0.123)	(0.121)	(0.122)	(0.114)	(0.112)	(0.131)	(0.124)	(0.188)	(0.102)	(0.00014)
Year 2007	-2.698***		-2.787***	-2.507	-0.346***	-2.628***	-2.613**	-3.013***	-2.998***	-2.997***	-2.629***	-3.353***	-3.079***	-4.251***	-1.943**	-3.380***
	(0.966)		(0.986)	(1.656)	(0.100)	(0.984)	(1.055)	(1.022)	(1.123)	(0.963)	(0.938)	(1.090)	(1.050)	(1.562)	(0.855)	(0.860)
Year 2003		-0.717														
		(1.290)														
HHI			-0.622													
			(0.773)													
Market forces factor						1.672										
						(2.376)										
Area male wages							-0.311									
							(1.100)									
Surplus/deficit								-0.001								
								(0.007)								
Area standardized mortality rate									-1.459							
									(2.156)							
Charlson index										0.481**						
										(0.226)						
AMI admissions											-0.182					
											(0.125)					
Urgent ambulance calls responded												-1.510				
												(0.941)				
ISTCs within 30km													0.021			
													(0.018)			
Case mix controls (36) (p-value)	0.000	0.000	0.000	0.004	0.000	0.000	0.000	0.002	0.008	0.000	0.009	0.000	0.000	0.065	0.000	0.000
Cardiac treatment controls (p-value)														0.377		
Implied elasticity																0.241
Adjusted R-squared	0.483	0.265	0.479	0.451	0.489	0.481	0.476	0.508	0.481	0.498	0.495	0.470	0.484	0.486	0.371	0.568
Number of hospitals	133	133	133	133	133	133	132	133	133	133	133	132	133	126	148	133
Observations	251	238	251	251	251	251	249	237	251	251	251	234	251	210	294	251

Notes: Full estimation results for the baseline and other models presented in the robustness checks table (see notes to Table 5 for details on the variables used). Models estimated by OLS with standard errors (in parentheses under coefficients) robust to arbitrary heteroskedasticity. Row 1 presents the baseline results. Row 2 presents the placebo DiD test. Row 3 uses the time-varying HHI at the hospital level (years 2003 and 2007). Row 4 uses the HHI calculated based on actual patient flows. Row 5 replaces the continuous HHI measure with an indicator for whether the hospital belongs to the 25% of hospitals with the largest predicted HHIs measured in 2003. Rows 6-14 add the following controls (respectively): (6) market forces factor; (7) average male wages in the districts within 30 km from the hospital; (8) retained surplus or deficit; (9) area standardized mortality rate; (10) average Charlson index for admissions to the hospital; (11) number of AMI admissions; (12) share of urgent ambulance calls receiving an emergency response within eight minutes; (13) number of ISTCs within 30 km from the hospital; (14) six cardiac treatment controls (p-value for joint Wald test of significance). In row 15 regressions are weighted by AMI admissions using all hospitals regardless of their number of AMI admissions. In row 16 the model is estimated in levels instead of logs and the elasticity implied by the estimated coefficient (calculated at mean values) is displayed. The table also displays the p-values for the joint Wald test of significance of the 36 case mix variables corresponding to shares of AMI admissions within 5 year age-gender bands. All models include a constant and a full set of hospital dummies. * Significant at 10%; ** significant at 5%; *** significant at 1%.

Table A5: Full difference-in-differences estimates of the impact of market structure on the 28 day all-cause mortality rate (in-hospital)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
DiD coefficient	0.099***	0.025	0.102***	0.154***	0.079***	0.104***	0.075**	0.103***	0.074**	0.098***	0.092***	0.096***	0.100***	0.00003***
	(0.031)	(0.030)	(0.030)	(0.046)	(0.028)	(0.030)	(0.031)	(0.034)	(0.030)	(0.031)	(0.029)	(0.031)	(0.026)	(0.00001)
Year 2007	-1.033***		-1.077***	-1.556***	-0.260***	-1.075***	-0.831***	-1.041***	-0.684***	-1.029***	-0.894***	-1.010***	-1.022***	-0.367***
	(0.238)		(0.238)	(0.387)	(0.060)	(0.240)	(0.262)	(0.269)	(0.246)	(0.242)	(0.225)	(0.243)	(0.205)	(0.079)
Year 2003		-0.228												
		(0.241)												
HHI			-0.378											
			(0.249)											
Market forces factor						-0.471								
						(0.639)								
Area male wages							-0.163							
							(0.436)							
Surplus/deficit								-0.002						
								(0.001)						
Area standardized mortality rate									1.342**					
									(0.521)					
Charlson index										-0.007				
										(0.092)				
Total admissions											-0.394***			
											(0.113)			
ISTCs within 30km												-0.002		
												(0.007)		
Case mix controls (36) (p-value)	0.000	0.003	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Implied elasticity														0.084
Adjusted R-squared	0.977	0.974	0.977	0.977	0.977	0.977	0.981	0.976	0.978	0.977	0.979	0.977	0.964	0.942
Number of hospitals	162	162	162	162	162	162	160	162	162	162	162	162	162	162
Observations	324	310	324	324	324	324	320	303	324	324	324	324	324	324

Notes: Full estimation results for the baseline and other models presented in the robustness checks table (see notes to Table 5 for details on the variables used). Models estimated by OLS with standard errors (in parentheses under coefficients) robust to arbitrary heteroskedasticity. Row 1 presents the baseline results. Row 2 presents the placebo DiD test. Row 3 uses the time-varying HHI at the hospital level (years 2003 and 2007). Row 4 uses the HHI calculated based on actual patient flows. Row 5 replaces the continuous HHI measure with an indicator for whether the hospital belongs to the 25% of hospitals with the largest predicted HHIs measured in 2003. Rows 6-12 add the following controls (respectively): (6) market forces factor; (7) average male wages in the districts within 30 km from the hospital; (8) retained surplus or deficit; (9) area standardized mortality rate; (10) average Charlson index for admissions to the hospital; (11) total number of admissions; (12) number of ISTCs within 30 km from the hospital. In row 13 regressions are weighted by total admissions. In row 14 the model is estimated in levels instead of logs and the elasticity implied by the estimated coefficient (calculated at mean values) is displayed. The table also displays the p-values for the joint Wald test of significance of the 36 case mix variables corresponding to shares of admissions within 5 year age-gender bands. All models include a constant and a full set of hospital dummies. * Significant at 10%; ** significant at 5%; *** significant at 1%.

Table A6: Full difference-in-differences estimates of the impact of market structure on mean length-of-stay (days)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
DiD coefficient	0.230***	-0.158	0.237***	0.489***	0.103	0.227***	0.233***	0.205***	0.238***	0.219***	0.223***	0.215***	0.262***	0.00006***
	(0.057)	(0.099)	(0.057)	(0.091)	(0.063)	(0.056)	(0.060)	(0.060)	(0.061)	(0.057)	(0.055)	(0.062)	(0.064)	(0.00001)
Year 2007	-2.170***		-2.268***	-4.505***	-0.357***	-2.139***	-2.153***	-1.965***	-2.273***	-2.065***	-2.028***	-2.025***	-2.438***	-0.578***
	(0.462)		(0.465)	(0.781)	(0.104)	(0.459)	(0.511)	(0.494)	(0.546)	(0.468)	(0.445)	(0.513)	(0.520)	(0.122)
Year 2003		1.300												
		(0.877)												
HHI			-0.853**											
			(0.391)											
Market forces factor						0.343								
						(0.878)								
Area male wages							-0.296							
							(0.726)							
Surplus/deficit								-0.002						
								(0.003)						
Area standardized									-0.394					
mortality rate									(0.972)					
Charlson index										-0.159				
										(0.148)				
Total admissions											-0.403**			
											(0.174)			
ISTCs within 30km												-0.012		
												(0.014)		
Case mix controls (36) (p-value)	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Implied elasticity														0.206
Adjusted R-squared	0.871	0.851	0.872	0.882	0.859	0.870	0.869	0.864	0.870	0.872	0.876	0.872	0.853	0.937
Number of hospitals	162	162	162	162	162	162	160	162	162	162	162	162	162	162
Observations	324	295	324	324	324	324	320	303	324	324	324	324	324	324

Notes: Full estimation results for the baseline and other models presented in the robustness checks table (see notes to Table 5 for details on the variables used). Models estimated by OLS with standard errors (in parentheses under coefficients) robust to arbitrary heteroskedasticity. Row 1 presents the baseline results. Row 2 presents the placebo DiD test. Row 3 uses the time-varying HHI at the hospital level (years 2003 and 2007). Row 4 uses the HHI calculated based on actual patient flows. Row 5 replaces the continuous HHI measure with an indicator for whether the hospital belongs to the 25% of hospitals with the largest predicted HHIs measured in 2003. Rows 6-12 add the following controls (respectively): (6) market forces factor; (7) average male wages in the districts within 30 km from the hospital; (8) retained surplus or deficit; (9) area standardized mortality rate; (10) average Charlson index for admissions to the hospital; (11) total number of admissions; (12) number of ISTCs within 30 km from the hospital. In row 13 regressions are weighted by total admissions. In row 14 the model is estimated in levels instead of logs and the elasticity implied by the estimated coefficient (calculated at mean values) is displayed. The table also displays the p-values for the joint Wald test of significance of the 36 case mix variables corresponding to shares of admissions within 5 year age-gender bands. All models include a constant and a full set of hospital dummies. * Significant at 10%; ** significant at 5%; *** significant at 1%.

Not for Publication

Appendix B: Estimation of predicted HHIs

Assigning hospital market competitiveness based on which hospital patients *actually* attended - rather than, for example, their area of residence - can induce a correlation between competitiveness and unobservable determinants of outcomes, because patients' hospital of admission may depend on unobserved determinants of their hospital's quality and their own health status. We therefore follow Kessler and McClellan (2000) and Gowrisankaran and Town (2003) in assigning a level of market competition to a hospital based on predicted patient flows from neighborhoods to hospitals. Hospitals are assigned the predicted level of market competition based on the neighborhoods from which they draw their patients.

To do this, we estimate a logit model for patient choice. Having estimated these models, *predicted HHIs* at the hospital level are then computed as functions of the patient level predicted probabilities. First, neighborhood level predicted HHIs are computed as the sum of squared (predicted) shares of patients from the neighborhood attending each hospital and second, the hospital level predicted HHI is calculated as a weighted average across these neighborhood HHIs, where the weights are the predicted proportions of the hospital's patients from each neighborhood. The neighborhood is defined as an MSOA (middle layer super output area).[60]

The details are as follows.

Estimated HHIs

The probability π_{ij} that patient i chooses hospital j is given by:

$$\pi_{ij} = \Pr\left(y_{ij} = 1\right) = \frac{\exp\left(\beta_1 d_{ij}\right)}{\sum_{j=1}^{J_i} \exp\left(\beta_1 d_{ij}\right)}$$

The log-likelihood function is:

$$\log L = \sum_{i=1}^{n} \sum_{j=1}^{J} \log\left(\pi_{ij}\right)$$

The predicted HHI for patient i is the sum of their squared probabilities:

$$H\hat{H}I_i = \sum_{j=1}^{J} \hat{\pi}_{ij}^2$$

Following Kessler and McClellan (2000) we compute the predicted HHI for hospital *j* as the weighted average across neighborhood level predicted HHIs where the weights equal the predicted proportions of patients from hospital *j* that live in neighborhood k .

[60] There are approximately 7,000 MSOAs in England each containing approximately 7,200 people, so they are similar in size if not a little smaller than a US zipcode. MSOAs are constructed to have maximum within MSOA homogeneity of population characteristics.

$$\hat{HHI}_j = \sum_{k=1}^{K} \left(\frac{\hat{n}_{kj}}{\hat{n}_j} \right) \hat{HHI}_k, \qquad \hat{HHI}_k = \sum_{j=1}^{J} \left(\frac{\hat{n}_{jk}}{\hat{n}_k} \right)^2$$

$$\hat{n}_j = \sum_{i=1}^{n} \hat{\pi}_{ij}, \qquad \hat{n}_k = \sum_{i=1}^{n_k} \sum_{j=1}^{J} \hat{\pi}_{ij} = \sum_{i=1}^{n_k} 1 = n_k, \qquad \hat{n}_{kj} = \hat{n}_{jk} = \sum_{i=1}^{n_k} \hat{\pi}_{ij}$$

The predicted HHI for neighborhood k is the sum of the squared shares of patients from neighborhood k who attend each hospital j.[61]

Specification of the utility function

We estimate alternative specific conditional logit models using the following specification of the patient utility function:

$$\begin{aligned} U_{ij} = & \sum_{h=1}^{2} \left\{ \beta_1^h \left(d_{ij} - d_{ij^+}^h \right) \times z_j^h + \beta_2^h \left(d_{ij} - d_{ij^+}^h \right) \times \left(1 - z_j^h \right) \right\} \\ & + \sum_{h=1}^{2} \left\{ \beta_3^h \left(d_{ij} - d_{ij^-}^h \right) \times z_j^h + \beta_4^h \left(d_{ij} - d_{ij^-}^h \right) \times \left(1 - z_j^h \right) \right\} \\ & + \sum_{h=1}^{2} \left\{ \begin{array}{l} \beta_5^h \, female_i \times z_j^h \\ + \beta_6^h \, young_i \times z_j^h + \beta_7^h \, old_i \times z_j^h \\ + \beta_8^h \, lowseverity_i \times z_j^h + \beta_9^h \, highseverity_i \times z_j^h \end{array} \right\} + e_{ij} \end{aligned}$$

where z_j^1 is a binary indicator of whether hospital j is a teaching hospital, z_j^2 is a binary indicator of whether hospital j is a big hospital (defined as being in the top 50% of the distribution of admissions), d_{ij} is the distance from the geographic centre of the neighborhood (the MSOA) for patient i to the geographic centre of the neighborhood (the MSOA) for hospital j, $d_{ij} - d_{ij^+}^h$ is the additional distance from patient i to the alternative under examination j over and above the distance to the nearest alternative j^+ which is a good substitute in terms of hospital characteristic h, $female_i$ indicates gender, $young_i$ and old_i are binary indicators of whether patient i is below 60 years old or above 75 years old respectively, and $lowseverity_i$ and $highseverity_i$ are binary indicators

[61] The predicted HHI for hospital *j* can be calculated in different ways. Gowrisankaran and Town (2003) compute the predicted HHI for hospital *j* as the weighted average across patient level predicted HHIs where the weights are equal to the predicted probability that they attend hospital *j*,

$$\hat{HHI}_j = \frac{1}{\hat{n}_j} \sum_{i=1}^{n} \hat{\pi}_{ij} \hat{HHI}_i; \qquad \hat{n}_j = \sum_{i=1}^{n} \hat{\pi}_{ij} \,.$$

When each patient lives in their own neighborhood, our approach will give the same predicted hospital level HHIs as Gowrisankaran and Town (2003). However, the larger the geographic scale of the neighborhoods, the more the HHIs based on this approach will differ from those based on the Gowrisankaran and Town (2003) approach.

of whether patient i has one ICD diagnosis or three or more ICD diagnosis respectively. All patient level variables are interacted with the variables z_j^1 and z_j^2.[62]

Following Kessler and McClellan (2000), no individual or hospital level variables are entered as main effects and as Kessler and McClellan (2000) and Gowrisankaran and Town (2003), we explicitly omit hospital level fixed effects to prevent predicted choice being based on unobserved attributes of quality. The error term, e_{ij}, is assumed i.i.d, Type I extreme value and captures the effects of unobservable attributes on patient choice.

The model is estimated for years 2003 and 2007, and undertaken separately for each of the nine Government Office Regions of England, thus allowing parameter estimates to be region-specific.[63]

Products
The sample of admissions is all elective admissions.

Sample of hospitals
We restrict our analysis to those hospitals which have 50 or more elective admissions. Hospitals with fewer admissions are dropped from the sample as are the patients who attend these hospitals.[64]

Travel distance
We restrict the distance travelled to be 100km, subject to ensuring that each patient's choice set includes the hospital actually attended and the first and second nearest hospital with each binary characteristic switched on and off.

To see why choice of both the first and second hospital is included, the following alternatives are included in all patients' choice sets, irrespective of distance: the hospital

---

[62] For example, consider the teaching hospital dimension $h=1$ and suppose that the hospital under examination is a non-teaching hospital $z_j^1 = 0$, then the differential distance $d_{ij} - d_{ij^+}^1$ is the distance to the hospital under examination over and above the distance to the nearest hospital which is also a non-teaching hospital.

[63] To make the model computation more efficient, we collapse patients who are identical in terms of model characteristics (i.e. who live in the same MSOA and go to the same hospital and have the same patient level characteristics) into groups. All patients within the group have the same choice set. Similarly, all patients within the group also have the same distances to each hospital within the choice set as distances are measured from MSOA centroids to hospital locations. Frequency weights are used in the estimation to reflect the number of patients within each group.

[64] It is possible for some alternatives within patients' choice sets to be never chosen. This is likely to happen since hospitals located outside the region under investigation will be included in the choice set of those patients living close to the boundary, even if no patients from the region under investigation go to that hospital. These faraway hospitals should not cause any problems with the statistical identification of the model parameters. This is because, unlike standard alternative-specific conditional logit models, our model does not include any hospital-specific intercepts.

actually chosen, the nearest non teaching hospital ($z^1 = 0$), the nearest teaching hospital ($z^1 = 1$), the nearest small hospital ($z^2 = 0$) and the nearest big hospital ($z^2 = 1$).

If the hospital under examination is, for example, the nearest hospital for which $z^1 = 0$, then the nearest alternative which is a good substitute will actually be the second nearest hospital where $z^1 = 0$ and so the differential distance is negative. To compute the value of this differential distance, we must also ensure that we include the second nearest hospital for which $z^1 = 0$ in patient's choice sets. The same argument can be made when the hospital under examination is the nearest hospital that has each of the other hospital characteristics (i.e. $z^1 = 1$, $z^2 = 0$, $z^2 = 1$). Thus, the following alternatives must also be included in all patients' choice sets, even if they are beyond the cut-off distance: the second nearest non teaching hospital ($z^1 = 0$), the second nearest teaching hospital ($z^1 = 1$), the second nearest small hospital ($z^2 = 0$), the second nearest big hospital ($z^2 = 1$).

Where patients actually travel further than 100km, we extend their choice set to additionally include the actual hospital attended. Each patient will thus always have at least four to nine alternatives within their choice set.

Model fit
The proportion of correct predictions is around 75%.[65] The results are robust to a range of model specifications including: (1) whether we allow model parameters to be region-specific; (2) the extent to which we expand patients' choice sets beyond the minimum set of hospitals required to estimate the model; and (3) whether we enter distance variables as linear or non-linear variables. Hospital HHIs based on predicted data are lower in value than HHIs based on actual data. The most important coefficient estimates are for distance, so that if patients were allocated to hospitals solely on a distance basis then hospitals would appear more competitive than they actually are. Actual choice of hospital is therefore based on additional factors that we have excluded from the model and these additional factors lead hospitals to become less competitive than they would otherwise be given geographical location.

[65] Parameter estimates available from the authors.