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Do Waiting Times Affect Health Outcomes?

Evidence from Coronary Bypass

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Abstract

Long waiting times for non-emergency services are a feature of several publicly-funded health systems. A key policy concern is that long waiting times may worsen health outcomes: when patients receive treatment, their health condition may have deteriorated and health gains reduced. This study investigates whether patients in need of coronary bypass with longer waiting times are associated with poorer health outcomes in the English National Health Service over 2000-2010. Exploiting information from the Hospital Episode Statistics (HES), we measure health outcomes with in-hospital mortality and 28-day emergency readmission following discharge. Our results, obtained combining hospital fixed effects and instrumental variable methods, find no evidence of waiting times being associated with higher in-hospital mortality and weak association between waiting times and emergency readmission following a surgery. The results inform the debate on the relative merits of different types of rationing in healthcare systems. They are to some extent supportive of waiting times as an acceptable rationing mechanism, although further research is required to explore whether long waiting times affect other aspects of individuals' life.

Keywords: England; Coronary bypass; Health outcomes; Hospitals; Waiting times.

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Introduction

Long waiting times for elective services are a prominent health policy issue in several OECD countries. They are prevalent in countries that combine public health insurance with low patient cost-sharing and constraints on capacity. They act as a non-price rationing mechanism which brings together the demand for and the supply of health services (Siciliani, Borowitz and Moran, 2013; Martin and Smith, 1999). Long waiting times may induce some patients to receive treatment in the private sector more swiftly at a positive price or to give up the treatment, therefore reducing the demand for public treatment. Similarly, if waiting times are long, individuals may prospectively buy private health insurance and opt for the private sector. On the supply side, when waiting times are high, providers may work harder if motivated by altruistic concerns or subject to performance targets (Cullis, Jones and Propper, 2000; Iversen and Siciliani, 2011).

A key concern with rationing by waiting is that waiting times may worsen health outcomes. Koopmanschap et al. (2005) provide alternative scenarios describing how waiting times may affect patients' health. For example, a patient may experience a health loss while waiting but her health might be restored if the treatment is effective. Alternatively, waiting times may affect not only patient's health, but also reduce treatment efficacy. If the patient waits too long, her health condition may have deteriorated so that treatment becomes less effective and health gains are reduced.

Analysing the effect of waiting times on health outcomes while on the list is important to understand whether patient's health deteriorates during the wait but does not inform us if patient's ability to benefit from surgery is also affected. Our analysis complements previous literature (reviewed below) which looks at the effect of waits on patients health while on the list (e.g. if the patient dies while waiting or is admitted to hospital as an emergency before planned CABG surgery) by investigating the effect of long waiting times on post-operative health outcomes. We measure health outcomes in terms of probability of (a) in-hospital mortality once admitted to the hospital for surgery, and (b) being admitted as an emergency for any cause in the 28 days following discharge from hospital after surgery.

The study contributes to the policy debate on the relative merits of different types of rationing in healthcare systems. If waiting times affect health outcomes, policymakers should consider alternative rationing mechanisms or introduce policies which further encourage effective prioritisation.

We focus on elective patients in need of a coronary bypass (CABG) in the English National Health Service (NHS). CABG is a common procedure for patients with serious heart conditions. Focusing on CABG is advantageous because (a) health outcomes can be unambiguously interpreted, as the risk of mortality and readmission is not negligible (more than 1% and about 4% respectively) and (b) CABG is nearly exclusively provided in the public sector, with the private sector performing only 2% of all heart surgeries, including CABG (Ludman, 2012). Therefore selection effects due to the private sector are likely to be negligible.

We employ a large sample of all patients receiving coronary bypass during 2000-2010. During this period, waiting times dramatically reduced from 220 to 50 days (Figure 1). Such reductions, unique to the United Kingdom, are the result of several policies that combined additional resources with stringent maximum waiting-times targets (Smith and Sutton, 2013). Below we argue that such policies generated changes in waiting times over time and across hospitals and provide a unique opportunity to assess whether long waiting times are associated with worse health outcomes. Although the reduction in waiting times through penalties may have also affected the referral criteria of patients added to the list, we control for patients' severity with a range of indicators and by employing an instrumental-variable approach.

Our analysis relies on three empirical strategies. First, for each year, we estimate patient-level linear probability models to analyse whether the probability of dying after admission (or being readmitted as an emergency) depends on waiting. Hospital fixed effects are included to control for variations in hospitals' resources and protocols which may act as confounders. A key issue for identification relates to prioritisation (Gravelle and Siciliani, 2008b): more severe patients wait less and have higher risk of in-hospital mortality. We address this issue introducing a range of controls. Nonetheless, we cannot exclude that unobserved severity remains, correlated with waiting times and health outcomes. This limitation is addressed with our other approaches.

Our second strategy exploits significant variations of waiting times over the years and across providers. We build a long panel with repeated observations at hospital level over eleven years. We test whether hospitals that experienced sharper reductions in waiting times resulted in better health outcomes by employing fixed-effects panel-data models, which control for time-invariant unobserved hospital heterogeneity. We account for time-varying unobserved factors by adopting an instrumental-variable approach.

Our third strategy involves patient-level models exploiting the whole panel. Waiting times are again potentially endogenous due to unobserved severity. We instrument patient-specific waiting times with the waiting time at hospital level for CABG and for Percutaneous Transluminal Coronary Angioplasty (PTCA), a less invasive procedure. Waiting times for PTCA should be correlated with waiting times for CABG, but not with CABG health outcomes, once we control for hospital characteristics.

Our results from panel-data models suggest no association of CABG waiting times with in-hospital mortality. Instead long CABG waiting times are associated with an increase in emergency readmission rates (although this effect has weak statistical significance). This is also generally the case when we employ patient-level regressions.

Literature

Limited evidence exists on whether waiting times affect post-surgery health outcomes of elective patients. Most studies are from the medical literature and focus on CABG. They tend to be small-scale studies with samples from selected providers. Légaré et al. (2005) and Carriér et al. (1993) find that CABG waiting times in Canada do not predict the probability of dying during hospitalization or other adverse outcomes (i.e. length of stay in intensive care units). Sari et al. (2007) compare health outcomes for CABG patients who waited less or more than 7 days and find no difference in morbidity, in-hospital mortality and adverse cardiac events. Sampalis et al. (2001) employ a sample of 266 patients in three hospitals and find no association between waits and mortality after surgery but evidence of reduced physical functioning, vitality and other indicators for long waiters (more than 97 days). Rexius et al. (2005) conclude that there is no evidence that prolonged CABG waits increase post-operative mortality in two Swedish hospitals. We follow the medical literature in measuring health outcomes as in-hospital mortality and probability of a post-surgery emergency admission. However, we employ a much larger sample, which includes the whole population of CABG patients over eleven years in England.

Sobolev and Fradet (2008) provide a review of the literature for CABG and suggest that long waits may worsen symptoms and clinical outcomes. Waits may also increase the probability of pre-operative death (while waiting) and unplanned emergency admission (Rexius et al., 2004; Sobolev et al., 2006, 2012; Sobolev and Kuramoto, 2010). The main difference of these studies with ours is the focus on the experience of patients while waiting, as opposed to their health once admitted for surgery, which is instead our focus.

There is an analogous literature that investigates the impact of waits for hip or knee replacement (Hajat et al., 2002; Fienden, 2005; Hirvonen et al., 2007a; Tuominen, 2013), suggesting that long waits are not associated with higher mortality and this is due to the low mortality risk. Some analyses find however an effect of long waits on quality of life. The systematic review by Hoogeboom et al. (2009) concludes that there is strong evidence that pain does not worsen during a six-month wait (Hirvonen, 2007b, for an earlier review). Self-reported functioning also does not deteriorate for patients awaiting a hip replacement, while the evidence is conflicting for knee replacement. While most studies have modest sample size, Nikolova et al. (2015) employ all patients undergoing four common procedures (hip and knee replacement, varicose veins and inguinal hernia) in English NHS hospitals for which Patient-Reported Outcome Measures (PROMs) are available and linked to HES. They find that long waits reduce health-related quality of life for hip and knee replacement patients. No evidence is found for varicose veins and inguinal hernia.

In the economics literature, Hamilton et al. (1996) analyse the impact of waits following hip fracture on the probability of death and further hospitalisation in Canada, finding no effect. A similar result for England is obtained by Hamilton and Bramley-Harker (1999). Hamilton, Ho and Goldman (2000) compare waiting times and outcomes in the US and Canada. Although waits are longer in Canada, they do not affect mortality rates. These studies differ from ours since they deal with waiting times following hip fracture that are very short (less than a week), as opposed to much longer waiting times for CABG surgeries (weeks or months). Our focus is on elective care, where waiting times are notoriously long.

Institutional background

Waiting times have been persistent in the British NHS since its inception. The NHS provides universal access to healthcare and is funded by taxation and free at point of use. Local purchasers receive budgets from central government. Family doctors act as "gatekeepers" to specialist care. Patients need a referral to access a specialist. Most hospital care is provided by public hospitals, which are separate from the purchasers. Hospitals are subject to regulatory control and receive a fixed price per patient treated. In 2003 private providers entered the market but treat a small proportion (2%) of NHS patients.

The 2000 NHS Plan recognised that the health system was underfunded relative to other European countries and had long waits. The intention has been to provide considerably extra funding in exchange of marked improvements in performance, in particular in relation to waiting times (Smith and Sutton, 2013). Such increase in resources is likely to be a key determinant in the observed reductions in waiting times. While maximum waiting-time guarantees have been implemented for at least 20 years, it was only from 2000 that sanctions were introduced for hospitals not satisfying such targets, which contributed to substantive reductions in waiting times (Propper et al., 2008).

In the English NHS, patients who need to see a specialist are usually referred by a General Practitioner (GP) to a hospital. Patients with symptoms of coronary artery disease (e.g. chest pain) are referred to a cardiologist. Direct access to hospital specialists is only possible through the

Accident and Emergency department for emergency patients. At the time of the referral, patients have the right to choose which hospital to go for their outpatient appointment and the consultant-led team who will be in charge of the patient in the first appointment at the hospital.

During the outpatient visit, the cardiologist assesses the health conditions and chooses the appropriate treatment. She may perform a non-surgical procedure to unblock the artery. If this fails the patient is referred for a CABG, usually performed by a cardiac surgeon, and placed on the waiting list (Gaynor et al, 2012). This is the time we start to calculate the wait. During the wait, a pre-assessment clinic appointment is arranged to prepare the patient for surgery and give her the opportunity to ask questions. The surgeon reviews patient’s medical history and does a physical check. The wait ends when the patient enters the hospital for the treatment. We are not aware of a formal urgency categorization to prioritise patients on the list, neither of corresponding recommended maximum waiting times for different groups, though that does not imply that doctors prioritise patients informally on the list.

Methods

We employ three specifications, whose advantages and drawbacks are outlined below.

Patient level analysis – cross-section

For each financial year, we estimate a patient-level specification using a linear probability model (LPM):

$$m_{ij} = \alpha + \beta_1 \log(w_{ij}) + \beta_2 S_{ij} + \mu_j + \varepsilon_{ij} \qquad (1)$$

where m_{ij} is a binary variable equal to one if patient i died (or was readmitted as an emergency following surgery) in hospital j, w_{ij} is waiting time, S_{ij} is a vector of measures of patients' severity and other controls, ε_{ij} is the idiosyncratic error term. μ_j is a hospital fixed effect, which controls for hospital unobservables (differences in resources, demand conditions, protocols and quality).

Our key interest is in estimating β_1. Failing to control for severity may lead to (downward) bias estimates because (a) mortality risk depends on patients' pre-operative severity and (b) more severe patients are prioritised and wait less. To account for this we include a large set of covariates. The identification strategy relies on residual unobserved severity being negligible.

We run Equation (1) for each year. Waiting times have changed over time and the relation between waiting and health outcomes may also have changed.

The models are estimated with Ordinary Least Squares (OLS), allowing for clustered standard errors at hospital level. The advantage of LPMs is that we can interpret the coefficients as marginal effects (Angrist and Pischke, 2008). We also estimated fixed-effects logit regressions but results were similar.

Hospital level analysis – panel data

Our second specification employs data aggregated at hospital level:

$$m_{jt} = \alpha + \beta_1 \log(w_{jt}) + \beta_2 S_{jt} + \mu_t + \mu_j + \varepsilon_{jt} \quad (2)$$

where m_{jt} is the mortality (or readmission) rate in hospital j in year t, w_{jt} (mean) waiting time, S_{jt} is a vector of variables which includes severity and other controls and ε_{jt} is the error term. μ_t

includes year dummies allowing for a time trend (e.g. due to technology). μ_j is a hospital fixed effect controlling for time-invariant unobservables.

To obtain an unbiased estimate of β_1 we exploit large variation in waiting times over time and across hospitals. We argue that such changes can be considered exogenous, since they were driven by general policy initiatives rather than specific ones on coronary bypass or aimed at specific hospitals. Additionally, by estimating a fixed effect model, we are able to control for time-invariant confounding factors, such as hospital organization, that could impact both waiting times and health outcomes.

Even if we control for unobservables at hospital level, there may be concurring events (e.g. other policy initiatives) that lead to omitted-variable bias. For example, the introduction of penalties for hospitals not respecting maximum waiting times may have contributed to the reduction in waiting times but also in the referral criteria to add patients on the list. We control for this possibility by instrumenting CABG waiting times with PTCA waiting times: waiting for PTCA should be correlated with waiting for CABG, but not with CABG mortality or emergency readmission rates (in Robustness Checks section we test for possible substitution between CABG and PTCA).

Under this approach β_1 can be interpreted as the effect of waiting times on health outcomes due to variation "within" hospitals.

Patient level analysis – panel data

We extend our analysis by running the following patient-level specification:

$$\mathrm{m}_{\mathrm{ijt}} = \alpha + \beta_1 \log(\mathrm{w}_{\mathrm{ijt}}) + \beta_2 \mathrm{S}_{\mathrm{ijt}} + \mu_{\mathrm{t}} + \mu_{\mathrm{j}} + \varepsilon_{\mathrm{ijt}} \qquad (3)$$

This model is analogous to (1) but it estimated using all years. Like (2), it exploits variations across hospitals and years. Like (1), Equation (3) could be subject to omitted-variable bias, if in each year patients with higher severity wait less and have a higher risk of mortality (or readmission). To address this form of endogeneity, we control for patients' severity using the extensive controls employed in (1) and we instrument the patient-level waiting times, w_{ijt} , with the mean wait in each hospital and in each year for both CABG and PTCA, i.e. we use $\bar{w}_{jt,CABG}$ and $\bar{w}_{jt,PTCA}$ as instruments. Since aggregated CABG waiting times are constructed using individual waiting times, it is reasonable to assume a strong positive correlation between (hospital-level) aggregated waits and the individual (patient-level) waits. Moreover, there is no reason to believe that aggregated waiting times are correlated with in-hospital mortality (or emergency readmission) once we control for patient's severity and hospital (unobserved) characteristics. A similar reasoning could be applied to (hospital-level) aggregated PTCA waiting times, which are expected to be correlated with patient's health status only through patient's waiting time for CABG surgery.

Data

The analysis uses the Hospital Episode Statistics (HES). This is an administrative data set which contains records of all inpatient admissions, outpatient appointments and Accident & Emergency attendances at English NHS hospitals. HES are secondary data collected by the Health and Social Care Information Centre. All patient level data are anonymised and no personal information is included. No ethical approval has been required to conduct the analysis since the project does not involve patients directly and patients are not identifiable from the data.

We use data for 11 financial years (April-to-March) between 2000-01 and 2010-11. Our sample includes all patients (elective inpatient admissions) who had a CABG surgery (OPCS-4 codes K40-K46). We exclude 30,486 patients for whom CABG was combined with PTCA and/or a heart valve procedure. These procedures can be substitutes or complements to CABG and contribute to mortality risk.

From this original sample of 150,805 CABG patients, we lose 92 patients without a valid hospital identifier and 7,276 patients with invalid waiting times (either missing or larger than two years). We exclude 10,085 patients treated in hospitals performing less than 20 CABG in a year. We drop 186 patients younger than 34 or older than 95 years. The final sample includes 133,166 patients.

For each patient we construct (a) a dummy equal to one if the patient dies in hospital between 0-29 days (inclusive) from admission for the first eligible procedure in the spell in the respective financial year; (b) a dummy if the patient is re-admitted as an emergency within 28 days following CABG surgery. These are our dependent variables measured at patient and hospital level.

(Inpatient) waiting time is the difference between the time the patient is added to the list, following specialist assessment, and the time the patient is admitted for surgery (Dixon and Siciliani, 2008).

Our control variables include patients' age, gender, number of diagnoses at admission, number of past emergency admissions to NHS hospitals the year preceding the surgery and the Charlson index (Charlson et al, 1987), which is a powerful predictor of 10-year mortality risk. We employ a proxy of patient's socioeconomic status built on income domain of Indexes of Multiple Deprivation (IMD) 2004, 2007 and 2010: the proportion of population living in low-income households in the Lower Super Output Area (LSOA) of patient's residence (Noble et al., 2004). The distribution of this indicator is split into five quintiles. We do not have explicit measures of prioritisation collected on a routine basis.

Table 1 provides descriptive statistics. On average, the waiting time for CABG is 107 days. Waiting times have reduced over the years (Figure 1). In 2000, a CABG patient waited more than 7 months before having a surgery, but only 2 months in 2010. The reduction was more substantive in 2000-2005 and then relatively stable.

In addition to a large overall variation, waiting times for CABG varied significantly within each hospital over the years. Such variation is important since fixed-effects models rely on this variability. A graphical representation of this variability is in Figure 2, which shows the mean waiting time distribution across hospitals over 11 years. There is large variability in waits both between (standard deviation is 46.74) and within hospitals (standard deviation is 107.33).

In-hospital mortality and emergency readmission rates have been stable at around 1.2% and 4.1% respectively. The average age is 65 years. 8% of patients have no comorbidities; more than half have at least five concurrent diagnoses. 82% are men. On average, patients had 0.27 emergency hospitalizations in the year preceding surgery.

Results

Table 2 provides the results for the patient-level analysis, as in Equation (1). All models include hospital fixed effects. Column (1) has no controls, while column (2) include all controls.

Our key results show that for all years (except for 2003) there is no association between waiting times and in-hospital mortality. For 2003, we find instead that longer waits are associated with a reduction in mortality. A 100% increase in waiting times (doubling the wait) reduces the probability of mortality by 0.23 percentage points (with an underlying risk of mortality of 1%). In 2003 waiting times were still high (100 days) and prioritisation was more likely to be pronounced.

A long wait may indicate that patient's severity is low. This counter-intuitive result may therefore be due to waiting times acting as a residual indicator of urgency, even after extensive severity controls.

When health outcomes are measured as probability of emergency readmission following surgery, we find a positive association (except for 2002, 2004 and 2010 when negative) but generally not statistically significant. The only exception is 2006 with a positive and statistically significant coefficient: an increase in waiting times by 100% increases the risk of an emergency readmission by 0.32 percentage points.

Table 2 does not report coefficients on control variables, which are available in the on-line Appendix. [INSERT LINK TO ONLINE FILE A] The sign of the coefficients are as expected. The risk of mortality and readmission generally increases with age, number of diagnoses, past utilization and Charlson index.

Tables 3a and 3b provide the results for the hospital-level analysis (using unbalanced and balanced sample). In Table 3a we find that waiting times are not associated with in-hospital mortality. The control variables are generally insignificant, which is not surprising given that patients' case-mix is unlikely to vary significantly over time for a given provider. Table 3b instead suggests a positive association between waiting times and the probability of an emergency readmission, although the coefficient is significant at 10% level only when additional controls are included. An increase in waits by 100% increases readmission rates by 0.49 percentage points. Given a baseline risk of 4.05%, doubling waiting times increases the risk of readmission from 4.05% to 4.54%. The effect appears therefore relatively modest.

Most waiting time reductions happened between 2000 and 2005 (Figure 1). The results may be biased towards zero if waits varied to a lower extent after 2005. In Tables 4a and 4b we split the sample in two sub-periods 2000-2005 and 2006-2010. We do not find an association of waiting times with mortality (Table 4). On the contrary, we find a positive and significant association of waiting times on emergency readmissions only during 2006-2010.

Another concern is that time trends in CABG mortality may differ across hospitals. To test this hypothesis, we interact hospital fixed effects with a linear time trend allowing each hospital' mortality to vary differently over time. Results, available from the authors, are similar to those in Table 4.

While fixed effect models control for time-invariant unobserved heterogeneity at hospital level, some residual endogeneity may persist if worse outcomes lead providers to change (possibly reduce) waiting times. We account for this by employing an instrumental variable (IV) approach (results in Table C of online Appendix [INSERT LINK TO ONLINE FILE B]). The first-stage regression suggests that waiting time for PTCA is a good predictor of waiting time for CABG. The second-stage results confirm that CABG waiting times are not associated with mortality and emergency readmissions.

One advantage of Equation (2) is that it allows focusing on exogenous variations in waits over time within hospitals. On the other hand, it may suffer from aggregation bias. To address this issue, we employ the specification in Equation (3) running a patient-level regression using the sample over 11 years, and instrument patient-level waiting times with hospital-level waiting times for CABG and PTCA. Table 5 provides the results when we do not account for endogeneity. We find that longer waiting times are associated with lower probability of dying after admission. This can be explained by unobserved heterogeneity (note that in the patient-level analysis most of the coefficients are negative, see Table 2). In contrast, we find that waiting times is not associated with emergency readmission. This is consistent with Table 2, where most coefficients are positive but not significant.

The IV results are reported in Table 6. The first-stage estimates show that the mean CABG waiting time ($\bar{w}_{jt}$) strongly predicts individual waiting time w_{ijt} (p-value close to zero), both when it is the only instrument and also when combined with mean PTCA waiting time (which varies by year and hospital). The F-statistic is well above the critical value, suggesting that the instruments are jointly significant. The second-stage results are consistent with those from Equation (2). The effect of waiting times on in-hospital mortality is negative and, differently from Table 5, not significant, indicating that when individual heterogeneity and unobserved severity are accounted for, waiting times do not affect the probability of death. On the contrary, the effect of waiting times on emergency readmission is still significant (at 10% level). The size of the coefficient is comparable to the one derived in Table 3b. An increase in waiting times by 100% increases readmission rates by 0.35 percentage points, which translates in an increase in the risk of readmission from 4.05% to 4.40%. Again, the effect appears modest.

Since we have two instruments, we can run the over-identification test and compute Hansen J-statistic. The high p-value demonstrates that the instruments are valid and correctly included as exclusion restrictions.

In summary, our key results are robust to alternative specifications. The result that longer waits affect weakly emergency readmission rates without producing any effect on mortality seems plausible. Mortality is a more extreme outcome than readmissions. It may be that longer waits deteriorate patients' health, but not enough to impair her survival rate.

Robustness Checks

A potential source of bias is selection into treatment due to waiting times: in response to long waits, some patients may be admitted as emergency or receive an alternative treatment, such as PTCA. To test whether the choice of elective CABG is uncorrelated with waiting times, we (a) replicate the analysis measuring the mortality rate for all CABG patients, including emergency ones; (b) test directly whether waiting times are positively associated with the proportion of patients admitted as an emergency and (c) test whether waiting times for CABG affect the proportion of patients who receive PTCA (as a ratio of total CABG and PTCA patients). Table 7 shows that the results are qualitatively similar when we use the mortality rate for all patients (elective and emergency). We find no evidence of hospitals admitting more CABG emergency patients or of substituting CABG with PTCA when waiting times are longer.

Since waiting times have a skewed distribution, we test whether our hospital-level results change when we replace the mean hospital wait with the median wait (which is shorter) and the 90^{th} percentile of each hospital wait distribution in each year. Table 8 suggests that there is no association between waiting time and in-hospital mortality. The association between waiting times and emergency readmission is more statistically significant when waiting time is measured at the 90^{th} percentile.

If hospitals differ in the volume of CABG patients, our estimates will be less precise. We already excluded hospitals performing less than 20 surgeries per year. As an additional check we re-estimated the models using the number of CABG surgeries in each year and hospital as weights.

The estimated coefficients are coherent with our baseline specification (available from the authors).

We discuss some possible limitations. We exploit variation in waiting times in 2000-2010. Several policies have been implemented during this period (e.g. more reliance on prospective payment systems, competition and enhanced choice). Our results may be biased if different hospitals responded differentially to these policies over time. For example, good hospitals may respond to competition policies by attracting more severe patients and higher demand, therefore increasing waiting times. A bias may arise only if changes in severity are unobserved. We control for a range of severity measures. Moreover, Gaynor et al (2013) find that competition had no effect on waiting time. Such bias is therefore unlikely.

A prospective payment system was introduced in 2003/2004, which links payments to hospital activity. This policy makes profitable for hospitals to up-code secondary diagnoses and label patients as more severe. However, we use information on total number of diagnoses and not on the type of diagnosis, which might instead proxy the severity of concurrent illness. As a check, we run Model (3) setting the maximum total number of diagnoses to the pre-reform level, and the results are unchanged. Our results are unlikely to be affected by the change in the payment system. Also, as mentioned, we run sensitivity checks allowing for a linear trend for each hospital, which will capture differential hospital responses to policies introduced over the period.

We treat mortality and emergency readmissions as independent outcomes. Using data on emergency admissions for hip fracture, Laudicella, Li Donni and Smith (2013) show that measures of emergency readmissions may be biased if they do not take into account that higher readmissions may depend on patients' survival. Our analysis finds that waiting times are never associated with mortality and therefore the association between waits and emergency readmissions is unlikely to be biased. We also test for selection due to mortality by running a Heckman selection model on patient-level data for readmissions conditional on patient's survival. The inverse Mills ratios are not significant and the results are similar to those without accounting for selection (available from the authors). Survivorship bias may be more relevant for emergency patients than elective ones.

Finally, if patients are dying while waiting, there may be a potential issue of selectivity through mortality on the list. Our data does not contain information on mortality while waiting. However, a review of the previous literature on CABG waiting times suggests that the mortality rate for patients waiting for the surgery is low and ranges between 0.5% to 2.6% (Koomen et al. 2001; Carrier et al. 1993) and this risk includes also deaths occurred for reasons unrelated to cardiac events.

Conclusions

Waiting times are a major health policy issue. Previous work showed that waiting times act as a rationing mechanism to bring demand for and supply in equilibrium. We still do not know whether rationing by waiting is an efficient and desirable rationing mechanism compared to others (such as co-payments and direct rationing).

A key policy concern with rationing by waiting is that prolonged waiting times may worsen health outcomes following surgery. This study contributes to the limited empirical evidence that informs this question. It shows that for coronary bypass variations in waiting times may lead to variations in health outcomes as measured by emergency readmission rates. No positive and significant association is found with in-hospital mortality.

Our results are important for policy and contribute to the understanding of the role and relative merits of different forms of rationing within publicly-funded systems. In the presence of excess demand, health care can be rationed in at least three different ways. One possibility is to introduce co-payments. Rationing by price implies that those patients who seek treatment can obtain it without significant waits, and therefore their health outcomes will not be affected. On the other hand, positive prices may deter some patients (those who are poor and sick) from seeking treatment, and therefore health outcomes for those may be reduced because of lack of adequate treatment. Rationing by waiting is less likely to deter poor patients from seeking treatment but long delays may potentially affect health outcomes. Another difference is that while higher co-payments raise additional resources for the funder, waiting times do not (Gravelle and Siciliani, 2008a) though when waiting times are short, an increase in waiting times may help to reduce cost of provision by reducing idle capacity (Siciliani, Stanciole and Jacobs, 2009). A third possibility is to implement 'direct' rationing where doctors refuse treatment to some patients based on low clinical need. This implies the existence of clear prioritisation rules and protocols, which is only to some extent observed in practice (and can be politically unsustainable). In practice, these three forms of rationing can co-exist. For example, an increase in waiting times could induce hospitals to reduce referrals for surgery by suggesting instead medical treatments.

Our results are to some extent supportive of waiting times as an acceptable rationing mechanism since waiting times do not appear to be associated with extreme measures of health outcomes such as in-hospital mortality and only to some extent with emergency readmissions. This may be interpreted as a sign that prioritisation works well within the NHS. We cannot however exclude that the quality of life of CABG patients is reduced when waiting times are long. Mortality and emergency readmission are extreme negative outcomes and therefore capture only one end of the distribution. For relatively healthier patients who do not die in hospital or are not admitted as an emergency, waiting times may still worsen health outcomes. Testing whether this is the case would require more refined outcome measures that are generally not available from large routine administrative databases (the only exception, to the best of our knowledge, is Sampalis et al., 2001, which tests whether long waits affects patients' quality of life though with a sample of 266 patients in three hospitals). This could be an interesting issue to explore in future research. It may also be interesting to test the effect of waiting times on health outcomes for conditions with different degree of urgency, including more urgent procedures (e.g. cancer patients) or less urgent ones (e.g. cataract surgery).

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Table 1: Descriptive Statistics. CABG elective patients: 2000/01-2010/11

Variables	Mean	Standard	Min	Max
Individual Level Variables (sample size =133,166)				
Patient died in hospital	0.0117	0.1073	0	1
Patient had an emergency readmissions within 28 days from discharge	0.0407	0.1975	0	1
Waiting time (days)	106.97	110.82	1	728
Waiting time (days) in 2000-05	139.16	128.57	1	728
Waiting time (days) in 2006-10	57.85	42.51	1	643
Age				
35-44 years old	0.0197	0.1390	0	1
45-54 years old	0.1190	0.3238	0	1
55-64 years old	0.3185	0.4659	0	1
65-74 years old	0.3977	0.4894	0	1
75-84 years old	0.1423	0.3494	0	1
85-94 years old	0.0028	0.0526	0	1
Male patient	0.8222	0.3823	0	1
Number of diagnoses at admission				
one diagnosis	0.0804	0.2719	0	1
two diagnoses	0.0780	0.2681	0	1
three diagnoses	0.1169	0.3212	0	1
four diagnoses	0.1422	0.3492	0	1
five diagnoses	0.1479	0.3550	0	1
six diagnoses	0.1384	0.3453	0	1
seven diagnoses	0.1256	0.3314	0	1
more than seven diagnoses	0.1708	0.3763	0	1
Number of past emergency admissions	0.2698	0.6352	0	41
Charlson Comorbidity Index	0.5768	0.8520	0	10
Income Deprivation Score (quintiles)				
least income deprived quintile	0.2224	0.4158	0	1
2nd income deprived quintile	0.1825	0.3862	0	1
3rd income deprived quintile	0.1972	0.3979	0	1
4th income deprived quintile	0.2056	0.4041	0	1
most income deprived quintile	0.1924	0.3942	0	1
Provider Level Variables (sample size = 324)				
Average waiting time (days)	97.32	62.37	7.21	328.98
Average patients' age	64.94	1.23	61.21	68.02
Average number of diagnoses per patient	5.32	1.93	1.00	11.56
Proportion of male patients	0.8227	0.0315	0.4742	0.8947
Number of past emergency admissions	0.2685	0.0689	0.0263	0.5046
Proportion of patients who died in hospital	0.0118	0.0078	0	0.0695
Proportion of patients with emergency readmissions	0.0405	0.0130	0	0.0867

Table 2: Patient Level Analysis. Cross-Section Models.
Effect of waiting times on mortality and readmissions

Year	Sample Size	Dependent variable: in-hospital mortality		Dependent Variable: emergency readmission	
		(1)	**(2)**	**(1)**	**(2)**
2000	13,646	-0.0002987	-0.0000842	0.0005680	0.0002792
		[-0.34]	[-0.10]	[0.42]	[0.21]
2001	13,837	-0.0007831	-0.0007331	0.0020397	0.0017898
		[-1.14]	[-1.13]	[1.24]	[1.12]
2002	14,104	-0.0015086	-0.0018849*	-0.0003972	-0.0010744
		[-1.53]	[-1.97]	[-0.28]	[-0.78]
2003	13,654	-0.0015245*	-0.0022725***	0.0030209*	0.0024057
		[-1.90]	[-2.81]	[2.01]	[1.51]
2004	13,593	-0.0004476	-0.0008855	-0.0008238	-0.0010015
		[-0.44]	[-0.84]	[-0.51]	[-0.61]
2005	11,614	0.0004950	-0.0002925	0.0014051	0.0008729
		[0.48]	[-0.26]	[0.70]	[0.44]
2006	11,151	0.0002042	-0.0007617	0.0038495**	0.0032247*
		[0.14]	[-0.51]	[2.32]	[1.92]
2007	11,844	0.0013930	0.0009229	0.0000530	0.0004030
		[1.37]	[0.99]	[0.02]	[-0.17]
2008	11,429	-0.0015440	-0.0022566	0.0008658	0.0005588
		[-1.14]	[-1.65]	[0.42]	[0.27]
2009	9,655	0.0000941	-0.0005145	0.0020787	0.0009075
		[0.06]	[-0.33]	[0.81]	[0.35]
2010	8,639	-0.0011555	-0.0016716	-0.0002173	-0.0008502
		[-0.68]	[-0.96]	[-0.07]	[-0.29]

Notes: * p <0.10, ** p<0.05, *** p<0.01. t-statistics are shown in brackets. Linear probability models with clustered robust standard errors. All specifications include hospital fixed effects. Specification (1) includes no control variables. Specification (2) includes patient's age at admission, gender, the number of diagnoses at admission, the number of emergency hospitalizations in the previous 365 days (for any cause) and income quintiles.

Table 3a: Hospital Level Analysis. Panel Data Models.
Effect of waiting times on mortality

	Dependent variable: in–hospital mortality rate			
	Unbalanced Sample		Balanced Sample	
	(1)	(2)	(1)	(2)
log(waiting times)	-0.0008503	-0.0004554	0.0016345	0.0021272
	[-0.42]	[-0.25]	[1.09]	[1.35]
Charlson Index		0.0063638		0.0075576
		[1.11]		[0.96]
male patients		-0.0100155		-0.0236277
		[-0.77]		[-1.02]
patients' age		0.0023973*		0.0018882
		[1.68]		[1.00]
n of diagnoses		-0.0005961		-0.0005331
		[-0.84]		[-0.68]
income deprivation		0.0144657		0.0243816
		[0.38]		[0.51]
past hospitalizations		0.0037953		0.0137693
		[0.31]		[0.91]
constant	0.0187076*	-0.1303792	0.0059890	-0.1041939
	[1.73]	[-1.35]	[0.60]	[-0.82]
hospitals FE	yes	yes	yes	yes
Observations	324		220	

Notes: * p <0.10, ** p<0.05, *** p<0.01. t-statistics in brackets. Fixed-effects model including year dummies; standard errors clustered at hospital level.

Table 3b: Hospital Level Analysis. Panel Data Models.
Effect of waiting times on readmissions

	Dependent variable: emergency readmission rate			
	Unbalanced Panel		Balanced Panel	
	(1)	(2)	(1)	(2)
log(waiting times)	0.0057596**	0.0048840*	0.0035729	0.0035660
	[2.64]	[1.88]	[1.55]	[1.52]
Charlson Index		0.0015954		0.0075049
		[0.21]		[0.93]
male patients		0.0217516		0.0065585
		[0.74]		[0.14]
patients' age		0.0012880		0.0029504**
		[0.96]		[2.57]
n of diagnoses		0.0003138		0.0007106
		[0.31]		[0.63]
income deprivation		0.2264362**		0.2777286**
		[2.61]		[2.41]
past hospitalization		-0.0041702		-0.0002132
		[-0.23]		[-0.01]
constant	0.0093396	-0.1176979	0.0204009	-0.2148450**
	[0.81]	[-1.26]	[1.69]	[-2.50]
hospitals FE	yes	yes	yes	yes
Observations	324		220	

Notes: * p <0.10, ** p<0.05, *** p<0.01. t-statistics in brackets. Fixed-effects model including year dummies; standard errors clustered at hospital level.

Table 4: Hospital Level Analysis. Panel Data Models in two sub-periods. Effect of waiting times on mortality and emergency readmission

	Dependent variable:			
	in-hospital mortality rate		emergency readmission rate	
	2000-05	2006-10	2000-05	2006-10
log(waiting times)	-0.0013290	0.0032361	0.0041493	0.0095292**
	[-0.44]	[1.06]	[1.06]	[2.37]
Charlson Index	0.0133009*	0.0205238*	-0.0012584	-0.0105682
	[1.74]	[1.99]	[-0.15]	[-0.57]
male patients	-0.0107531	-0.0147130	0.0112005	0.0406979
	[-0.31]	[-1.59]	[0.21]	[1.33]
patients' age	0.0042726*	-0.0005928	0.0014022	0.0038950**
	[1.84]	[-0.44]	[0.73]	[2.15]
n of diagnoses	-0.0003913	-0.0016946	0.0009951	0.0006015
	[-0.28]	[-1.27]	[0.81]	[0.27]
income deprivation	0.0687616	0.0381129	0.0199846	0.3594544**
	[0.52]	[0.57]	[0.11]	[2.37]
past hospitalization	0.0050440	0.0029481	-0.0106507	-0.0190434
	[0.27]	[0.24]	[-0.44]	[-0.68]
constant	-0.2554769	0.0424847	-0.0839738	-0.3302220**
	[-1.53]	[0.45]	[-0.58]	[-2.38]
hospitals FE	yes	yes	yes	yes
year dummies 2001-05	yes	no	yes	no
year dummies 2007-10	no	yes	no	yes
Observations	175	149	175	149

Notes: * p <0.10, ** p<0.05, *** p<0.01. t-statistics are shown in brackets.
Unbalanced panel. Fixed-effects model with standard errors clustered at hospital level.

Table 5: Patient Level Analysis. Whole Panel.
Effect of waiting times on mortality and readmissions

	Dependent variable: in-hospital mortality		Dependent variable: emergency readmission	
	(1)	(2)	(1)	(2)
log(waiting time)	-0.0005946*	-0.0007689***	0.0012020**	0.0007985
	[-2.02]	[-2.70]	[2.46]	[1.63]
hospitals FE	yes	yes	yes	yes
Observations	133,166	133,166	133,166	133,166

Notes: * p <0.10, ** p<0.05, *** p<0.01. t-statistics in brackets. Fixed-effects model including year dummies, with standard errors clustered at hospital level. Column (1) includes no control variables. Column (2) includes additional controls (see Table 2).

Table 6: Patient Level Analysis. Whole Panel.
FE Instrumental Variable Model

	CABG waiting times as IV		CABG and PTCA waiting times as IV	
First stage				
log(CABG wt)	1.112029***	[26.08]	1.106934***	[26.59]
log(PTCA wt)			0.0554588*	[1.74]
F-statistic	680.23	(0.0000)	360.22	(0.0000)
Second stage – dependent variable: in-hospital mortality				
log(wt)	0.0029032	[1.40]	0.0028265	[1.37]
overidentification test			0.662	(0.4160)
endogeneity test	2.736	(0.0981)	2.280	(0.1311)
Second stage – dependent variable: emergency readmission				
log(wt)	0.0036514*	[1.85]	0.0035337*	[1.81]
overidentification test			1.843	(0.1746)
endogeneity test	2.207	(0.1374)	1.604	(0.2053)
Observations	133,166		133,035	

Notes: * p <0.10, ** p<0.05, *** p<0.01. t-statistics in brackets. Fixed effects instrumental variable model including all control variables and year dummies. The F-statistic for the significance of the exclusion restriction(s) is distributed as a chi-squared with one (two, in the last specification) degree of freedom, under the null. The endogeneity test provides the Hausman statistic testing the null hypothesis that the baseline and IV estimates are the same. Under the null, the statistic is distributed as a chi squared with one degree of freedom. The over-identification test provides the Hansen J statistic for the joint validity of instruments.

Table 7: Hospital Level Analysis. Robustness Checks.

	In-hospital mortality rate for elective and emergency CABG patients	Proportion of emergency CABG patients	Proportion of PTCA patients
log(waiting times)	-0.0017926	-0.0053354	0.0204818
	[-1.00]	[-0.73]	[0.96]
hospitals FE	yes	yes	yes
Observations	324	324	318

Notes: * p <0.10, ** p<0.05, *** p<0.01. t-statistics in brackets. Unbalanced panel. Fixed-effects model including year dummies with standard errors clustered at hospital level.

Table 8: Hospital Level Analysis. Additional Robustness Checks.

	Average waiting time		Median waiting time		90^{th} percentile of waiting time distribution	
	In-hospital mortality	Emergency readmission	In-hospital mortality	Emergency readmission	In-hospital mortality	Emergency readmission
log(waiting times)	-0.0004554	0.0048840*	-0.0011880	0.0028426	0.0007398	0.0065304**
	[-0.25]	[1.88]	[-1.07]	[1.32]	[0.29]	[2.57]
hospitals FE	yes	yes	yes	yes	yes	yes
Observations	324	324	324	324	324	324

Notes: * p <0.10, ** p<0.05, *** p<0.01. t-statistics are shown in brackets. Unbalanced panel. Fixed-effects model including year dummies with standard errors clustered at hospital level.

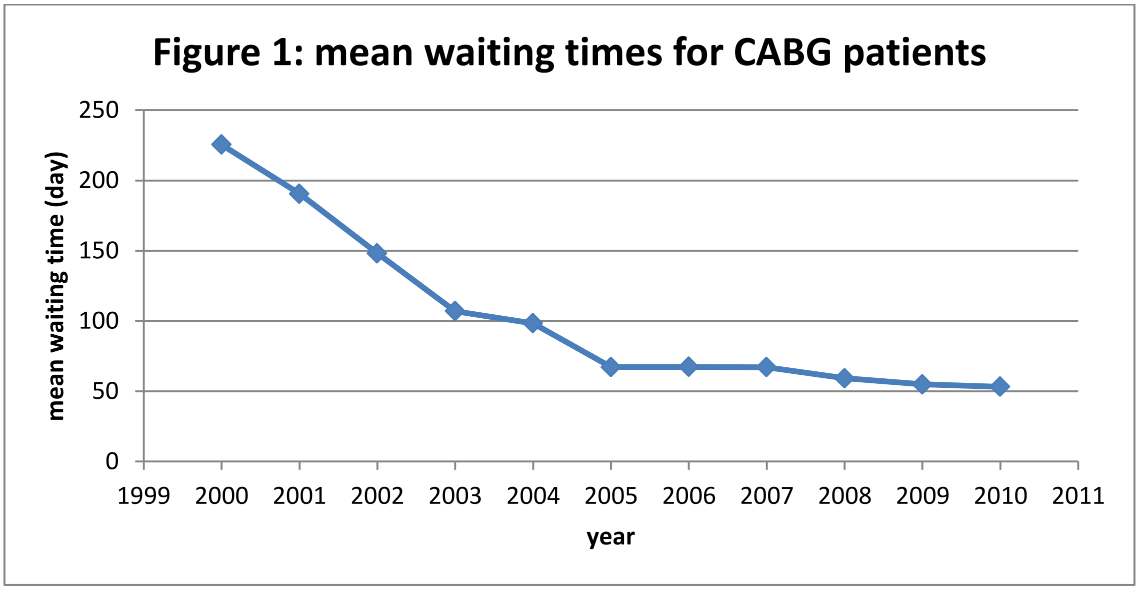
Figure 1: mean waiting times for CABG patients
mean waiting time (day)
250
200
150
100
50
0
1999
2000
2001
2002
2003
2004
2005
2006
2007
2008
2009
2010
2011
year

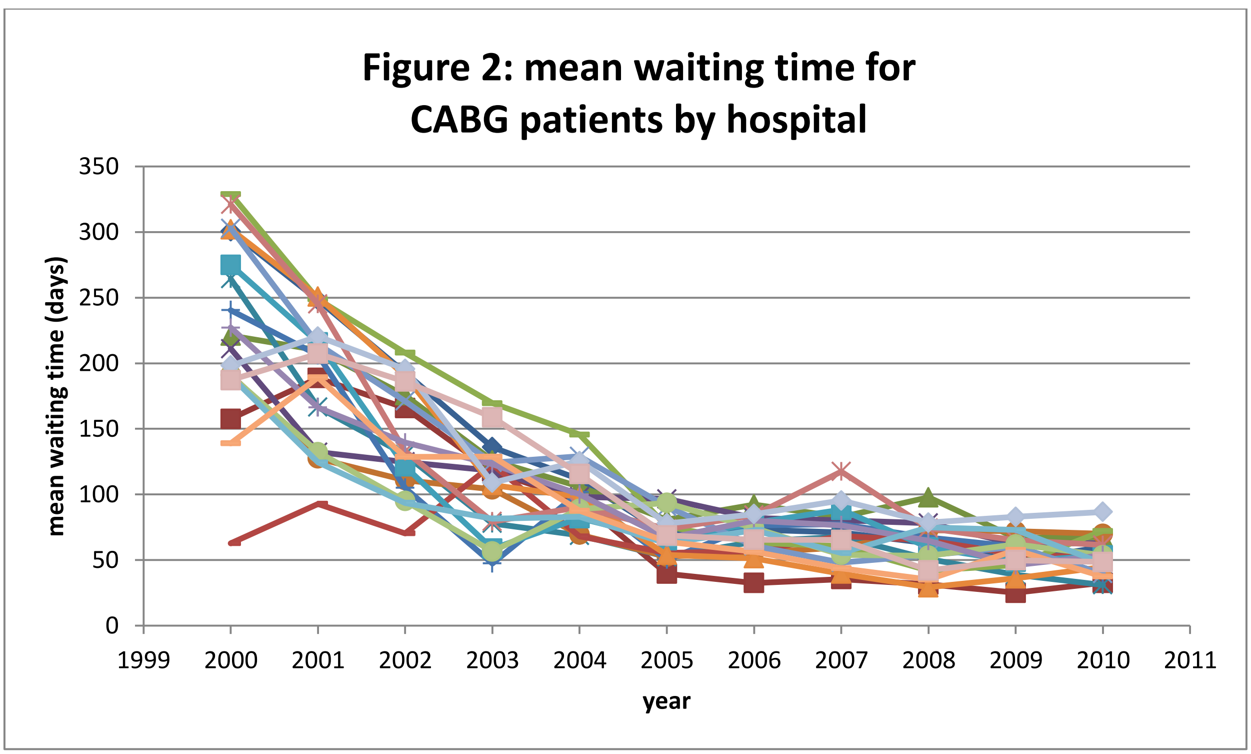
Figure 2: mean waiting time for CABG patients by hospital
mean waiting time (days)
350
300
250
200
150
100
50
0
1999
2000
2001
2002
2003
2004
2005
2006
2007
2008
2009
2010
2011
year