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RESEARCH ARTICLE

Do Reduced Hospital Mortality Rates Lead to Increased Utilization of Inpatient Emergency Care? A Population-Based Cohort Study

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Objectives. To measure the impact of the improvement in hospital survival rates on patients' subsequent utilization of unplanned (emergency) admissions.

Data Sources/Study Setting. Unplanned admissions occurring in all acute hospitals of the National Health Service in England between 2000 and 2009, including 286,027 hip fractures, 375,880 AMI, 387,761 strokes, and 9,966,246 any cause admissions.

Study Design. Population-based retrospective cohort study. Unplanned admissions experienced by patients within 28 days, 1 year, and 2 years of discharge from the index admission are modeled as a function of hospital risk-adjusted survival rates using patient-level probit and negative binomial models. Identification is also supported by an instrumental variable approach and placebo test.

Principal Findings. The improvement in hospital survival rates that occurred between 2000 and 2009 explains 37.3 percent of the total increment in unplanned admissions observed over the same period. One extra patient surviving increases the expected number of subsequent admissions occurring within 1 year from discharge by 1.9 admissions for every 100 index admissions (0.019 per admission, 95% CI, 0.016–0.022). Similar results in hip fracture (0.006[0.004–0.007]), AMI (0.006[0.04–0.007]), and stroke (0.004(0.003–0.005)).

Conclusions. The success of hospitals in improving survival from unplanned admissions can be an important contributory factor to the increase in subsequent admissions.

Key Words. Risk adjustment for resource use or payment, health care costs, hospitals, quality of care/patient safety (measurement)

Health systems in many high-income countries are struggling to cope with an increasing demand for hospital services coupled with increasing pressures to reduce costs. In the United States, the number of emergency department (ED) visits increased by more than a third between 1997 and 2007 (Tang et al. 2010) and hospital admissions from EDs have increased by 50.4 percent from 11.5 million in 1993 to 17.3 million in 2006 (Schuur and Venkatesh 2012). Hospital-based emergency care has made remarkable progress in research, training, and technical capabilities, but such achievements might be compromised by growing demand and declining financial support (Kellermann et al. 2013).

In spite of large differences in the health system, similar trends are observed in England. Accident and Emergency (A&E) departments are facing an ever-increasing demand and in recent years many have failed to achieve a four-hour waiting time target for treatment set by the regulator (Blunt, Edwards, and Merry 2015). Unplanned admissions to hospital (i.e., emergency admissions) increased by 47 percent in the past 15 years from 3.6 million in 1997 to 5.3 million in 2012 with emergency admissions now accounting for 67 percent of all hospital bed days in England (NAO 2013).

Three main factors are often cited to explain the rapid increase in emergency admissions: the aging population, the introduction of new financial incentives, and new performance targets for health care providers. Numerous studies show that demographic trends and population health are able to explain only 40–50 percent of the increment in the utilization of emergency care (Strunk, Ginsburg, and Banker 2006; Tang et al. 2010; Cowling et al. 2014; Wittenberg et al. 2014). This has been often considered sufficient to conclude that health care providers must be responsible for the remaining unexplained growth.

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In England, the switch from block contracts, where providers were paid a fixed budget irrespective of the number of patients treated, to prospective case payments, under which providers are paid per unit of care, has provided hospitals with a financial incentive to admit more patients and reduce length of stay (Farrar et al. 2009). Also, the introduction of waiting time targets in A&E has increased pressure on consultant specialists to assess patients quickly and might have resulted in more patients being admitted to hospital unnecessarily for short periods. However, no causal relationship can be found between the introduction of the new payment system and the growth in hospital activity (Farrar et al. 2009), nor between waiting time targets and the growth of short-stay emergency admissions (Nuffield Trust 2010; Weber et al. 2012). Moreover, emergency admissions were growing before the introduction of these two policies.

This study investigates an alternative potential driver of emergency admissions that to our knowledge has not been considered by existing empirical investigations. We examine the impact that hospitals' increasing success in reducing patients' mortality may have had on the subsequent utilization of emergency services over a long period. In the past decades, hospitals have achieved marked improvements in survival rates by investing in new medical technologies, adopting more effective surgical interventions and implementing new policies on patient safety (Wang et al. 2014). Overall hospital mortality rates in the United States dropped by 20 percent from 2.5 to 2.0 deaths per 100 admissions between 2000 and 2010 (Hall, Levant, and DeFrances 2013). Similar trends are observed in England and reported in this study. A growing number of patients survive after surgery in a large number of acute life-threatening conditions both in the United States and in England (Krumholz et al. 2009; Ovbiagele 2010; Finks, Osborne, and Birkmeyer 2011; Smolina et al. 2012; Lisk and Yeong 2014). However, an unintended consequence of this success may be a growing population of patients who are increasingly frail and at high risk of more hospital admissions over time (Laudicella, Li Donni, and Smith 2013).

Our study tests the hypothesis that reducing hospital mortality rates (i.e., improving hospital survival rates) increases utilization of emergency admissions over time. Our analysis also provides an estimate of the scale of such an effect in the context of the growth of emergency admissions occurring between 2000 and 2009. Our study contributes to the wider literature on the effect of medical improvements and new technologies on the growing costs of health care and population health (Cutler and McClellan 2001; Cutler, Rosen, and Vijan 2006).

METHODS

Our main hypothesis is that hospitals that are more successful in saving the lives of their patients generate a population of patients with a higher risk of using emergency care after their first admission. The main purpose of this study was to estimate the share of the total increment in emergency admissions that can be explained by the improvement in hospital survival rates. The second aim was to test if hospital performance in improving survival rates adversely affects their performance in 28-day emergency readmissions. Our data come from an extract of the Hospital Episode Statistics (HES) linked to patient mortality data from the Office for National Statistics (ONS).

Study Sample

We examined a number of separate cohorts of patients with an emergency admission to NHS acute hospital trusts for a life-threatening condition occurring between April 2000 and March 2010 (i.e., fiscal years 2000 and 2009). We label a patient's first admission as the *index admission* to facilitate illustration. We follow patients in each cohort for 2 years after discharge from their index admission and for 2 years before. Therefore, the actual data on admissions examined stretches from April 1998 to March 2012 to allow for a complete follow-up of each patient. We excluded patients with any history of emergency admissions in the 2 years before the index admission to prevent confounding effects from the care received during past admissions and interactions with unobservable patient characteristics.

We selected three separate cohorts of patients with a primary diagnosis of a hip fracture at age 65+, acute myocardial infraction (AMI) at age 55+, and stroke at age 55+. These conditions require immediate treatment, and thus, patients are normally admitted to the closest hospital rather than a hospital of their choice, so the potential confounding effects arising from patient selection of hospitals with different quality levels are minimal. We also constructed a fourth aggregate cohort of patients aged 55+ with an index emergency admission from any heath condition, that is, all-cause emergency admissions.

We examined a total of 286,027 patients in the hip fracture cohort, 375,880 in the AMI cohort, and 387,761 in the stroke cohort. We studied 9,966,246 patients with any cause emergency admission.

Dependent Variables

We counted the total number of emergency admissions experienced by patients in our cohorts within 28 days, 1 year, and 2 years from the discharge date of their index admission. Subsequent emergency admissions are defined as unplanned hospital admissions for any condition occurring in any NHS hospital in England after the index admission. They include admissions via the A&E department (65 percent of total), via urgent GP referral (25 percent), via urgent outpatient specialist referral (3 percent), and other routes (7 percent). Descriptive statistics for these variables and for each cohort are presented in Appendix SA2. Patients with a very large number of subsequent admissions (beyond the 99.95 percentile) are excluded.

Exposure and Control Variables

Hospital survival rates are defined as survival events per 100 admissions and thus range from 0 to 100. We consider all deaths within 30 days from the index admission regardless of whether they occurred in hospital or after hospital discharge. We then use a patient-level probit model to estimate risk-adjusted survival rates for each hospital in each year for each cohort. Separate survival rates and models are estimated for each of the selected cohorts.

Differences in patient health risk are controlled for by using the Charlson index, a set of dummies for the conditions included in the Charlson index, patient age and gender, and indicators for education, health and income deprivation associated with the patient's place of residence. Additional controls for cohort-specific risk factors are also included the following: fixation, prostatic replacement, or no operation for hip fracture; subsequent infarction (ICD-10 code I22) for AMI; and type of stroke (I60, I61, I62, I63, or I64) for stroke patients. The model for any cause emergency admissions also includes a set of dummies for the specialty of patients' admission.

We controlled for the accessibility of hospital services and primary care services in the patient's area of residence as these might influence the ability of different patients to access secondary care services. To this end, we included indicators for patients living in a rural or urban area, population density and the number of hospitals (within a 15 km radius) and GP practices (within a 3 km radius) of patient place of residence.

Finally, we included controls for hospital risk-adjusted length of stay, volume of admissions, and prevalence of day cases admissions to prevent confounding effects from other indicators of hospital quality that might be linked to hospital mortality and emergency admissions (Heggestad 2002; Southern and Arnsten 2015).

Statistical Analysis

For each cohort we used a patient-level negative binomial model to estimate the effect of variation in hospital risk-adjusted survival rates (i.e., the exposure variable) on patient future utilization of emergency admissions within 1 and 2 years of discharge from index admission. We also use a patient-level probit model to estimate the same effect on the probability of any emergency admission occurring within 28 days of discharge from the index admission.

The effect of hospital survival rates is estimated separately in each cohort, and we use year–level and hospital-level fixed effects, and clustered robust standard errors to allow for patients clustered within hospitals. The identification of the effect of hospital survival rates on patient subsequent utilization of emergency services is therefore obtained from the within hospital variation in survival rates over time, rather than from the comparison of different hospitals with different survival rates. In other words, the effect on emergency admissions is determined by the difference in how and when hospitals begin to improve their survival rates over time. All time-invariant hospital characteristics, such as teaching status, long-term quality reputation, and long-term integration with other care services, are controlled by the hospital-level fixed effect. The common time trend is controlled by the year-level fixed effect.

Our study sample is distributed across a total of 200–300 hospitals depending on the cohort examined (Table 1). This provides a sufficiently large number of admissions and mortality events per hospital-year for the identification of hospital survival rates.

The models include the same set of controls for differences in patient health risk described in the exposure variable section above.

The statistical analysis was conducted using *STATA* 13 (StataCorp. 2013. Stata Statistical Software: Release 13. College Station, TX: StataCorp LP).

Robustness Check, Placebo Test, and Sensitivity Analysis

An instrumental variable approach was used to test the robustness of our findings to potential endogeneity of hospital survival rates to subsequent admissions. We constructed an indicator for index admissions occurring over bank holidays and weekends and use it as an instrument to predict hospital survival

Table 1: Patients' Characteristics in the Study Sample and at the Point of the Index Admission

	Number of Patients	*Number of Hospitals*	*Share of Male Patients*	*Average Patients' Age*	*Average Number of Diagnoses**	*Average Charlson Index*	*Average Income Deprivation†*
Hip fracture	286,027	202	20.6%	82.7	4.579	0.565	0.137
Acute myocardial infarction (AMI)	375,880	214	62.4%	72.6	3.737	1.614	0.143
Stroke	387,761	234	47.3%	76.4	3.720	1.572	0.144
All-cause emergency admissions	9,966,246	303	46.6%	73.6	3.606	0.878	0.146
Cataract	660,437	148	37.4%	76.2	2.455	0.252	0.140

Notes: The study sample includes patients with an index admission for a hip fracture age 65+, AMI age 55+, stroke age 55+, and all-cause emergency admissions age 55+. An index admission is defined as the first hospital admission experienced by the patient in the past 2 years. Patients with a hospital admission in the 2 years before the index admission are not included in the study, that is, 40% in the hip fracture, 31% in the AMI, 37% in the stroke, 31% in all-cause emergency, and 24% in the cataract cohort. Data on admissions occurring in 1998 and 1999 are used to allow for calculating index admissions in 2000 and 2001.
*Includes primary and secondary diagnoses.
†The IMD 2004 income deprivation index measures the share of people relying on income benefits in the patient's area of residence.

rates. Previous studies show that unplanned admissions during bank holidays and weekends are associated with greater mortality risk due to lower availability of specialists and senior staff over these periods[1] (Aylin et al. 2010, 2013). In contrast, being admitted during the weekend should have no direct effect on total admissions 1 and 2 years after discharge, since patients surviving the weekend admission receive similar care to other patients during the rest of their hospital stay.[2]

We estimated a Two Stage Least Square model (2SLS) on the aggregated cohort of patients with an index emergency admission for all cause as they provide a large number of observations per hospital per year and can be approximated by a linear model. The 2SLS model includes the same control variables as the non-IV model, including hospital and year fixed effects. The predictions of the 2SLS are similar to the non-IV model (Table 2).

We also conducted a placebo test by estimating the effect of the four emergency hospital survival rates on the utilization of emergency care in a cohort of patients aged 65+ admitted for elective cataract surgery. We

Table 2: Effect of Improving Hospital Survival Rates on Subsequent Admissions Experienced by Patients within Twenty-Eight Days, One Year, and Two Years of Discharge from an Index Admission

	Subsequent Emergency Admissions per Patient Within					
	28 days	*95% CI*	*One year*	*95% CI*	*Two years*	*95% CI*
+1 patient surviving						
per 100 hip fracture admissions	0.001***	(0.000–0.002)	0.006***	(0.004–0.007)	0.007***	(0.005–0.010)
per 100 AMI admissions	0.001***	(0.000–0.001)	0.006***	(0.004–0.007)	0.009***	(0.007–0.011)
per 100 stroke admissions	0.001***	(0.000–0.002)	0.004***	(0.003–0.005)	0.007***	(0.005–0.008)
per 100 all-cause emergency admissions	0.002***	(0.001–0.004)	0.019***	(0.016–0.022)	0.028***	(0.024–0.033)
per 100 all-cause emergency admissions – Instrumental Variable Model	0.070***	(0.065–0.076)	0.028***	(0.011–0.045)	0.030**	(0.005–0.055)

Notes: Estimates from a patient-level Probit (28 days admissions) and Negative Binomial model (1 and 2 years admissions) with cluster robust standard errors. The model includes controls for patient-level characteristics (age, gender, primary diagnosis, and comorbidities), hospital-level characteristics (hospital fixed effects, volume of admissions, and risk-adjusted length of stay, and share of day-case activity), area-level characteristics (area deprivation, distance from closest hospital, total hospitals within 15 km, Primary Care Services within 15 km), and year fixed effect. A Two Stage Least Squares model with Instrumental Variable is used as a robustness check to the main analysis.

The study sample includes patients with an index admission for a hip fracture age 65+, AMI age 55+, stroke age 55+, and all-cause emergency admissions age 55+. An index admission is defined as the first hospital admission experienced by the patient in the past 2 years. Patients with a hospital admission in the 2 years before the index admission are not included in the study. Data on admissions occurring in 1998 and 1999 are used to allow for calculating index admissions in 2000 and 2001.

**p-value < .05; ***p-value < .01.

expect that improvements in hospital survival rates should not be associated with the utilization of emergency care in the placebo. Also, we expect these patients to be at risk of an emergency admission due to their age 65+. Hospital survival failed to achieve statistical significance in any of the estimated placebo models.

We conducted a number of other sensitivity analyzes. We re-estimated our models using unadjusted mortality rates to avoid potential risk-adjustment bias from variation in reporting patient health conditions across hospitals and over time (Mohammed et al. 2009). We also tested the sensitivity of our results to the control variables included in the models. Our estimates are very robust to these sensitivity analyzes, and full results are included in Appendix SA2.

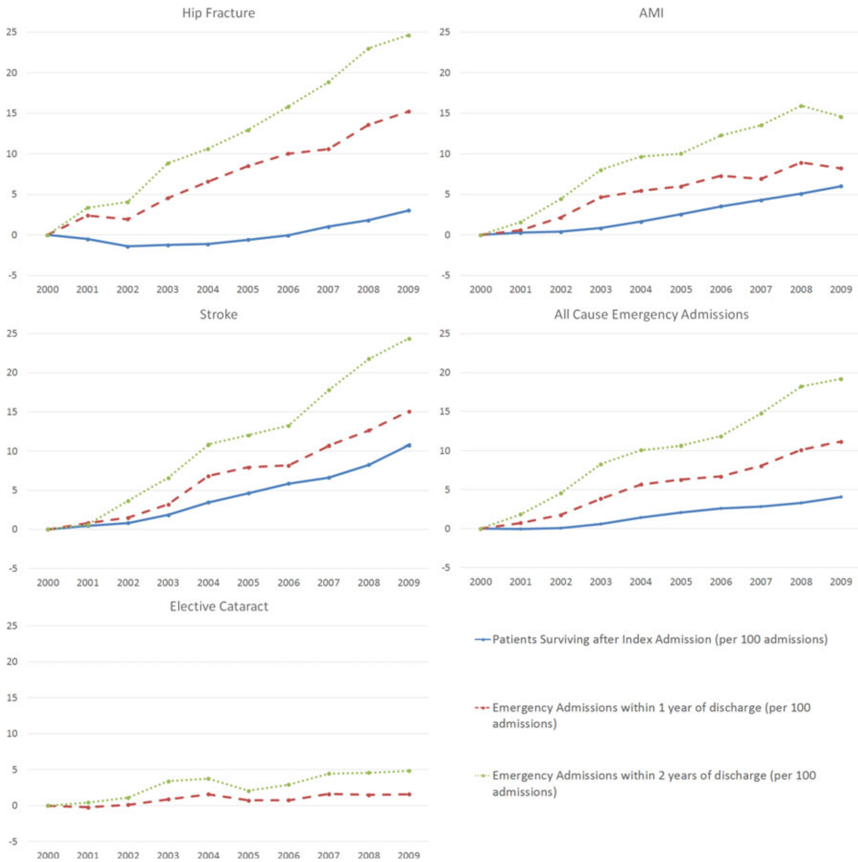
STUDY RESULTS

Table 1 reports the number of patients included in our study and their health and socioeconomic characteristics.

Figure 1 displays the variation in the number of patients surviving an index admission in each of our patient cohorts from 2000 to 2009. Values are reported in terms of additional patients surviving per 100 admissions from 2000 (the baseline year) and before any statistical adjustment for age and comorbidity. We observe a marked rise in the number of patients surviving their index admission across all patient cohorts between 2000 and 2009: +3.0 patients survive per 100 admissions for hip fractures, +6.0 patients for AMI, +10.8 for stroke, and +4.1 patients for all-cause emergency admissions. Similar trends have been reported elsewhere (Ovbiagele 2010; Smolina et al. 2012; Lisk and Yeong 2014). Figure 1 also shows the variation in the number of subsequent emergency admissions experienced by patients within 1 and 2 years of discharge from an index admission as compared with 2000 baseline year. A marked growth is observable across all cohorts between 2000 and 2009: +15.2 emergency admissions within 1 year of discharge per 100 admissions for a hip fracture, +8.2 for AMI, +15.0 for stroke, and +11.1 for all-cause emergency admissions.

Figure 2 shows similar upward trends for emergency admissions occurring within 28 days of index admissions. These trends are compared with the emergency admissions experienced by patients after an elective cataract admission, which we use as a placebo test. Emergency admissions in the placebo increased by a much more modest amount or actually dropped between the same years: +1.6 emergency admissions within 1 year of

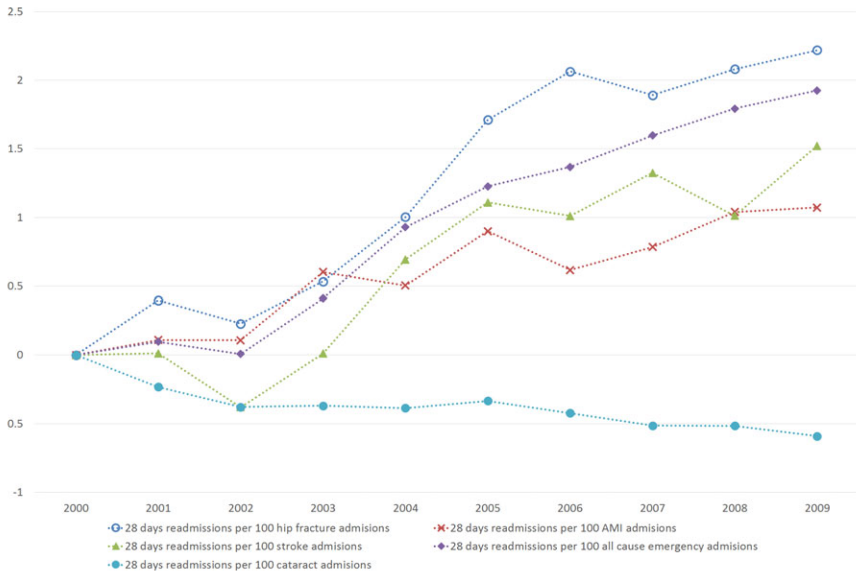
Figure 1: Patients' Survival and Subsequent Emergency Admissions Following an Index Admission [Color figure can be viewed at wileyonlinelibrary.com]



Note: Line graphs in Figure 1 report the unadjusted variation in total number of events from the baseline year 2000 and per 100 index admissions. An index admission is defined as the first hospital admission experienced by the patient between 2000 and 2009. Patients with a history of hospital admissions in the past 2 years are excluded. The blue continuous line reports the number of additional patients surviving after 30 days for every 100 index admissions as compared with the baseline year. The red dashed line reports additional emergency admissions within 1 year of discharge for every 100 index admissions as compared with the baseline year. The green dotted line reports the same information for emergency admissions occurring within 2 years.

discharge from the cataract admission (per 100 admissions) and −0.59 emergency admissions within 28 days of discharge from the cataract admission (per 100 admissions).

Figure 2: Emergency Admissions within Twenty-Eight Days of Discharge from an Index Admission [Color figure can be viewed at wileyonlinelibrary.com]



Note: Dotted lines in Figure 2 report the number of additional emergency admissions occurring within 28 days of discharge for every 100 index admissions as compared with the baseline year. An index admission is defined as the first hospital admission experienced by the patient between 2000 and 2009. Patients with a history of hospital admissions in the past 2 years are excluded.

Table 2 reports results from our statistical analysis. We use risk-adjusted models described in the Method section to estimate the effect of the variation in hospital survival rates on patients' subsequent emergency admissions within 28 days, 1 year, and 2 years from their index admission. Results are reported in terms of average marginal effects (AME), that is, the change in total emergency admissions experienced on average by each patient after a one unit increase in hospital risk-adjusted survival rates. The latter measures the number of patients surviving per 100 admissions; hence, a one unit increase represents one extra patient surviving per 100 admissions.

Results in Table 2 show that improvements in hospital survival rates had a significant effect on the number of subsequent emergency admissions across all examined cohorts. To put the scale of these effects into context, Table 3 combines results from our statistical model with the increment in hospital survival rates observed in our patient cohorts. Survival rates improved

Table 3: Increment in Emergency Admissions Explained by Improvements in Hospital Survival Rates

	Effect of+ 1 Patient Surviving on Subsequent Emergency Admissions within 1 year of Discharge, per 100 Index Admissions (A)	*Increment in Patient Surviving per 100 Index Admissions, 2000–2009 (B)*	*Effect on Emergency Admissions within 1 year, per 100 Index Admissions (C) = (A) × (B)*	*Total Index Admissions ("), 2009 (D)*	*Total Effect on Emergency Admissions within 1 year, 2009 (E) = (D) × (C)/100*	*Total Increment in Emergency Admissions, 2000–2009 (F)*	*Share of Total Increment Explained by Improvement in Hospital Survival Rates, 2000–2009 (G) = (E)/(F)*
Patient cohorts							
Hip fracture	0.6	3.02	1.81	28,267	513	n/a	n/a
AMI	0.6	6.01	3.60	31,138	1,122	n/a	n/a
Stroke	0.4	10.78	4.31	36,649	1,580	n/a	n/a
All-cause emergency	1.9	4.07	7.73	1,090,264	84,316	225,779	37.3%

Notes: (A) Estimated effects from negative binomial model (Table 2) rescaled per 100 admissions; (B) Unadjusted survival data in our patient cohorts (Figure 1); (D) Index admissions in our patient cohorts; (F) Emergency admissions in our patient cohorts.
The study sample includes patients with an index admission for a hip fracture age 65+, AMI age 55+, stroke age 55+, and all-cause emergency admissions age 55+. An index admission is defined as the first hospital admission experienced by the patient in the past 2 years. Patients with a hospital admission in the 2 years before the index admission are not included in the study. Data on admissions occurring in 1998 and 1999 are used to allow for calculating index admissions in 2000 and 2001.

by +3.02 patients per 100 hip fracture admissions between 2000 and 2009, resulting in +1.81 subsequent emergency admissions per 100 hip fracture admissions in 2009 according to our statistical model. This effect is then rescaled to our population of patients admitted in 2009 for a hip fracture, that is, 28,267 patients; thus, it can be estimated that +513 emergency admissions were generated by these patients within 1 year from their discharge as a result of improvement in their in-hospital survival rates since 2000. These additional admissions would not have occurred had survival after a hip fracture remained at its 2000 level.

Similar calculations can be made for patients in other cohorts. For AMI patients, hospital survival rates improved by +6.01 patients per 100 admissions between 2000 and 2009, resulting in +3.60 subsequent emergency admissions per 100 AMI admissions in 2009. This means that 31,138 AMI patients admitted in 2009 generated +1,122 subsequent admissions within 1 year as a result of the improvement in their in-hospital survival rates since 2000.

Between 2000 and 2009, hospital survival rates for stroke patients show an impressive improvement of +10.78 patients per 100 admissions, resulting in +4.31 subsequent emergency admissions per 100 stroke admissions in 2009. This means that 36,649 stroke patients admitted in 2009 generated +1,580 subsequent admissions within 1 year of discharge as a result of the improvement in their in-hospital survival rates since 2000.

Finally, our analysis of all-cause emergency admissions shows an even larger effect. Survival rates improved by +4.07 patients per 100 admissions between 2000 and 2009, resulting in +7.73 subsequent emergency admissions per 100 admissions in 2009. If we rescale this effect to our study population of 1,090,264 patients admitted in 2009, we can estimate that +84,316 subsequent admissions within 1 year of discharge were generated by these patients as a result of improvement in their in-hospital survival rates. In other words, there would have been 84,316 fewer emergency admissions in 2009 had hospital survival rates remained unchanged since 2000. This estimate accounts for 37.3 percent of the total increment in emergency admissions observed between 2000 and 2009 in our study population, that is, +225,779 admissions. The 2SLS instrumental variable model confirms these predictions.

We also found that improving hospital survival rates had a small but statistically significant effect on the probability that a patient experiences an emergency admission within 28 days of discharge from the index admission. Our model estimates that a hospital saving the life of one extra patient per 100 admissions increases the probability of a 28-day readmission by +0.001 for hip fracture, AMI, and stroke admissions and +0.002 for all-cause emergency admissions. The 2SLS instrumental variable model predicts a larger effect, +0.07, in the aggregated cohort of all-cause emergency admissions.

Complete results from statistical analysis are included in Appendix SA2.

DISCUSSION

This study produces new evidence explaining the surge in utilization of unplanned hospital admissions in England, which also affects similar high-income countries. We find evidence of a strong link between the improvement in hospital survival rates and the increment in utilization of emergency care. We examined a large population of patients admitted to hospital for an acute condition, such as hip fracture, AMI, and stroke, and we found that improvements in hospitals' success in saving the lives of these patients had a notable effect on their subsequent emergency admissions for any cause and occurring at any hospitals within 1 and 2 years of discharge from their index admission. We also examined this relationship in the aggregate population of patients with an initial acute admission for any cause to assess the overall impact. Our analysis shows that the marked improvement in hospital survival rates achieved between 2000 and 2009 resulted in an additional 4.04 patients surviving per 100 admissions, but with an additional 7.73 emergency admissions occurring within 1 year per 100 admissions. This accounts for 37.3 percent of the overall increment in emergency admissions observed in our sample of 10 million admissions between 2000 and 2009.

To appreciate the cost implications of our findings, we can extrapolate our estimates and apply them to recent data on the whole population of NHS patients in England. Between 2000 and 2012, a total increment of 1.63 million emergency admissions is reported (NAO 2013) with a total incremental cost to hospitals of £3.45 billion[3] ($5.45 billion). Our analysis suggests that 37.3 percent of this amount may be explained by the variation in hospital mortality rates that occurred between 2000 and 2009, that is, £1.29 ($2.04) billion or 10.32 percent of the total cost of emergency admissions reported by all NHS hospitals in 2012 (£12.5 billion).

Although it seems plausible to argue that more patients surviving hospital surgery might generate future costs for the health system, there is a dearth of evidence on the magnitude of such effects. So far as we are aware, our study is the first empirical investigation attempting to quantify the relationship between improved survival and future utilization of care in the context of the rising number of emergency admissions affecting many modern health systems. Previous studies have shown evidence of a weak association between hospitals' survival rates and 28-day readmission rates suggesting that high-performing hospitals in survival rates have no better or worse performance in 28-day readmissions as compared with other peers (Gorodeski, Starling, and Blackstone 2010; Krumholz et al. 2013; Laudicella, Li Donni, and Smith 2013; Brotman et al. 2016; Sabbatini et al. 2016). Such evidence is consistent with our findings. A number of studies have shown that end-of-life care is one of the main drivers in the per capita costs of acute care (Seshamani and Gray 2004; Raitano 2006). Such evidence is consistent with our findings that saving the life of patients who are at high risk of dying after an acute admission generates additional demand for acute care in the short term. This process derives from patient mortality risk at the point of the first admission (i.e., the index admission) and seems independent of patient age. Our study also finds a positive effect of improving hospital survival rates on emergency admissions occurring within 28 days of initial admission. Although the average effect is small, it is statistically significant in all cohorts, and it should receive careful consideration. As hospitals continue to improve their performance in survival rates, discharged patients may become increasingly frail. Therefore, the effect that we found between 2000 and 2009 is likely to be magnified in future years. Moreover, the average effect could become more marked for better performing hospitals at the top end of the distribution of survival rates.

Our findings have important policy implications. Policymakers in the United States and England have adopted a series of measures to contain emergency admissions, informed by the common assumption that a significant share of emergency admissions is avoidable and driven by hospitals' financial incentives and inefficiencies. In England, the marginal rate rule reduces hospital payments to 30 percent of the prospective tariff for emergency admissions exceeding the hospital level of activity in past years. The marginal rate rule results in a total of £530 ($837)[4] million lost income per year, an average of £3.2 ($5) million lost income per hospital trust (Foundation Trust Network 2013). A similar policy imposes a reduction in payments for emergency admissions occurring within 28 days of a previous admission for NHS patients in England and Medicare patients in the United States. In England, total penalties associated with 28-day emergency readmissions are estimated to cost hospitals £584 ($915) million in lost income per year, which is an average of £4 ($6) million per hospital trust (Foundation Trust Network 2011). Our study shows evidence that hospital volumes of emergency admissions and readmissions are affected in part by their performance in reducing mortality rates.

The assumption that health care providers can directly control their flow of emergency admissions may therefore to some extent be flawed. Indeed, current policies may generate unwanted consequences for the health system, by draining funding from high-performing hospitals that are particularly successful in saving their patients' lives (Laudicella, Li Donni, and Smith 2013; Gu et al. 2014).

Finally, the emergency care services of the NHS and many other health systems are coming under great strain, and there have been widespread calls for additional hospital funding to cope with an increasing demand for services. Hospitals have over time succeeded in saving the lives of an increasing number of patients who are likely to be frail and at high risk of accessing emergency care in the future. Managing and targeting the health conditions of these patients after their hospital discharge may, therefore, be a key to containing a large element of the demand for emergency services in the future.

Study Strengths and Limitations

A number of policy reforms are likely to have generated different shocks in hospital survival rates over the period examined by our study supporting the identification of our statistical models. Among the most relevant are the introduction of the Healthcare Commission and hospitals ratings and league tables "naming and shaming" poor-performing hospitals (Kmietowicz 2001); the introduction of the Foundation Trust status giving hospitals new autonomies and financial incentives in providing quality of care (Lewis 2005); and the impact of the Patient Choice and Hospital Competition reform on quality of care as measured by a reduction in hospital mortality rates (Gaynor, Moreno-Serra, and Propper 2013). Finally, a series of national clinical audits periodically reviewed the quality of care for hip fractures, AMI, and stroke patients, highlighting large geographical variation and making recommendations for improvements to hospitals with poor standards (Herrett et al. 2010; Party et al. 2011; Treml et al. 2011).

We undertook a number of tests for the hypothesis that unobservable confounding factors might be responsible for the effect of hospital survival rates on patients' future emergency admissions identified in our study. These should be factors associated with both within-hospital variation in survival rates over time and within-hospital variation in volume of emergency admissions. For instance, hospitals might have lowered their threshold for emergency admissions and improved their survival rates as a result of large investments in emergency services. Moreover, confounding effects might be generated by policy changes occurring over the examined period. For instance, the introduction of a new hospital reimbursement scheme based on prospective case payments might have incentivized hospitals to lower their threshold for emergency admissions in order to increase revenue; the introduction of waiting time targets in A&E might have also lowered the threshold for an emergency admissions by reducing the time for assessing patient conditions before a decision to admit. It is worth noting that, although these policies might be responsible for part of the rise in emergency admissions over time *across all hospitals*, they have to be correlated with within-hospital variation in survival rates over time in order to be a confounding factor to our analysis. In other words, hospitals that are more successful in improving their survival rates should also be more responsive to these policies as compared to other hospitals.

We are confident that such concerns are unlikely to be a major factor. First, we found evidence that hospital survival rates predict future emergency admissions in three separate cohorts of patients with an index admission for three distinct life-threatening conditions as well as in an aggregated cohort of all-cause admissions. Second, we found no evidence that hospital survival rates predict future emergency admissions in a placebo cohort of patients age 65+ admitted for a nonacute and non-life-threatening condition, but at risk of using emergency care in the future due to their age. Finally, we used an instrumental variable approach and controlled for some of the potential unobservable confounders described above using exogenous shocks to hospital survival rates generated by patients admitted over bank holidays and weekends. The 2SLS model predictions are similar to the non-IV model for one and two-year subsequent admissions, while a larger effect is predicted for 28-day admissions.

We tested for the timing of the effect of improvements in hospital survival rates by including a one-year lagged covariate in the models (see Appendix SA2). We find the effect is entirely captured by the nonlagged covariate. However, this should not be interpret as evidence of a contemporaneous effect, as the model-dependent variable captures a share of admissions generated in $t + 1$ and $t + 2$ years from when hospital survival rates are calculated.

We use hospital administrative data extracted from the HES database that are not collected for the purpose of our study. Patient secondary diagnoses are not reported consistently over the 10-year period examined and data quality may vary across hospitals (Mohammed et al. 2009). To mitigate these limitations, we made a number of sensitivity analyzes including and excluding patient comorbidities and using adjusted and unadjusted hospital mortality rates.

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Notes

1. The hypothesis of week instrument was rejected using Kleibergen-Paap Wald rk F statistic for robust standard errors.
2. Bank holiday and weekend admissions are more likely to be nondeferrable and present higher mortality risk than weekday admissions, providing further support to the identification of our IV model. However, they could also be correlated with unmeasured determinants of the risk of subsequent readmissions after discharge, invalidating our IV model. A simple correlation between weekend admissions and readmissions is likely to be contaminated by the effect of the mortality risk associated with weekend admissions. We tested for a direct effect of weekend admissions on postdischarge readmissions by including both the instrument and the endogenous variable in our models. We find no evidence of an effect of weekend admissions on postdischarge readmissions (see Appendix SA2).
3. Assuming an average cost of £2,118 per emergency admission reported by the Department of Health in Reference costs 2012–13.
4. £1 = $1.58 average exchange rate between April 2012 and March 2013.

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Supporting Information

Additional supporting information may be found online in the supporting information tab for this article:

Appendix SA1: Author Matrix.
Appendix SA2: Technical Details of the Empirical Analysis.