
Quality of Care and Mortality

Do Hospitals Provide Lower Quality Care on Weekends?

David J. Becker

Objective. To examine the effect of a weekend hospitalization on the timing and incidence of intensive cardiac procedures, and on subsequent expenditures, mortality and readmission rates for Medicare patients hospitalized with acute myocardial infarction (AMI).

Data Sources. The primary data are longitudinal, administrative claims for 922,074 elderly, nonrural, fee-for-service Medicare beneficiaries hospitalized with AMI from 1989 to 1998. Annual patient-level cohorts provide information on ex ante health status, procedure use, expenditures, and health outcomes.

Study Design. The patient is the primary unit of analysis. I use ordinary least squares regression to estimate the effect of weekend hospitalization on rates of cardiac catheterization, angioplasty, and bypass surgery (in various time periods subsequent to the initial hospitalization), 1-year expenditures and rates of adverse health outcomes in various periods following the AMI admission.

Principal Findings. Weekend AMI patients are significantly less likely to receive immediate intensive cardiac procedures, and experience significantly higher rates of adverse health outcomes. Weekend admission leads to a 3.47 percentage point reduction in catheterization at 1 day, a 1.52 point reduction in angioplasty, and a 0.35 point reduction in by-pass surgery ($p < .001$ in all cases). The primary effect is delayed treatment, as weekend-weekday procedure differentials narrow over time from the initial hospitalization. Weekend patients experience a 0.38 percentage point ($p < .001$) increase in 1-year mortality and a 0.20 point ($p < .001$) increase in 1-year readmission with congestive heart failure.

Conclusions. Weekend hospitalization leads to delayed provision of intensive procedures and elevated 1-year mortality for elderly AMI patients. The existence of measurable differences in treatments raises questions regarding the efficacy of a single input regulation (e.g., mandated nurse staffing ratios) in enhancing the quality of weekend care. My results suggest that targeted financial incentives might be a more cost-effective policy response than broad regulation aimed at improving quality.

Key Words. Temporal variation, myocardial infarction, Medicare beneficiaries

The fact that patients admitted to the hospital on a weekend are more likely to die than patients admitted to the hospital on a weekday has stimulated

renewed debate over the role of regulation versus incentives in enhancing the quality of health care. The apparent differences in weekend versus weekday care are substantial. A recent study found that, for certain medical conditions, patients admitted on the weekend were over 15 percent more likely to die in the hospital than patients admitted during the week (Bell and Redelmeier 2001). Because hospitals employ fewer nurses and other support staff on weekends, some researchers and policy makers attribute this “weekend effect” to hospitals’ reductions in weekend staffing, and have recommended mandatory staffing legislation as a solution.

In this paper, I investigate an alternative hypothesis: that the weekend effect is caused by the delayed provision of specific intensive treatments which may be difficult to remedy with regulation alone. Under Medicare’s reimbursement system, payments to hospitals are not related to the timeliness of treatment, or its suitability for a particular patient’s health needs. Hospitals are reimbursed according to a patient’s diagnosis-related group (DRG), which depends both on the patient’s illness and the treatments provided by the hospital (McClellan 1997). The fixed-price DRG payment is determined by the mean cost across all hospitals of treating similar patients, and does not vary with the actual costs of treating a specific patient. As providing intensive treatment on the weekend may entail fixed costs, or higher marginal costs, it may be optimal for a hospital that sought to maximize revenues over costs to decline to do so. If the rapidity of intensive treatment has important effects on health outcomes for a sufficiently large number of patients, then this might not be in society’s best interests.

Distinguishing between these hypotheses is a special case of an important general health policy problem: should observed shortfalls in quality be addressed with an input regulation or with a reimbursement system that rewards superior performance? Several states have passed legislation requiring that hospitals employ a minimum number of nurses and other support staff per patient at all times, which has substantially increased operating costs of hospitals. If understaffing causes the weekend effect, then intervention in the form of staffing ratios could be socially optimal. But if the weekend effect is caused by another form of inappropriate treatment not easily remedied through regulation, then stronger reimbursement incentives would be a more cost-effective policy response.

Address correspondence to David J. Becker, Ph.D., Assistant Professor, Department of Health Care Organization and Policy, School of Public Health, University of Alabama at Birmingham, RPHB 330, 1530 3rd Avenue South, Birmingham, AL 35294-0022.

In this paper, I examine the impact of a weekend admission on the timing and incidence of specific intensive treatments received by elderly Medicare beneficiaries diagnosed with heart attack, or acute myocardial infarction (AMI). I explore how the timing of the three major invasive treatments associated with AMI—cardiac catheterization (CATH), angioplasty, and bypass surgery—varies for patients admitted on weekends and weekdays. I also examine the effects of weekend hospitalization on aggregate treatment intensity (1-year inpatient and outpatient expenditures) and adverse health outcomes (including mortality and cardiac related readmissions).

WEEKEND EFFECT: BACKGROUND

The existing empirical literature provides conflicting evidence regarding the relationship between weekend hospitalization, medical treatment decisions, and quality of care. Several studies conclude that weekend hospitalization is associated with worse health outcomes. The earliest research (MacFarlane 1978; Mangold 1981) examines the effect of weekend birth on neonatal or perinatal mortality, and finds higher mortality on the weekend than during the week. Bell and Redelmeier (2001) find significantly higher rates of in-hospital mortality for patients hospitalized on the weekend with one of three conditions (ruptured abdominal aortic aneurysm, acute epiglottitis, and pulmonary embolism) whose treatment was identified by the authors as being particularly sensitive to variations in staffing. The study also examines differences between weekend and weekday mortality for the top 100 diagnoses associated with in-hospital death. For 23 of these conditions the authors find evidence of significantly higher in-hospital mortality for patients admitted on the weekend. These studies are supported by other work that finds a positive correlation across hospitals between annual average staff-to-patient ratios and quality of care (Aiken et al. 2002; Needleman et al. 2002). Staffing is generally lower on the weekend, so evidence of a positive correlation between staffing and quality is consistent with an adverse weekend effect that operates through staffing.

Other researchers have argued that higher weekend mortality, at least in part, reflects differences in the unobservable characteristics of patients hospitalized on weekends versus weekdays. Because of the preferences of physicians and/or their patients for scheduled visits during the week, elective admissions are less likely to occur on weekends, which would lead patients admitted on the weekend to be sicker than those admitted during the week.

More recent studies of neonatal mortality (Dowding et al. 1987; Gould et al. 2003) have found that higher weekend mortality results from the increased incidence of low-risk elective deliveries during the week, and the correspondingly higher proportion of high-risk, spontaneous weekend births. Halm and Chassin (2001) note that of the 23 conditions for which Bell and Redelmeier find evidence of higher in-hospital mortality, over half are cancers. Differences in short run outcomes between weekend and weekday patients admitted with *nonacute* illnesses are likely to be caused by unobserved differences in the health of such patients on admission. Consistent with this hypothesis, Dobkin (2003) finds no evidence of a weekend effect using data similar to those of Bell and Redelmeier, after controlling for patient heterogeneity.

A limited number of clinical studies have examined how day of week (and time of day) of admission might influence the use of specific medical treatments and health outcomes. A recent study by Magid et al. (2005) compares door-to-drug and door-to-balloon (angioplasty) times for patients admitted with acute ST-segment elevation myocardial infarction during “regular-hours” (Monday–Friday, 7:00 A.M.–5:00 P.M.) and “off-hours.” Conditional on receiving reperfusion within 6 hours of hospitalization, patients admitted during off-hours periods face longer door-to-balloon times, but no significant difference in door-to-drug times. The authors find evidence of elevated in-hospital mortality for the pooled sample of off-hours patients receiving angioplasty or drug reperfusion, but do not find a significant effect on the in-hospital mortality for patients receiving angioplasty alone. By limiting their sample to patients who received reperfusion therapy within 6 hours, their analysis considers only the effects of delay conditional on receiving treatment, and does not explore how hospitalization during off-hours affects the mix of treatment patients receive.

Other research has explored the role of physician bias or “convenience” in medical decision making. Fraser et al. (1987), use data on all births at Royal Victoria Hospital in Montreal, Canada from 1978 to 1984 to examine the relationship between time of day and the rate of cesarean section. Controlling for duration of labor, they find significantly higher rates of cesarean section for dystocia during the evening hours (6:00 P.M.–11:59 P.M.). This is consistent with the convenience hypothesis—physicians prefer to perform these procedures in the evening hours rather than during the nighttime hours of sleep or the daytime hours when they have scheduled appointments with patients. Finally, Varnava et al. (2002) examine the effects of weekend hospital service reductions on the treatment of AMI patients in a British hospital. The authors observe substantially lower rates of discharge on weekends, suggesting that

attending physicians prefer to be present on the day of discharge, but also prefer to be absent from the hospital on weekends.

In summary, one arm of existing research has documented correlations between mortality and either weekend admission or staffing, but has not identified whether these correlations are causal, and if they are, the mechanism through which they operate. In addition, the measures of outcomes used in this work have been either so short term as to fail to measure the likely consequences of weekend admission, or subject to important biases. Another arm has estimated the consequences of weekend admission on specific treatments, but has largely failed to explore the impact of these changes in treatment decisions for the cost and quality of care. Although policy makers have hypothesized that generalized reductions in staffing may be a source of excess weekend mortality in hospitals, there is little empirical evidence to support this claim. Without such evidence, the basis for regulatory intervention is weak at best.

DATA

The principal data used in my analysis are longitudinal claims from the 100 percent Medicare Provider Analysis and Review files, which are used to construct cohorts of nonrural, elderly Medicare beneficiaries hospitalized with a “new” diagnosis of AMI in 1989–1998. My definition of a “new” AMI excludes patients who had been hospitalized with AMI in the previous 365 days. I also exclude patients who were in HMOs, patients suffering from end-stage renal disease, and patients who were not initially admitted through the hospital’s emergency room. The claims data are matched to patient demographic characteristics obtained from the Health Care Financing Administration’s¹ HISKEW enrollment files, which also provide validated dates of death from Social Security Administration death reports. These death dates allow the construction of mortality indicators at various periods following the initial hospitalization.

From the Medicare claims, I construct measures of medical expenditures, treatment intensity, and health outcomes in various periods subsequent to the initial hospitalization. I calculate total inpatient and outpatient expenditures (expenditures = Medicare hospital reimbursement + deductibles + co-payments) in the year following the initial admission with AMI. Claims data also are used to construct a series of indicator variables denoting whether, and when, a patient received the three intensive procedures commonly used in the treatment of AMI—cardiac catheterization (CATH), percutaneous translu-

minimal coronary angioplasty (PTCA or angioplasty), and coronary artery bypass graft surgery (CABG or bypass). Using administrative claims, I am able to observe the day on which the procedure was performed, but not the specific time of day. Claims data are also used to construct measures of the occurrence of cardiac complications, including a readmission with a primary diagnosis of AMI or congestive heart failure (CHF) within 365 days of the initial admission. In constructing readmission variables I exclude transfers and readmissions within 30 days, which may reflect treatment of the initial health episode.

I use the claims to create several measures of patient health status at the time of the AMI admission. First, I calculate total Medicare inpatient expenditures for any illness in the prior year and create an indicator variable denoting whether or not the patient received any inpatient care during that period. Second, I use inpatient and outpatient claims from the prior year to construct the Charlson comorbidity index (Charlson et al. 1987) at the time of the AMI admission.² In addition to the traditional Charlson index, I also construct an adapted Charlson index which uses the odds ratios obtained from a logistic regression of 1-year mortality on each of the Charlson conditions as weights. While this adapted index was intended to provide a more relevant measure of health status for AMI patients, my results were not sensitive to the use of either the traditional or adapted index. Finally, I control for the physical location of the infarction using the fourth digit of the primary ICD-9 code reported on the patient's initial AMI claim.

EMPIRICAL MODELS

The fundamental question this paper asks is straightforward: does weekend hospitalization lead to differences in medical treatments and health outcomes? To answer this question I use annual cohorts of elderly Medicare patients who were hospitalized, through the emergency room, with a primary diagnosis of AMI between 1989 and 1998. In zip code k during year $t = 1, \dots, T$, I analyze the effect of weekend admission on expenditures, intensive procedures and outcomes for individuals $i = 1, \dots, N_{kt}$ who are admitted to hospital j with an incident occurrence of AMI. Each patient has a vector of observable characteristics X_{it} : a nine-dimensional vector denoting the location of the infarction (ICD-9 410.0 \times "anterolateral wall" is the omitted category), four age indicator variables (70–74, 75–79, 80–89, 90–99; and 65–69 years is the omitted group), gender, and a complete set of interaction effects between age and gender. Each patient has two measures of health status upon admission: A_{it}

which denotes an inpatient admission in the prior year; and two indicator variables C_{it} which indicate whether the patient's Charlson comorbidity index at the time of admission is "very high" (Charlson > 90th percentile) or "high" ($0 < \text{Charlson} < 90\text{th percentile}$), with Charlson = 0 being the omitted group.

I examine the impact of weekend hospitalization on rates of the three intensive cardiac procedures—catheterization, angioplasty, and bypass—in four time periods subsequent to the initial hospitalization. To assess the effect on immediate treatment, I use indicator variables denoting whether the patient received a given procedure on the day of ($P0_{it}$), or within 1 day of, their hospitalization ($P1_{it}$). The same day rates best capture the effect on immediate procedure use, particularly primary angioplasty which is typically performed within hours of hospital presentation. The 1-day rates provide a broader window for immediate treatment, accounting for the fact that some patients admitted late in the evening may receive treatment within several hours, but on the next calendar day. In order to examine the persistence of delays in treatment, I also estimate the effect of weekend admission on variables indicating whether each of these cardiac procedures was received within 7 days but not within 1 day of the initial admission (procedure was received 2–7 days after admission), $P7_{it}$. Additionally, I estimate the effect of weekend admission on 365-day (cumulative) procedure rates $P365_{it}$ to identify whether differences in procedure rates persist at 1 year. I examine the effect of weekend admission on Medicare expenditures using the logarithm of Y_{it} where Y is total expenditures in the year after and including the admission to the hospital with AMI. I examine separately the effects of weekend admission on inpatient and outpatient expenditures, setting the logarithm of outpatient expenditures to zero for patients with no outpatient utilization in the following year. Lastly, the patient has health outcomes O_{it} where $O = 1$ indicates the patient suffered an adverse health outcome in various periods following the initial admission, including mortality, readmission with AMI, and readmission with CHF.

To identify the effects of a weekend admission on expenditures, treatments and health outcomes, I estimate linear models of the following form:

$$\frac{P_{it}}{O_{it}} = \alpha_j \pi + \sigma_t M_k \chi + X_{it} \psi + A_{it} \delta + C_{it} \lambda + W_{it} \beta + \varepsilon_{ikt} \quad (1)$$

where α_j is a hospital fixed effect, σ_t is a year fixed effect which is allowed to vary with M_b a five-dimensional vector of indicators denoting the size of patient i 's MSA (largest MSAs with population > 2,500,000 are the omitted group), and ε_{ikt} is an error term where $E(\varepsilon_{ikt} | \dots) = 0$. By including hospital

fixed effects I control both for variation in practice patterns across hospitals and the nonrandom sorting of patients across hospitals.

The coefficients of interest are β_{P0} , β_{P1} , β_{P7} , β_{P365} , β_Y and β_O , which reflect differences in treatments, expenditures, and health outcomes for patients admitted on the weekend. The coefficients from cardiac procedure models make it possible to parse the weekend effect into “delay” and “cumulative use” components. The “delay” attributable to weekend admission is defined as $\beta_{P365} - \beta_{P1}$. Delay in treatment occurs when a reduction in 1-day procedure rates for weekend patients is offset by procedures these patients receive later in time. Estimates of β_{P7} make it possible to compare the proportion of weekend patients experiencing a brief delay (β_{P7}) versus a longer delay ($\beta_{P365} - \beta_{P1} - \beta_{P7}$) in receiving invasive procedures. The “cumulative use” component of weekend admission is simply β_{P365} .

The coefficients β_Y and β_O can be used to assess the welfare consequences of weekend admission. Even if weekend admission reduces treatment intensity, the weekend effect would be welfare improving if such a reduction in intensity reduced total expenditures but did not increase rates of adverse outcomes ($\beta_Y < 0$, $\beta_O \leq 0$). If weekend admission leads to increased total expenditures without decreased rates of adverse health outcomes ($\beta_Y > 0$, $\beta_O \geq 0$), then it is welfare reducing. If weekend admission leads to reduced expenditures and increased rates of adverse outcomes ($\beta_Y < 0$, $\beta_O > 0$), or increased expenditures and reduced rates of adverse outcomes ($\beta_Y > 0$, $\beta_O < 0$), then I can calculate the implied cost-effectiveness per year of life saved, or year of cardiac health achieved, of weekend versus weekday admission.

In order to further investigate the mechanism through which the weekend effect operates, I examine whether the results differ by patients’ ex ante health status. To do this, I estimate models that include an interaction between weekend admission and the prior inpatient admission indicator, A_{it} . These models have the following form:

$$\ln\left(\frac{P_{it}}{O_{it}}\right) = \alpha_j\pi + \sigma_t M_{k\chi} + X_{it}\psi + A_{it}\delta + C_{it}\lambda + W_{it}\beta + A_{it}W_{it}\theta + \varepsilon_{ikt} \quad (2)$$

In these models, the coefficient θ reflects the difference in treatments, expenditures and outcomes for sicker versus healthier patients who are admitted with AMI on a weekend.

β and θ represent causal effects only under the assumption that $E(\varepsilon|W) = 0$, i.e., there is no unobserved heterogeneity between AMI patients admitted on weekends and during the week. I take a number of steps to control

for patient heterogeneity. First, in addition to choosing a well-defined acute illness for study to minimize the possibility of unobserved differences between weekday and weekend admissions, I further limit my sample to AMI patients initially admitted to the hospital through the emergency room. Restricting my analysis to ER admits reduces the potential for unobserved heterogeneity of weekend versus weekday patients, because every patient admitted through the ER is likely to have AMI symptoms that are sufficiently severe to require immediate hospitalization. I also construct an extensive set of medical history variables from patients' prior hospital utilization, including the Charlson comorbidity index and an indicator variable denoting whether the patient was admitted to an inpatient hospital in the year before their index admission. The inclusion of hospital fixed effects controls for differences in unobservable patient characteristics across hospitals.

Finally, I estimate the weekend effect comparing the treatments, expenditures, and health outcomes of weekend plus Monday admissions to those of Tuesday through Friday admissions. To the extent that the preferences of patients and/or their physicians for weekday over weekend admissions lead to delaying Saturday or Sunday admissions until Monday for patients with less severe forms of illness, a Saturday/Sunday/Monday versus Tuesday–Friday weekend effect would be free of bias owing to unobserved patient heterogeneity, even if a Saturday/Sunday versus Monday–Friday weekend effect were not. Even in the presence of differential unobserved heterogeneity arising from the delayed admission of patients with weekend symptom onset, the underlying health status of patients admitted on Saturday–Monday versus Tuesday–Friday would be the same.

RESULTS

Table 1 provides descriptive statistics showing the dramatic increase in treatment intensity during my study period of 1989–1998. With rates of reimbursement for a given treatment relatively constant, the rapid growth in inpatient expenditures is the result of increased treatment intensity during the study period (McClellan 1997). The table reveals sharp increases in the rates of all three major invasive procedures used to treat AMI patients—catheterization, angioplasty, and bypass surgery. Along with the increase in the volume of catheterizations, there also was a move toward earlier intervention, with the fastest growth occurring in 0- and 1-day CATH rates. In addition to the diffusion of these invasive procedures for treating AMI, the study period was marked by the diffusion of noninvasive technologies as well, such as

Table 1: Trends in Characteristics of Elderly Medicare AMI Patients Admitted through the Emergency Room

	<i>1989–1998 Aggregate Mean</i>	<i>1989 Mean</i>	<i>1998 Mean</i>	<i>1989–1998 Change</i>
1-year inpatient expenditures (1993 \$) (standard deviation)	\$19,196 (20,439)	\$16,016 (16,997)	\$22,742 (22,862)	42.0%
1-year outpatient expenditures (1993 \$) (standard deviation)	\$798 (2,107)	\$520 (1,246)	\$868 (2,007)	66.9%
0-day CATH rate	6.5%	2.5%	11.1%	8.6%
1-day CATH rate	10.3%	4.0%	17.5%	13.5%
7-day CATH rate	33.2%	16.4%	43.1%	26.7%
365-day CATH rate	42.4%	28.8%	48.9%	20.1%
0-day PTCA rate	3.9%	1.3%	7.1%	5.8%
1-day PTCA rate	5.0%	1.7%	9.3%	7.6%
7-day PTCA rate	11.8%	4.2%	19.3%	15.1%
365-day PTCA rate	16.4%	8.0%	22.9%	14.9%
0-day CABG rate	0.5%	0.2%	0.8%	0.6%
1-day CABG rate	0.9%	0.4%	1.5%	1.1%
7-day CABG rate	6.3%	2.4%	9.0%	6.6%
365-day CABG rate	14.4%	10.6%	14.6%	4.0%
30-day mortality rate	20.8%	23.8%	20.0%	–3.8%
90-day mortality rate	26.0%	29.2%	25.7%	–3.5%
180-day mortality rate	30.0%	33.0%	30.3%	–2.7%
1-year mortality rate	35.2%	38.2%	36.3%	–1.9%
1-year AMI readmission rate	5.6%	5.4%	5.9%	0.5%
1-year CHF readmission rate	9.2%	8.8%	9.2%	0.4%
1-year prior inpatient expenditures (1993 \$) (standard deviation)	\$3,289 (8,394)	\$2,876 (6,937)	\$4,178 (10,035)	45.3%
Inpatient admission 365 days before AMI	30.1%	30.1%	32.3%	2.2%
Charlson comorbidity index	0.760	0.607	0.953%	57.0%
Age	77.136	76.755	78.042	1.7%
Black	6.7%	6.6%	7.0%	0.4%
Female	51.3%	51.7%	52.5%	0.8%
N	922,074	90,713	94,382	

Notes: For dichotomous variables the means (and their changes) are reported in percentage points; for continuous variables changes are reported in percents. For 1996, I use a 20 percent sample which is upweighted to construct aggregate means. Outpatient expenditures are calculated using 20 percent sample in all years.

AMI, acute myocardial infarction; PTCA, percutaneous transluminal coronary angioplasty; CABG, coronary artery bypass graft; CHF, congestive heart failure; CATH, cardiac catheterization.

thrombolytic drugs, ACE inhibitors, β -blockers, and aspirin. All of these technologies have contributed to the steady drop in 1-year mortality rates (McClellan and Noguchi 1998). The increased survival rate has led to only

slightly higher 1-year heart failure and AMI readmission rates. Over my study period, patients have grown older, and less likely to live in large MSAs, with racial and gender composition roughly constant.

Table 2 presents separate descriptive statistics for weekend and weekday patients admitted during the entire 10-year study period. This table provides evidence suggestive of a weekend effect on the rapidity and probability of procedure use. Weekend patients are significantly less likely to receive immediate catheterization, angioplasty or bypass surgery on the day of their admission with AMI ($p < .001$). The gap in procedure rates is largest 1 day subsequent to the initial hospitalization, as Saturday patients continue to experience further weekend induced delays in treatment. Data not presented in the table show that weekend-weekday differences in immediate (both 0 day and 1 day) procedure rates became more pronounced over the duration of the study period, as these invasive procedures came to be used both more frequently and earlier in the treatment of AMI. The weekend-weekday differences in procedure rates decrease with time from the initial hospitalization, with most of this narrowing occurring during the first week. Only small reductions in rates of angioplasty and catheterization persist at 1 year. In turn, weekend patients experience higher 1-year mortality ($p < .10$), though the effects on cardiac readmission rates is somewhat mixed.

Table 2 also shows no evidence of systematic differences in the demographic characteristics or observable measures of ex ante health of weekend versus weekday patients. Weekend patients are no more likely to have been hospitalized in the prior year, and have mean prior year inpatient expenditures and Charlson indices that are statistically indistinguishable from those of weekday patients. For the purpose of this table only, I matched hospital characteristic data from the American Hospital Association Survey to patient-level data based on each patient's initial hospital of admission. These hospital characteristics show that there are no systematic differences in the quality of hospitals to which weekend versus weekday patients are admitted. While weekday patients are more likely to be admitted to large, teaching hospitals, weekend patients are more likely to be admitted to hospitals with catheterization laboratories.

In recent work, Dobkin (2003) argues that nonuniform incidence, resulting from patients' and/or physicians' preferences for weekday admission, indicates the presence of selection that will generate upward biased estimates of the impact of weekend admission on health outcomes. Table 2 reveals that the proportion of patients admitted on the weekend ($263,068/922,074 = 28.5$ percent) is exactly what would be expected if the true incidence of AMI were

Table 2: Descriptive Statistics for Elderly Medicare AMI Patients Admitted on Weekends and Weekdays (1989–1998)

	<i>Weekdays</i>	<i>Weekends</i>	<i>Difference p-Value*</i>	
<i>Patient-level variables</i>				
1-year inpatient expenditures (1993 \$) (standard deviation)	\$19,014 (19,494)	\$18,932 (19,261)	– \$82	0.066
1-year outpatient expenditures (1993 \$) (standard deviation)	\$754 (1,718)	\$737 (1,647)	– \$41	<0.001
0-day CATH rate	6.78%	4.92%	– 1.86%	<0.001
1-day CATH rate	10.77%	7.65%	– 3.12%	<0.001
7-day CATH rate	32.62%	32.04%	– 0.58%	<0.001
365-day CATH rate	41.78%	42.06%	0.28%	0.008
0-day PTCA rate	4.00%	2.91%	– 1.09%	<0.001
1-day PTCA rate	5.20%	3.81%	– 1.39%	<0.001
7-day PTCA rate	11.57%	11.01%	– 0.56%	<0.001
365-day PTCA rate	16.12%	15.86%	– 0.26%	0.018
0-day CABG rate	0.57%	0.41%	– 0.16%	<0.001
1-day CABG rate	1.03%	0.73%	– 0.30%	<0.001
7-day CABG rate	6.14%	5.98%	– 0.16%	0.068
365-day CABG rate	14.23%	14.43%	0.20%	0.013
30-day mortality rate	20.84%	20.97%	0.13%	0.033
90-day mortality rate	26.05%	26.17%	0.12%	0.324
180-day mortality rate	29.96%	30.05%	0.09%	0.680
1-year mortality rate	35.21%	35.39%	0.18%	0.070
1-year AMI readmission rate	5.62%	5.54%	– 0.08%	0.058
1-year CHF readmission rate	9.11%	9.24%	0.13%	0.133
1-year before index inpatient expenditures (1993 \$) (standard deviation)	\$3,266 (7,988)	\$3,257 (7,998)	– \$9	0.625
Inpatient admission 365 days before AMI	30.15%	30.20%	0.05%	0.345
Charlson comorbidity index	0.756	0.755	– 0.001	0.840
Age	77.129	77.076	– 0.053	<0.001
Black	6.69%	6.70%	0.01%	0.929
Female	51.26%	51.21%	– 0.05%	0.386
<i>Hospital-level variables</i>				
Large size (>300 beds)	29.49%	28.88%	– 0.61%	<0.001
Medium size (100–300 beds)	58.77%	59.01%	0.24%	0.078
Small size (<100 beds)	11.74%	12.11%	0.37%	<0.001
Teaching	29.42%	28.26%	– 1.16%	<0.001
For-profit	9.64%	9.70%	0.06%	0.142
Nonprofit	82.46%	82.13%	– 0.33%	<0.001
Public	7.90%	8.17%	0.27%	<0.001
Catheterization laboratory	66.32%	66.54%	0.22%	0.066
High AMI volume hospital	72.35%	72.15%	– 0.20%	0.333
N	659,006	263,068		

Notes: *The *p*-values come from a *t*-test of the difference in means.

AMI, acute myocardial infarction; PTCA, percutaneous transluminal coronary angioplasty; CABG, coronary artery bypass graft; CHF, congestive heart failure; CATH, cardiac catheterization.

uniform. Although there is a slight spike in the number of AMI patients admitted on Mondays, the absence of a weekend drop-off suggests that this is unlikely to be the result of less severely ill patients experiencing weekend symptom onset deferring admission until Monday. Dobkin acknowledges that AMI may be a condition where the true incidence is in fact nonuniform, as researchers have offered explanations for the increased Monday incidence ranging from heavy weekend drinking (Evans et al. 2000) to the stress associated with the start of the work week (Willich et al. 1994).

Tables 3 and 4 summarize the primary results of interest—the effect of weekend admission on procedure rates, expenditures and outcomes—under several specifications. Table 3 presents estimates and standard errors of the effects of weekend admission on 0, 1, 2–7, and 365 day cardiac procedure rates, while Table 4 presents the effects of weekend admission on inpatient and outpatient expenditures and health outcomes. These models all correspond to Equations (1) and (2) and control for patient health status, patient demographics, hospital fixed effects, and differential time trends across MSAs of differing sizes. The standard errors are based upon an estimator of the variance–covariance matrix that is consistent with the presence of heteroscedasticity.

The first row of Table 3 (specification 1a), shows evidence of substantial effects of weekend admission on the rapidity of intensive procedure use. Most of the weekend-effect induced delay in receiving CATH and PTCA occurs in the 7 days after the initial admission. Weekend admission leads to a decline in the probability of receiving CATH (PTCA) on the day of admission by 2.12 (1.18) percentage points with the weekend induced reduction in procedure rates rising further to 3.47 (1.52) at 1 day. Of that decline, 2.40 (0.80) percentage points are recovered in the 2–7 days after admission, with an additional 1.01 (0.33) percentage points recovered in the next 358 days (i.e., days 8–365).³ Expressed as a share of the number of patients in 1998 who received a CATH (PTCA) on the day of admission, weekend hospitalization delays the treatment of $19.1 = 2.12/11.1$ ($16.6 = 1.18/7.1$) percent of patients. Most of the weekend-effect delay in receiving CABG occurs after the first week; β_{P7} for CABG is small and statistically insignificant, while $(\beta_{P365} - \beta_P - \beta_{P7})$ remains large. There are no significant differences in the cumulative 365-day CATH and CABG rates for weekend versus weekday patients, although weekend patients are significantly less likely to receive PTCA within 1 year of admission with AMI.

The first row of Table 4 shows that these differences in treatment have significant consequences for both expenditures and health outcomes. Patients

Table 3: Effects of Weekend Admission on Treatment Rates for Elderly Medicare AMI Patients Admitted through the Emergency Room, 1989–1998

	<i>Dependent Variable</i>											
	<i>CATH</i> <i>0 Day</i>	<i>PTCA</i> <i>0 Day</i>	<i>CABG</i> <i>0 Day</i>	<i>CATH</i> <i>1 Day</i>	<i>PTCA</i> <i>1 Day</i>	<i>CABG</i> <i>1 Day</i>	<i>CATH</i> <i>2–7 Days</i>	<i>PTCA</i> <i>2–7 Days</i>	<i>CABG</i> <i>2–7 Days</i>	<i>CATH</i> <i>365 Days</i>	<i>PTCA</i> <i>365 Days</i>	<i>CABG</i> <i>365 Days</i>
<i>Specification 1a</i>												
Weekend admission	-2.12** (0.05)	-1.18** (0.04)	-0.19** (0.02)	-3.47** (0.06)	-1.52** (0.04)	-0.35** (0.02)	2.40** (0.09)	0.80** (0.06)	0.05 (0.05)	-0.06 (0.10)	-0.39** (0.08)	0.10 (0.08)
<i>Specification 1b</i>												
Weekend or Monday admission	-1.65** (0.05)	-0.95** (0.04)	-0.12** (0.01)	-2.25** (0.05)	-1.14** (0.04)	-0.22** (0.02)	1.15** (0.08)	0.42** (0.05)	-0.09** (0.04)	0.18** (0.09)	-0.32** (0.07)	0.14** (0.07)
<i>Specification 1c</i>												
Saturday admission	-2.17** (0.06)	-1.21** (0.05)	-0.19** (0.02)	-5.13** (0.07)	-1.88** (0.05)	-0.41** (0.03)	4.96** (0.11)	1.54** (0.08)	0.50** (0.07)	0.12 (0.13)	-0.30** (0.10)	0.20** (0.10)
Sunday admission	-2.07** (0.06)	-1.16** (0.05)	-0.19** (0.02)	-1.82** (0.08)	-1.17** (0.06)	-0.30** (0.03)	-0.13 (0.11)	0.07 (0.07)	-0.40** (0.06)	-0.25** (0.12)	-0.48** (0.10)	0.00 (0.10)
<i>Specification 2</i>												
Weekend admission	-2.38** (0.06)	-1.35** (0.05)	-0.21** (0.02)	-3.90** (0.07)	-1.73** (0.05)	-0.40** (0.03)	2.80** (0.11)	0.94** (0.07)	0.05 (0.06)	-0.00 (0.11)	-0.38** (0.10)	0.11 (0.09)
Weekend admission* inpatient	0.86** (0.10)	0.54** (0.08)	0.07** (0.03)	1.44** (0.12)	0.68** (0.09)	0.15** (0.04)	-1.31** (0.18)	-0.46** (0.11)	-0.01 (0.09)	-0.18 (0.21)	-0.06 (0.16)	-0.05 (0.15)

Notes: **Significance at 5 percent level.

*Significance at 10 percent level. All coefficients multiplied by 100 to facilitate interpretation. Coefficients are in percentage points. All models include controls for patient characteristics, hospital fixed effects and year fixed effects which are allowed to vary by six MSA size categories. Heteroscedasticity-consistent standard errors are shown in parentheses. $N = 922,074$.

PTCA, percutaneous transluminal coronary angioplasty; CABG, coronary artery bypass graft; CATH, cardiac catheterization.

Table 4: Effects of Weekend Admission on Medicare Expenditures and Patient Health Outcomes for Elderly Medicare AMI Patients Admitted through the Emergency Room, 1989–1998

	<i>Dependent Variable</i>							
	<i>ln (1-Year Inpatient Expenditure)</i>	<i>ln (1-Year Outpatient Expenditure)</i>	<i>AMI 365 Days</i>	<i>CHF 365 Days</i>	<i>Died 30 Days</i>	<i>Died 90 Days</i>	<i>Died 180 Days</i>	<i>Died 365 Days</i>
<i>Specification 1a</i>								
Weekend admission	-0.55** (0.20)	-4.40** (1.53)	-0.14** (0.05)	0.20** (0.06)	0.22** (0.09)	0.27** (0.09)	0.25** (0.10)	0.38** (0.10)
<i>Specification 1b</i>								
Weekend or Monday admission	0.16 (0.18)	-3.61** (1.39)	-0.12** (0.05)	0.20** (0.06)	0.14* (0.08)	-0.16* (0.08)	0.19** (0.09)	0.22** (0.09)
<i>Specification 1c</i>								
Saturday admission	-0.60** (0.27)	-2.72 (2.00)	-0.12* (0.07)	0.10 (0.08)	0.07 (0.11)	0.10 (0.12)	0.11 (0.13)	0.23* (0.13)
Sunday admission	-0.49* (0.26)	-6.09** (2.00)	-0.15** (0.07)	0.30** (0.08)	0.37** (0.11)	0.43** (0.12)	0.39** (0.13)	0.53** (0.13)
<i>Specification 2</i>								
Weekend admission	-0.32 (0.25)	-5.92** (1.81)	-0.01 (0.06)	0.18** (0.07)	0.18* (0.12)	0.27** (0.11)	0.28** (0.11)	0.36** (0.12)
Weekend admission* inpatient Admission in prior year	-0.76* (0.43)	5.06 (3.36)	-0.41** (0.12)	0.06 (0.15)	0.14 (0.19)	-0.02 (0.21)	-0.09 (0.22)	0.06 (0.22)

Notes: **Significance at 5 percent level.

*Significance at 10 percent level. All coefficients multiplied by 100 to facilitate interpretation. Coefficients are in percentage points. All models include controls for patient characteristics, hospital fixed effects, and year fixed effects which are allowed to vary by six MSA size categories. Heteroscedasticity-consistent standard errors are shown in parentheses. $N = 201,225$ in outpatient expenditure models, otherwise $N = 922,074$. AMI, acute myocardial infarction; CHF, congestive heart failure.

admitted with AMI on the weekend have approximately 0.55 percent lower inpatient expenditures in the year after their AMI, reflecting, in part, the lower volume of intensive procedures. Patients hospitalized on weekends also have 4.40 percent lower outpatient expenditures in the subsequent year. Weekend heart attack patients are more likely to die than their weekday peers in all time periods subsequent to the initial hospitalization; at 1 year, patients admitted on the weekend experience 0.38 percentage points higher mortality from AMI, which is statistically significant ($p < .01$). Expressed in relation to average AMI mortality in 1998 (36.3 percent), patients admitted on the weekend face a 1.0 percent increase in 1-year mortality. The effects of weekend admission on cardiac complications are mixed. While weekend admission has a positive and significant effect on readmission with CHF ($p < .01$), weekend patients are less likely ($p < .01$) to be readmitted with AMI in the year following hospitalization with the initial AMI.⁴

To address the issue of selection bias arising from delays in admission, the second rows of Tables 3 and 4 (specification 1b) present estimates that group weekend and Monday admissions together. According to Table 3, the effect of a Saturday/Sunday/Monday admission on the rapidity and probability of intensive treatment is smaller, but still statistically and economically significant. The effect of a Saturday/Sunday/Monday admission on 1-year mortality is reduced, but remains statistically significant. As in the weekend versus weekday models, Saturday/Sunday/Monday patients experience higher 1-year rates of readmission with CHF, but slightly lower 1-year AMI readmission rates. The effect on inpatient expenditures of Saturday/Sunday/Monday versus Tuesday–Friday admission is small and statistically insignificant.

I take an additional step to ensure that my results are not driven by selection bias attributable to patients experiencing less severe symptom onset on the weekend deferring admission until the following Monday. If emergency departments were less likely to diagnose patients with AMI who present with chest pain on the weekend, and these patients remain symptomatic and are hospitalized with AMI on Monday this could lead to bias. The direction of this bias is uncertain as, if deferred admission were harmless, estimates of the weekend effect based on admission date would *overstate* the true effect. However, if deferred admission were harmful, estimates of the weekend effect based upon the date of the AMI admission would *understate* the true effect. To explore this issue, I use a 20 percent sample of outpatient claims to identify patients who were treated in the ER with chest pain (but not hospitalized) in the 3 days before their AMI admission. Fewer than one in 700 AMI patients had such a visit before their hospitalization, and the incidence of such visits is

roughly constant across days of the week. I re-estimate all of the models in specification 1a, indexing each patient's admission date using the earlier of their actual AMI admission date or the date of any ER visit in the 3 days prior. This reassignment of index admission dates using prior ER visits with chest pain had no impact on any of the results.

Taken together, these results suggest that AMI patients admitted through the emergency room on the weekend receive lower quality health care, with the weekend effect on mortality not merely an artifact of selection. The 0.55 percent reduction in inpatient expenditures on the weekend amounts to approximately \$105 per weekend patient (based on the sample average 1-year inpatient expenditure of \$19,196). Using the point estimate of the effect of weekend admission on 1-year mortality, this suggests that the more intensive treatment provided to patients admitted through the ER during the week is cost effective assuming a value per year of life saved of approximately \$27,631 or greater.⁵

A number of researchers (see, e.g., Duan 1983; Manning 1998) have noted that log-transformed models may lead to significantly biased inference on the raw scale. In light of these concerns, I examine the sensitivity of my cost-effectiveness results using the two primary approaches for dealing with the retransformation problem—smearing estimates and generalized linear models (GLM). Using the homoscedastic smearing coefficient (Duan 1983), I find that weekend admission is associated with a \$102 reduction in 1-year inpatient expenditures. When I employ separate smearing coefficients for weekend and weekday patients to account for heteroscedasticity on the log scale, the effect of a weekend hospitalization on inpatient expenditures is significantly attenuated. Lastly, estimates from the expenditure model using GLM with a γ -distribution and a log link (Blough, Madden, and Hornbrook 1999) indicate that weekend patients have \$123 lower 1-year inpatient expenditures than their weekday peers. Under *all* of these alternative specifications, the treatment provided to weekend patients remains cost-ineffective.

To further explore the relationship between delayed provision of treatment and health outcomes, I estimate models where I examine separately the effects of Saturday and Sunday admissions. The coefficient estimates from these regressions (specification 1c) are reported in Tables 3 and 4. Saturday and Sunday patients experience similar declines in 0-day rates of all three cardiac procedures, suggesting the provision of immediate treatment is roughly similar on both weekend days. Patients admitted on Saturdays experience significantly larger declines in 1-day procedure rates. However, Saturday admissions make up most of the decline in 1-day procedure rates on days 2–7,

whereas Sunday admissions do not. Even at 365 days, Sunday patients experience slightly larger decreases in procedure rates. Patients admitted on Saturdays and Sundays experience similar declines in 1-year inpatient expenditures relative to weekday patients, while the decrease in 1-year outpatient expenditures is much larger for patients initially hospitalized on Sunday. The mortality effects in all periods subsequent to the AMI admission are significantly larger for Sunday admits. The effect of a Sunday admission on mortality is significant ($p < .001$) in all periods, rising from -0.37 percentage points at 30 days to -0.53 percentage points at 1 year. Only at 1 year do I observe statistically higher mortality for patients admitted on Saturday versus patients admitted during the week.

The final four rows of Tables 3 and 4 present estimates from models that allow the effects of weekend admission to vary with patients' ex ante health status. In Table 3, the effects of weekend admission on 0-day and 1-day procedure rates are negative for (base group) patients without a prior year's hospital admission, while the interaction effects between weekend admission and prior year admission are positive and smaller in absolute value than the base group effects. Conversely, in models of 2-7-day rates, the base group effects of weekend admission are positive, while the interaction effects are negative and roughly similar in magnitude to the interaction effects for 1-day rates. Interaction effects on 365-day cardiac procedure rates are insignificant. Taken together, these results indicate that sicker patients are less likely to experience delays in treatment, but when they do, those delays are similar in length to those experienced by healthy patients.

Specification 2 in Table 4 examines how the effects of weekend admission on expenditures and outcomes vary with observable patient health status at the time of initial admission with AMI. Weekend admission leads to a statistically significantly larger decrease in inpatient expenditures for sicker patients versus healthier patients. In contrast, the decrease in outpatient expenditures resulting from weekend admission is larger for observably healthier patients. Sicker patients experience similar increases in 1-year mortality and CHF as a result of weekend admission, although lower rates of 1-year readmission with AMI.

CONCLUSION

Is the positive correlation between weekend hospitalization and mortality the result of differences in patients' characteristics, hospitals' understaffing, or

another form of inappropriate treatment not easily remedied through regulation? Understanding whether a causal relationship exists, as well as its source, is necessary to address this particular case of a recurrent health policy question: should apparent shortfalls in quality be addressed with an input regulation or with reimbursement policy that rewards superior performance? Despite the critical importance of understanding how weekend hospitalization affects patient outcomes, existing research has largely failed to identify the mechanisms through which a weekend effect might occur.

I present three main findings. First, patients admitted on the weekend are significantly less likely to receive the primary intensive treatments associated with AMI within the first days of their admission. Weekend admission with AMI is associated primarily with delay and, to a lesser extent, with reductions in treatment intensity. Second, weekend admission with AMI leads to lower subsequent expenditures, but higher 1-year mortality and a higher rate of one major cardiac complication, readmission with CHF. By conventional standards of cost-effectiveness of medical treatment, weekend admission leads to inappropriate reductions in intensive medical care. Third, the effects of weekend admission do vary with patient health status. While weekend admission leads to delays in invasive treatments for all patients, sicker patients (defined by inpatient admission in the prior year) are less likely to experience such delays.

My analysis cannot identify the portion of the mortality effect of weekend admission caused by delay in the use of intensive procedures, the portion caused by reduced use of intensive procedures, and the portion caused by other changes in medical treatment or staffing. Empirically, weekend admission affects both delay and incidence of intensive procedures, and may also affect dimensions of treatment that I cannot observe. However, under certain assumptions, my analysis implies an upper bound on the mortality effect of particular intensive treatments. In this way, I can use previous clinical studies to validate my results.

As catheterization is primarily a diagnostic tool and immediate bypass is rarely performed, one could make the assumption that the entire weekend mortality effect results from the decline in the 0 day PTCA rate. In this case, the implied IV estimate of the effect of immediate PTCA on mortality is 0.322 (0.0038/0.0118). Patients who do not receive immediate PTCA as a result of weekend admission experience a 32.2 percentage point, or 87.5 percent (0.322/1998 base 1-year mortality of 0.363) increase in 1-year mortality. Alternatively, if we use the broader 1-day measure of immediate PTCA, the implied IV estimate falls to 0.250 (0.0038/0.0152), with weekend admits facing

a 67.9 percent increase in 1-year mortality.⁶ These upper bound estimates are somewhat higher, but of the same order of magnitude, as those obtained from randomized clinical trials examining the benefits of immediate or “primary” angioplasty. Grines et al. (1993), for example, found a 50 percent increased risk of mortality at 6 weeks among AMI patients treated with medical management versus those receiving primary angioplasty. Zijlstra et al. (1999) suggests long-term benefits from primary angioplasty, as patients receiving medical management experience a 50–100 percent increased risk of mortality and reinfarction both soon and long after their initial admission with AMI.

Other studies have presented mixed evidence regarding the mortality benefits associated with the intensive treatment of heart attack. In their landmark study using differential distance to technologically capable hospitals as an instrument for intensive treatment, McClellan, McNeil, and Newhouse (1994) find little effect of catheterization and revascularization (angioplasty) on the survival of the marginal patient. Another group of studies uses international variation in practice patterns to examine the efficacy of invasive cardiac procedures following AMI. Several of these studies, using randomized trial data from the United States and Canada (Rouleau et al. 1993; Tu et al. 1997), conclude that substantially lower rates of catheterization and angioplasty in Canada do not lead to higher mortality. However, other researchers have found evidence of both mortality (Kaul et al. 2004) and nonmortality (Mark et al. 1994) benefits associated with the more intensive management of AMI in the United States.

While I do not explicitly evaluate the merits of mandatory staffing legislation, or any other input regulation, my results suggest that such laws are unlikely to provide a comprehensive fix for the quality of care in American hospitals. The weekend effect is not the result of an inadequate stock of high-tech capital, as weekend patients are no less likely to be admitted to a hospital with a catheterization laboratory (Table 2). To further investigate the value of regulation, I examine whether the effect of weekend admission on intensive procedure use is the result of hospitals shutting down their catheterization laboratories over the weekend, or whether it is the result of an overall decrease in procedure rates across all hospitals. I identify “weekend shutdown” hospitals using the ratio of weekend to weekday cardiac procedures performed at each hospital during the entire study period. In models not reported in the tables, I find that patients admitted to hospitals which do not shut down on weekends experience larger decreases in immediate procedure rates. The fact that the weekend effect on procedures is *larger* at hospitals which do not typically shut down their catheterization laboratories suggests that neither

staffing regulation nor other rules mandating the availability of cardiac procedures on weekends will be effective in reducing the weekend effect.⁷

Instead, my results suggest that the weekend effect might be better addressed by reimbursing doctors and hospitals on the basis of the appropriateness of their treatment decisions. Although Medicare DRG payments are related to the volume of treatment a patient receives, the timeliness of these treatments is irrelevant. Given the costs of staffing and operating a catheterization lab, a hospital that sought to maximize the excess of revenues above costs might choose to limit such services on the weekend. To counter these incentives, hospitals could be paid a premium to provide such services on the weekend, perhaps offset by lower reimbursements for procedures provided during the week. An experimental reimbursement policy, whereby providers would receive a premium for performing emergency cardiac procedures on the weekend, would provide a relatively inexpensive way of confirming, or refuting, the results of this paper. The results presented in this paper must be qualified, as my analysis has focused on a single, albeit important, medical condition (AMI), and a single patient type (Medicare beneficiaries). My findings may, or may not, generalize to other conditions and patient populations. Further research is needed to identify other conditions where weekend treatment differences potentially lead to higher rates of adverse health outcomes.

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NOTES

1. For 1996 I use a 20 percent sample as I do not have access to Medicare claims from the prior year for the full sample. Observations for 1996 are upweighted by a factor of five in all regressions. For all years, I only have subsequent outpatient expenditures for the 20 percent sample. Sample sizes in log outpatient expenditure models are reported in Table 4.
2. The conditions used to construct the Charlson index are congestive heart failure, peripheral vascular disease, cerebrovascular disease, dementia, chronic pulmonary disease, rheumatological disease, ulcer, moderate liver disease, diabetes, dia-

betes w/complications, hemi- or paraplegia, renal disease, cancer, severe liver disease, metastatic cancer, and AIDS.

3. The effect of weekend admission on 365-day cumulative CATH (PTCA) rates is 1.01 (0.33) percentage points smaller at 365 days than at 7 days (e.g., $1.01 = -0.06 - (-3.47 + 2.40)$).
4. I examine the sensitivity of my results to the inclusion of various covariates. The effects on health outcomes and procedure rates are robust to the inclusion/exclusion of controls for infarction location, ex ante health status and patient demographic characteristics. Given the random nature of heart attack, comparison of weekend and weekday patients can be thought of as a crude regression discontinuity design—where fate assigns individuals a time when their AMI will occur. Similarity in observable characteristics across days of the week is consistent with this random assignment from nature.
5. $\$27,631 = (\text{average inpatient 1-year expenditures}) \$19,196 \times 0.0055/0.0038$. This calculation of cost-effectiveness is based upon inpatient expenditures alone. In models not reported in the paper I use the 20 percent sample to estimate the effect of weekend admission on total (inpatient+outpatient) 1-year expenditures with weekend care remaining cost-ineffective.
6. The results of specification 1c suggest that the difference in mortality rates for Saturday versus Sunday patients cannot be explained by differential use of immediate PTCA. However, models of 2–7-day procedure rates are consistent with a dose–response relationship, as patients admitted on Sundays experience more persistent delays in treatment, and worse health outcomes, than Saturday admits.
7. In models not reported in tables, I rule out two additional alternative mechanisms through which the weekend effect might operate—the availability of cardiologists and hospital volume. There is no evidence that more cardiologists (at three-digit zip code level) or a higher volume of patients mitigates the impact of weekend admission on treatments and outcomes.

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