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Do Unionized Registered Nurses Reduce AMI Mortality?*

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Abstract

Although recent research shows interesting relationships between hospital work organization and patient outcomes, the relationship between nurse unions and patient outcomes has not been explored. This study examines the relationship between the presence of a bargaining unit for registered nurses and the acute myocardial infarction (AMI) mortality rate for acute care hospitals in California. After accounting for patient and hospital characteristics, we find that hospitals with unionized RN workforces have 5.5 percent lower AMI mortality than do nonunion hospitals (−0.8 percentage point/14.6 percent average AMI mortality).

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Recent work by Seago and Ash (2001) shows an association between RN unionization and reduced AMI (heart attack) mortality after controlling for a host of observable hospital and patient characteristics. To make causal claims for the relationship, we need to consider the role of unobservable features of the hospital which may confound the measurement of the causal RN-union effect on patient outcomes.

The voice and image of unions for registered nurses (RNs) in California has changed in the past decade as RN unions have responded to changes in the healthcare environment as a threat both to their members (AONE 1994, Harris 1996, Sherer 1994) and to the well-being of their patients. On October 1, 1995, (California Nurse 1994, Moore 1995) the California Nurses Association (CNA), a union, severed its association with the American Nurses Association (ANA), a professional organization which has had an ambivalent orientation regarding the unionization of nurses. The activities that led to this disassociation produced tension in the California nursing community and among some nurse managers/educators and staff nurses.

During a recent lengthy and acrimonious dispute with Kaiser Permanente, the California Nurses Association (CNA) mounted a public relations campaign with the message that changes in the work of RNs and other healthcare workers jeopardizes patient safety and negatively influences patient outcomes (Sherer 1994). Kaiser, of course, denied those charges. It is not always clear whether changes in health care are associated with patient outcomes, but it is clear that hospitals/health systems and unions often engage in spirited rhetoric about what is best for patients with little objective evidence to support either view (Sherer 1994). Little systematic investigation has been done as to whether the presence or absence of a union for RNs is associated with patient outcomes. This study examines the relationship between the acute myocardial infarction (AMI) mortality rate and the presence of an RN union in the acute-care hospitals of California.

Organizational characteristics are known to influence patient outcomes. Showstack and others (Flood, Scott and Ewy 1984b, Flood, Scott and Ewy 1984a, Mitchell and Shortell 1997, Showstack, Kenneth and Garnick 1987) found that higher patient volumes in hospitals were associated with reduced patient mortality in both medical and surgical patients. The assumption is that patient volume is a proxy for experience or expertise of the hospital staff and care providers. In a quasi-experimental setting which controlled for hospital-patient-procedure selection, McClellan, McNeil and Newhouse (1994) and McClellan, Henson and Schmele (1994) found cardiac technology to be of limited value in reducing AMI mortality, but they found that the quality of care in the first 24 hours following the attack matters significantly. Aiken, Smith and Lake (1994), Aiken and Fagin (1997), and Aiken, Sloane, Lake, Sochalski and Weber (1999) found the organizational characteristics of magnet hospitals are

related to decreasing mortality rates and improving patient satisfaction with nursing care. Aiken's team discovered (Aiken et al. 1994) that good outcomes are not simply a matter of skill mix (RN hours/total nursing staff hours) but are also related to nurse autonomy and control of practice. These workplace practices have a beneficial impact on 30-day mortality and patient satisfaction for inpatients. More recently (Aiken et al. 1999) the researchers found that a richer RN/patient ratio is associated with reduced 30-day mortality and increased patient satisfaction. Kovner and Gergen (1998) found, after controlling for a wide variety of hospital characteristics, more care hours provided by RNs correlates with reduced instances of adverse events following surgery, including urinary tract infection and pneumonia. Employee satisfaction is related to positive patient outcomes (Aiken et al. 1994). As we discuss in the next section, unions have complex effects on many of these institutional factors.

Nurses Unions and Patient Health Outcomes

Unions may affect the quality of care, although the mechanism may be difficult to ascertain. In this section, we explore three venues by which unions may affect the quality of care. Under labor law, the unionization decision of registered nurses at a hospital designates the union as the exclusive bargaining agent for all registered nurses at that hospital and mandates collective bargaining between the nurses's union and hospital management. The institutional arrangement of collective bargaining mandates that management and the union negotiate in good faith on a defined set of conditions and terms of employment, including wage and benefits, overtime, work burden, and staffing levels.

Nurse quality

Unions often raise the wage that is paid to employees. Research indicates that unions are associated with slightly higher RN wages (Hirsch and Schumacher 1995, Wilson, Hamilton and Murphy 1990) although Hirsch and Schumacher (1998) found smaller union wage differences among healthcare workers than among similar workers in other industries and smaller differences among RNs than among lower skilled workers. The higher wage might improve the quality of the nursing staff by changing hiring practices. Facing the higher wage, employers may become more selective in hiring, for example, limiting hires of nurses with Associate's Degrees in Nursing in favor of nurses with BSNs or additional education.

Unions typically institute seniority systems which raise the wages and improve the working conditions of nurses with greater tenure. The higher wage for unionized nurses and the

seniority system may discourage quits, leading to a nursing work force with longer tenure at the particular hospital, greater experience with the procedures of the hospital, and better capacity to meet patient needs. The literature shows some evidence that higher wages reduce nurse turnover. Aiken (1989), Spetz (1996), and Seago and Ash (2001), demonstrate that nurse turnover is lower in hospitals with RN unions. On the other hand, nurses unions may offer job protection for nurses with poor work records, and this aspect of unionization would, presumably, be bad for patient outcomes.

Communication

Formal communication It is possible that the increased communication mandated by the collective bargaining process directly facilitates the transfer of information that may assist in reducing patient mortality. However, the particular contents of communication under collective bargaining, i.e., wage and working conditions, would not seem to be the materials of which reduced mortality is made, but we do not dismiss the possibility out of hand.

The collective bargaining relationship may facilitate other formal modes of communication. Labor law strongly precludes negotiation between management and nonunion employee organizations; the reason for this provision is to prevent the manipulation of such organizations into "company unions." In some cases, employers may be concerned that establishing employee organizations without the collective bargaining relationship would constitute a violation of labor law. Also, employees may have reservations about participating in nonunion employee organizations. So the collective bargaining relationship may facilitate other formal modes of communication between nurses, hospital management, and the medical staff, such as quality circles which Baskin and Shortell (1995), Shortell and Hull (1996), Shortell, Jones, Rademaker et al. (2000), Shortell, Zazzali, Burns et al. (2001), and Ferlie and Shortell (2001) identify as having a beneficial effect on patient outcomes.

Informal communication By offering protection from arbitrary dismissal or punishment and by implementing a grievance procedure, unions may reduce the intensely hierarchical character of the relationship between nurses and the medical staff. The protections offered by unionization may encourage nurses to "speak up" in ways that improve patient outcomes but might be considered insubordinate and, hence, career-jeopardizing without union protections (Gordon, Benner and Noddings 1996, Gordon 1998).

The increased tenure resulting from union negotiation of higher wages and a seniority system may also improve informal communication between RNs and the medical staff of

the hospital. Unions may affect the organization of nursing staff or the way nursing care is delivered in a fashion that facilitates RN-MD communication. This is the "voice" function of unions described by Freeman and Medoff (1984). This communication is constantly under stress from the hierarchical character of hospital workplace relations, but the physicians and surgeons of a hospital may come to trust the opinion of a particular nurse over time as the nurse's judgment is borne out by years of actual experience. By increasing the tenure of nurses, nurses' unions may increase the trust between nurses and MDs in a particular hospital setting.

On the other hand, unions may create an adversarial or rule-bound work environment that interferes with communication and care. There are indications in the literature that union activity is viewed as unprofessional (Breda 1997) and is associated with declining employee morale and job satisfaction (AONE 1994, Sherer 1994, Harper, Motwani and Subramanian 1994) and that the normal state is for managers and unions to be adversarial (Breda 1997, Flarey, Yoder and Barabas 1992). Managers and administrators typically view union activity negatively, and they often use strategies to prevent unionization. Most administrators assume that managing in a union environment makes their jobs more difficult (Harper et al. 1994, Flarey et al. 1992).

Staffing levels

The work burden is one of the mandated topics of collective bargaining, and staffing levels have been a particular focus of activism by nursing unions in California. Unions may improve the quality of care by negotiating increased staffing levels which, according to Kovner and Gergen (1998), improve patient outcomes. On the other hand, unions may raise wages to an extent that the employer slows hiring or stops hiring nurses and staffing is adversely affected. There is some indication that unionization of hospitals is related to increasing costs (Wilson et al. 1990).

In summary, there is a substantial literature on patient outcomes related to hospital organizational variables, but there are unanswered questions. With the current data, we can explore the association between the presence of RN unions and patient outcomes. It is possible that the wage is the important factor in attracting and retaining high quality nurses and that unions only function is to win higher wages. In this case, the wage, not the union, would be the causative factor, and the union would be only an instrumental factor (Hirsch and Schumacher 1995, Hirsch and Schumacher 1998, Spetz 1996). The situation is analogous for increased staffing as a result of collective bargaining. By controlling directly for wages and staffing, we can examine to what extent unions have an effect on care over and

above their direct effect on wages and staffing. The current data will permit only limited resolution of the mechanisms by which unions affect care. Furthermore, we overlook ways in which unions cause organizational change in hospitals but the change occurs in areas that do not affect patient care.

Confounding Factors

Many complex factors determine the survival of patients suffering from AMI. To make our analysis useful and our estimates of the RN union effect unbiased, we need to consider to what extent these various factors are correlated with RN unionization.

Patient characteristics and hospital selection

Patient characteristics such as sex, age, race, insurance category, risk characteristics and illness severity are associated with patient outcomes such as mortality and other adverse events (Aiken et al. 1999). Because unionized hospitals are more likely to be located in urban areas, the AMI patients they treat are likely to differ systematically from those AMI patients treated in nonunion hospitals. We do not directly examine the selection of patients to union or nonunion hospitals, but we surmise that patients treated in unionized hospitals have average characteristics that in some dimensions increase and in other dimensions decrease their probability of AMI mortality. For example, patients treated in unionized hospitals are more likely to be African American (an increased risk factor for AMI mortality) but are less likely to have had the substantial delay in treatment associate with transit time to rural hospitals (a decreased risk factor for AMI mortality).

Our strategy for addressing the nonrandom matching of patients and hospitals is to use mortality data which have been carefully risk-adjusted for patient age, gender, type of AMI, and chronic illnesses. By using both adjusted and raw data, we can directly test whether unionized hospitals indeed receive a nonrandom patient load and whether the patients are, on average, positively or negatively selected. Furthermore, because urban location appears to be such an obvious basis of patient selectivity, we directly include controls for the location of the hospital in a rural or urban area.

Hospital characteristics

Mitchell argues that mortality rates and complications seem to be more closely related to patient variables, while adverse events may be a more sensitive marker of organizational

variables (Mitchell and Shortell 1997).

Hospital Ownership and Control Arrow (1963) suggests that the not-for-profit hospital may have advantages over the for-profit form in in meeting patient need under conditions of uncertainty in diagnosis and treatment. It is also possible that the not-for-profit organization is an easier environment for unionization, for example, because management lacks strong incentives to resist it. If not-for-profit organization is associated both with improved partient outcomes and with unionization, then an analysis which failed to incorporate hospital ownership and control would spuriously associate these characteristics. We control directly for the ownership and contol of hospitals by including indicator variables for public and for-profit hospitals, i.e., not-for-profit hospitals are the excluded category.

Expensive technologies Mitchell and others have consistently found high technology in hospitals to be related to lower mortality (Mitchell and Shortell 1997, Hartz, Krakauer and Kuhn 1989, Manheim, Feinglass, Shortell and Hughes 1992), although in a quasi-experimental setting, McClellan et al. (McClellan, McNeil and Newhouse 1994, McClellan, Henson and Schmele 1994) found cardiac technology to be of limited value in reducing AMI mortality. Sophisticated cardiac technology, i.e., the high-value capital assets of a hospital, may be attractive to unions as the basis of extracting union rents; so unions may tend to form at these hospitals.

If expensive technologies are both attractive to unions and useful in reducing AMI mortality, a naïve analysis would spuriously find unions associated with reduced AMI mortality. We have a three-fold strategy for addressing this issue. First, we directly control for the presence of expensive cardiac technologies using several measures of these technologies described later. Second, we test the effect of non-healthcare bargaining units, e.g., unionized engineers or food service workers, at the hospital on the AMI mortality outcome. Unions of nonhealthcare workers would have the capacity to reap the same rents created by the presence of these expensive technologies but, by assumption, should have no effect on AMI mortality. If our analysis finds an effect of nonhealthcare unions on AMI mortality, then the union effect is likely spurious.

Lastly, in addition to union status in the outcome-measurement period (1991-1993), we also observe whether hospitals not unionized as of 1991–1993 have subsequently unionized over the period 1993 to 1998. Hospitals that subsequently unionize are comparable in their characteristics to hospitals that have already unionized. If our analysis shows an effect of subsequent unionization on the earlier AMI mortality outcome, then the association of the effect with unionization is spurious.

Other sources of capital rent With some diseases, particularly with surgical and technical procedures, higher volumes of patients treated at the hospital has been shown to improve patient outcomes (Flood et al. 1984b, Luft and Romano 1993, Showstack, Kenneth, Garnick, Luft, Schaffarzick and Fowles 1987). There is also some indication that hospitals larger than 100 beds have better patient outcomes (Luft and Romano 1993, Grumbach, Anderson, Luft, Roos and Brook 1995, Phillips, Luft and Ritchie 1995). Again, these larger, high-volume hospitals may be a source of rents for organized labor at the hospital, and the same spurious correlation might appear. Our strategy is similar to our strategy for addressing the spurious correlation with high-tech apparatus: we directly control for hospital size and volume of activity, and we use nonhealthcare unions and future unionizers as specification tests.

Human capital rent As unions may extract rents for the members from the capital stock of a hospital, so too can they extract rents from the human capital stock of other persons working at the hospital. So RN unions might be attracted to hospitals that have strong prospects of yielding these human-capital rents, for example, by attaching to teaching hospital or to hospitals with many or highly-skilled MDs. There are studies that indicate increased RN hours, increased MDs, and increased total staff hours have a positive impact on patient outcomes (Shortell, Zimmerman, Rousseau et al. 1994, Zimmerman, Knaus, Draper et al. 1991, Zimmerman, Rousseau, Duffy et al. 1994). Again we control directly for human capital characteristics of the medical and care staff of the hospital, and we use nonhealthcare unions and future unionizers to test for spurious association.

Adversarial industrial relations Adversarial industrial relations may be the cause of unionization rather than an effect. As noted above, employee satisfaction is related to positive patient outcomes (Aiken et al. 1994). If adversarial industrial relations and diminished employee satisfaction cause both unionization and worsened patient outcomes, then our analysis might spuriously associate the union with worsened patient outcomes. We have no direct measure of already existing adversarial industrial relations. So the measured effect of RN unions on patient outcomes will be inclusive of already existing adversarial industrial relations and, hence, biased towards finding that unions are bad for patient outcomes. To the extent that the presence of nonhealthcare unions or future unionizers are markers for adversarial industrial relations at the hospital, we will address this issue.

To explore further the adversarial workplace relations, we note that relations are frequently most adversarial in the period immediately after unionization (and often calming with the settlement of the first contract). In addition to union status in the outcome-measurement period (1991-1993), we also observe the date of unionization. So among union-

ized hospitals we can distinguish between hospitals with recently formed (after 1987) and those with longstanding unions (unionized before or by 1987). We divide the unionized hospitals into hospitals with recently formed (after 1987) and those with longstanding unions (unionized before or by 1987). We run additional specifications of the model with indicator variables entered separately for hospitals in these two categories. The test of the effect of adversarial relations in a unionized setting is carried out by examining the coefficients on both indicator variables.[1]

In summary, it is clear that mortality is significantly predicted by patient characteristics, as well as hospital characteristics and environmental factors. This study seeks to determine for acute care hospitals in California if, controlling for these factors, there is a relationship between the acute myocardial infarction (AMI) mortality rate and the presence of a bargaining unit for registered nurses. Because geographic regions in California are very different from each other and the presence or absence of unions within geographic areas may influence each other, we are controlling for regional (health service area) characteristics as well as hospital and patient characteristics.

Methods

The outcome measure is risk-adjusted AMI mortality, a measure described in more detail below. Because hospital mortality can be a function of many factors other than whether hospital nurses are unionized, we estimate multivariate regression models to control for these other characteristics. After we explored descriptive statistics stratified by union status, we used a multivariate regression model to test the significance of the variable of interest in the presence of covariates.

Sample

The target population was all acute care hospitals in California.[2] California was selected because (1) it has risk-adjustment mortality data for AMI, a common diagnostic related group; (2) there is variation among hospitals regarding union status; and (3) the investigators

[1]The difference suggests that first-differenced approaches to the effect of unionization, such as that employed by Hoxby (1996) in her study of teacher unionzation may mistakenly capture adversarial relations in the unionization process rather than the effect of the union itself.

[2]A power analysis indicated that a multiple linear regression model including 16 covariates and one predictor (union/nonunion) with a sample size of 236 would have 80 percent power to detect a squared multiple correlation (R^2) of 0.08 at alpha ¡ 0.05. So our sample should be adequate for inference.

are familiar with the health care system of the state and associated databases. We made a determined effort to contact every acute care hospital in California, although non-response is potentially nonrandom. Inclusion criteria were all (non-Federal) acute care hospitals in the state of California that reported discharge data in the 1991–93 OSHPD Hospital Disclosure database. Exclusion criteria were children's hospitals, psychiatric hospitals, rehabilitation hospitals, and long term care hospitals because there are too few specialty hospitals in the state to make an adequate comparison. All the Kaiser Foundation hospitals (24 hospitals whose RNs are represented by unions) were excluded because they do not report Hospital Disclosure (financial and accounting) data to OSHPD.[3]

Variables and the database

The dependent variable is the risk-adjusted 30-day AMI mortality rate for each hospital which was developed in the California Hospital Outcomes Project (CHOP) for years 1991–93 (Zach, Romano and Luft 1997, Romano, Luft, Rainwater and Zach 1997, Luft and Romano 1997, OSHPD 1997).[4] The risk-adjustment model used in this study accounted for the following patient characteristics: age; sex; type of heart attack; and the presence of chronic diseases (specifically, congestive heart failure, central nervous system disease, renal failure, diabetes, malignant neoplasm, hypertension, previous coronary artery bypass graft surgery, thyroid disease) (Luft and Romano 1997, Romano et al. 1997, Luft and Romano 1997) that were present on admission to the hospital. Outcome data from the 3 study years were pooled to reduce the degree of measurement error in the outcome variable, although unpooled variants of the model were examined as well. In some specifications, we use the raw AMI mortality rate for each hospital as an alternative dependent variable in order to explore the effect of the risk-adjustment. By using the risk-adjusted outcome data and by weighting the analysis by caseload, our analysis is similar to a patient-level analysis clustered by hospital.

The independent variable of interest in this study is RN union status. RN union status was found using web pages such as the National Labor Relations Board (NLRB) web page, from various bargaining unions, or by calling the hospitals directly. The first attempt to ascertain union status was by exploring the hospital or public web pages. We then called the RN union offices, or calling the hospital personnel office or nursing office between December 1998 and April 1999. The telephone survey questions included the presence or absence of

[3]We were able to estimate the bivariate regression with the Kaiser hospitals in the sample, and the results were unchanged.

[4]Because the outcome variable is a rate between zero and 100 percent, a logistic regression may be more appropriate. On the other hand, the simple linear regression offers the benefit of easy interpretation of the coefficients. We estimated all models with the logistic specification, i.e., $\ln \frac{MR}{1-MR}$ as the dependent variable, where MR is the mortality rate. The signs and significance of the results were identical.

an RN union, name of the union, presence or absence of unions for other employees in the hospital, the names of all the other unions, and when all unions were formed.

Additional variables were obtained from the California Office of Statewide Health Planning and Development (OSHPD) Hospital Disclosure Report database. These data are collected annually from all hospitals in the state and include information about service provision, finances, and resource utilization. The database also includes information about capital acquisition, labor staffing, and the provision of medical care in each revenue unit of the hospital. The reported and derived variables include: the number of hospital beds and discharges from nursing units that care for AMI-related patients, specifically medical/surgical intensive care units, coronary care units, other intensive care units and acute medical/surgical nursing units; a technology index derived from a principal components factor analysis of all the cardiac specialty services; availability of providers, specifically, RN hours, total staff hours, and the number of active medical staff with clinical specialties in any cardiovascular specialty, internal medicine, or therapeutic radiology; and the average wage of registered nurses adjusted for the local area cost of living (Table 1).

The average hourly wage of RNs adjusted for the local cost of living was computed as follows. The average hourly wage was computed for RNs in hospital units associated with AMI: medical/surgical intensive care; coronary care; other intensive care; and medical/surgical acute units. The hospital average wage was regressed on a set of dummy variables for each Health Service Area (HSA),[5] and then residuals from this regression were used as the wage index variable in the final models. The measure indicates how much the hospital's wage for RNs in AMI care exceeds that of other hospitals in the HSA. The within-HSA wage residual was used as the wage index to control for price differences across HSAs, and is, hence, analogous to a real wage. Alternative specifications used the hospital wage alone or the hospital wage relative to per capita county income and found very similar results.

The cardiac technology index was derived from a principal components analysis of all the cardiac specialty services (coronary intensive care unit, cardiac catheterization, echocardiology lab, heart surgery services, open heart surgery services, mobile cardiac care, and cardiac clinic). It replaces all the cardiac service variables. The advantage of using the cardiac technology index over individual dummy variables is a reduction in dimensionality that increases the adjusted R-squared. The key results with the cardiac specialty services included separately are very similar to those with the index and are available from the authors. See

[5]The federal government designates HSAs for planning and resource analysis of hospitals and other health services. There are 205 HSAs in the US and they are constructed of either entire states or aggregates of counties (Makuc, 1991 340) (Seago and Ash 2001, Grumbach, Ash, Seago, Coffman and Spetz Forthcoming 2001). California has 58 counties and 14 Federally designated HSAs.

Table 1 for variable definitions and data sources.

Based on the conceptual grouping of variables, several regression specifications were estimated. All the models include the RN union variable as a predictor because it is the variable of interest. The first specification examines the bivariate relationship between mortality and union status. The second specification includes all of the confounding factors discussed above, including volume (hospital size classes by number of beds and AMI-related hospital discharges), staff hours per AMI discharge, the number of MDs per AMI discharge, an indicator teaching hospital, the cardiac technology index, and hospital-location variables. Specifications 3, 4, and 5 add the local area wage premium and RN hours per AMI discharge, variables that are themselves likely affected by unionization. Hence, specification 5 examines the effect of unionization over and above its effect on both of these specific outcomes of collective bargaining.

Results

Of the 385 eligible acute care hospitals during this time period, we obtained complete data on 344. Summary statistics for RN-union and RN-nonunion hospitals are reported in Table 2. As of 1993, 27 percent percent of California hospitals had RN unions. Smaller and nonmetropolitan hospitals are less likely to be unionized. Union hospitals have more AMI-related discharges, more MDs per AMI-related discharge, more complex technology, specifically heart surgery services and open heart surgery. Union hospitals also pay higher RN hourly wages, both in absolute terms and relative to nonunion hospitals in the same market, and unionized hospitals apply more RN hours to the average AMI case. We report 5 specifications for the regression analysis in Table 3.

Union results The most interesting finding is that the presence of an RN union significantly predicts lower risk-adjusted AMI mortality in all specifications. In specification 1 of Table 3, the bivariate regression with no controls for confounding factors, the presence of a nurse's union is associated with a 1.0 percentage point reduction in AMI mortality. The average AMI mortality in our sample is 14.6 percent; so the presence of a nurse's union is associated a 6.8 percent reduction in mortality. The adjusted R-squared for this regression is 0.02.

Specification 2, the estimation with full controls, is our preferred specification for estimating the effect of RN unions on AMI mortality. When we introduce the full set of controls for confounding factors in specification 2, the adjusted R-squared increases to 0.10, and the presence of an RN union is associated with a 0.84 percentage point, or 5.8 percent, reduction

in AMI mortality. That is, the full set of independent variables reduces the estimated coefficient on RN union by approximately one-fifth, and the coefficient remains highly significant with a standard error of 0.35 and a p-value of 0.02.

In specifications 3, 4, and 5, we introduce, respectively and then together, RN hours per AMI discharge and the wage (expressed as the within-HSA premium); the logic of their inclusion is that they are likely to be affected by RN unionization. Neither variable has a statistically significant coefficient, although both terms have negative point estimates, suggesting that a higher wage and more RN hours are both associated with decreased AMI mortality. The adjusted R-squared does not increase from that of the preferred specification for any of these variants. At the point estimates for wage, the implication is that paying a ten-dollar per hour premium above the HSA average would reduce AMI mortality by about 0.7 percentage point. At the point estimate for RN hours, the implication is that increasing RN hours by about 5 hours (16 percent) per AMI-related discharge would reduce AMI mortality by about 0.1 percentage point.

Furthermore, neither the inclusion of RN wage nor that of RN hours substantially affects the coefficient on RN union, which falls only slightly in magnitude, from -0.84 to -0.76, with their joint inclusion in specification 5. The estimate of the union effect on AMI mortality through changes in nurse quality and communication, over and above the effect of unionization through its effect on wages and hours, is a 0.8 percentage point, or 5.8 percent, reduction in mortality. Thus, the union has an effect beyond simple increases in wage and simple increases in RN hours.

Adversarial industrial relations To test the hypothesis that adversarial industrial relations around the unionization process adversely affect patient outcomes—distinction to the presence of the unions themselves—we examine the difference in patient outcomes between hospitals that have recently undergone unionization and those with longstanding unions. Of hospitals with unionized RNs in 1993, Table 2 shows that 9 percent had unionized since 1987.

The results of these specifications are reported in Table 5. Recently unionized (between 1988 and 1993) hospitals have a small, positive but insignificant point estimate in the AMI mortality estimation, indicating that they are essentially indistinguishable from nonunionized hospitals. Hospitals with longstanding RN unions (unionized before or by 1987) have slightly lower mortality than do all unionized hospitals; the point estimate increases in magnitude from -0.84 for all RN-unionized hospitals to -0.96 for longstanding RN-unionized hospitals.

These results suggest possible explanations which warrant further investigation. First, the very process of unionization may generate adversarial industrial relations which prevent

the immediate realization of the care-quality improvements associated with the presence of unions. Second, adversarial industrial relations may both adversely affect patient outcomes and induce unionization. Or, third, the benefits of unionization may only arrive after reorganization of the work process, i.e., after the implementation of selection, communication, and retention policies and processess described in the section on theory. The current data do not permit discrimination among these theories, but the difference in care outcomes between recently and longer-unionized hospitals is intriguing.

Other results Few other variables were significant at conventional levels in the specifications reported in Table 3. The volume measure, i.e., having more cardiac related hospital discharges, was also related to lower mortality rates, a finding consistent with past studies, e.g., Showstack, Kenneth and Garnick (1987). Having more cardiac services as measured by the cardiac technology index (cardiac catheterization, a cardiology clinic, coronary intensive care unit, mobile cardiac services, and heart surgery services) was significantly related to lower mortality rates. The cardiac technology index is significant in the expected direction (although in the specification with all of the individual specialty services entered separately, no one service was significantly different from zero). A one standard-deviation change in the index is associated with a reduction in AMI mortality of about 0.9 percentage points. Larger hospitals (those with more than 300 beds) and smaller hospitals (those with fewer than 100 beds) had point estimates of lower AMI mortality than did mid-range hospitals, although this result was not statistically significant. For-profit hospitals had a higher mortality rate than did the other ownership forms, and the result was significantly different from zero in some specifications. Additional MDs per AMI discharge was associated with reduced AMI mortality, although again the effect was not statistically significant.

Specification-Test and Selection Results

Patient characteristics and risk adjustment We first examine the effect of risk adjustment by using raw rather than adjusted mortality as the dependent variable. In column 1 of Table 4, we use the same preferred set of regressors as in specification 2 of Table 3, but we use unadjusted, rather than risk-adjusted, AMI mortality at the dependent variable. The estimated coefficient on RN union with the unadjusted mortality data is -0.15, and the estimate is not statistically significant. This result means that unionized hospitals receive a substantially more at-risk AMI caseload than do non-unionized hospitals. Failure to risk adjust would have meant incorrectly attributing the higher mortality at unionized hospitals to their unionized status rather than to their more difficult caseload.

Non-healthcare unions As we indicated in our discussion of methods, we are worried about the interpretation of causality in our analysis. Our proposed remedy is to examine the "effect" of non-healthcare unions at the hospital on AMI mortality. To the extent that non-healthcare unions are associated with reduced AMI mortality, we may be concerned that the relationship between unionization and mortality is spurious, caused by some third factor, for example, high-tech capital, which both reduces mortality and attracts unionization.

The specification test is to include the presence of a nonhealth care union and to exclude the presence of an RN union, with all the other covariates from specification 2, i.e., the preferred specification. The hypothesis that RN unions indeed affect AMI mortality suggests that this equation is misspecified, but the relatively high correlation of 0.70 between RN and non-healthcare unions will attribute some of the RN union effect to the non-healthcare union. The results are reported in column 2 of Table 4. When the non-healthcare union indicator is included instead of the RN union indicator, the coefficient is an insignificant -0.4, with standard error equal to 0.4. When we exclude hospitals which do have RN unions and compare mortality at hospital with and without nonhealthcare unions, we find that unionized hospitals have lower mortality—a point estimate of -0.72—but the estimate is far from significant with a standard error of 1.23. On balance, These results suggest that RN unions, rather than hospital unionizability, are associated with the reduction in AMI mortality.

An alternative test is an attempt to control for factors that may cause both union selection and affect patient outcomes. This test is to include both the RN union indicator and the non-healthcare union indicator. In this case, reported in column 3 of Table 4, the coefficient on non-healthcare union is a positive although insignificant 0.4, and the coefficient on RN union is significant and actually increases slightly in magnitude to -0.89. Again, this specification test provides evidence for the hypothesis that RN unions indeed have a causal association with reduced AMI mortality. The results of the specification tests are essentially unchanged when we repeat the exercise with the additional inclusion of an indicator for unionized Licensed Vocational Nurses and Nurses' Aides.

Our next test is to apply the Heckman two-step selection correction to control for nonrandom selection of hospital union status. Although no exclusion restriction is computationally necessary to estimate this method, we propose presence of a nonhealthcare union as an appropriate excluded variable.[6] In the first stage, we estimate a probit for the presence of an

[6]We have a rare instance in which we have an appropriate exclusion restriction for the first-stage probit of the Heckman selection correction which is not an appropriate instrument for 2SLS. The presence of non-healthcare unions is inappropriate for 2SLS; we argue that the presence of nonhealthcare union is correlated with the unobservable features of a hospital that make it both high quality with respect to patient outcomes and a likely candidate for unionization.

RN union at the hospital, including all of the variables in our preferred specification as well as the presence of a nonhealthcare union. The only significant variables in the first stage are the presence of a nonhealthcare union with a large coefficient and a p-value below 0.0001 and public ownership of the hospital (with a p-value of 0.04).[7] The low levels of significance of other variables in this estimation suggest that union and nonunion hospitals do not differ substantially in the observables. In the second stage OLS estimation, we include the inverse Mill's ratio computed from the probit results to capture features of the hospital which are correlated both with propensity to unionize and potentially with patient outcomes. The results are reported in column 4 of Table 4. The coefficient on the inverse Mill's ratio is small and far from significance suggestiong a small role for nonrandom hospital-union selection correlated with patient outcomes. The coefficient on RN union remains negative and stable at -0.74, but it no longer significant, having a standard error of 0.61. Increases in standard error are endemic in two-stage methods such as the Heckman selection correction, and so this result on balance adds weak confirming evidence for the relationship between the presence of an RN union and reduced AMI mortality.

Future unionizers An alternative approach to these specification and selection corrections is to examine hospitals that were not unionized during the study period but subsequently did unionize. As Table 2 shows, about 5 percent of hospitals which did not have RN unions in 1993 unionized over the 1993–1998 period. Again, we use the specification test approach, i.e., including future unionizers instead of the currently unionized and excluding the currently unionized from the analysis.

When future unionizers are included instead of the currently unionized, the estimated effect, -1.0 percentage points, is negative and of the same magnitude as that on currently unionized, but the coefficient is quite far from significance at conventional levels. The standard error is 0.86 and the p-value is 0.23.

The high and negative, albeit insignificant, coefficient on future unionizers does give pause and leaves concern that union-hospital selection contaminates the causal interpretation of the results. One possible interpretation is that there is a variety of practice in future unionizers and that some achieve the reduced mortality rates of which they are capable before unionization. The high, negative, and significant coefficient on the currently unionized suggests that RN unions help to implement the potentially improved practice. So the presence of a current RN union is associated with a reduction in mortality of about 0.8 percentage point even after controlling for hospital characteristics and our best effort at controlling for union-hospital selection.

[7] Other results of the probit are available from the authors.

Discussion and Implications

Potential Methodological Limitations

A key issue for this study is the question of causation. It is possible, for example, that union status is highly correlated with features of the geographic area that influence patient outcomes, regardless of whether the hospital has an RN union. For example, union density, that is the number of union hospitals divided by total hospitals in an area, in the San Francisco Bay Area is much higher than it is in other parts of California. In the SF Bay Area it ranges from 65 to 95 percent while in other parts of California it ranges from 10 to 60 percent. To test the robustness of the model within regions, we estimated the model with fixed effects for each of California's 14 Health Service Areas to examine the effect of within-HSA variation in union status on within-HSA differences in the outcome. This allows for comparing union/nonunion hospitals within each HSA. Unfortunately, using the HSA indicator reduces the variation in the explanatory variables, most importantly, in RN union status, so the effects may become harder to observe. Even in the HSA fixed effects models, the coefficient remains relatively stable at and negative at -0.45 although it ceases to be statistically significant (p-value of 0.43). There is no increase in the adjusted R-squared for the HSA fixed effect regression.

The mechanism by which unions affect patient outcomes is another area for further explanation. The type of data employed in this study cannot shed light on the particular mechanisms by which unions affect patient outcomes. However, we do observe that the relationship persists even when we control for the wage and hour changes which unions may cause.

Another concern is that hospitals with very small numbers of AMI cases did not generate a reliable risk adjusted mortality rate. In order to test this concern, correlation matrices were generated across years of risk-adjusted mortality rates stratified by the number of cases. The median hospital had an average of 51 or more AMI cases per year. The cross-year correlation of risk-adjusted mortality for hospitals that had 51 or more cases per year was statistically significant (Table 7, Panel A). The cross-year correlation of risk-adjusted mortality of hospitals that had fewer than 51 cases per year was not statistically significant (Table 7, Panel B). This suggests that hospitals with lower numbers of cases produce less reliable measures of risk-adjusted mortality, that is the risk-adjusted mortality in one year is a poor estimate of the risk-adjusted mortality in another year. Thus, in all models, the data were weighted by the number of AMI cases because hospitals with higher numbers of cases have lower variance estimates of risk-adjusted mortality. This approach is analogous

to treating the data as if it were patient-based rather than hospital based with hospital-wide values assigned to all patients at the hospital. Although the risk-adjusted mortality numbers were unbiased for all hospitals, the lower variance estimates are closer to a true measure of hospital quality. Specifically, the risk-adjustment model is a more accurate measure of quality for hospitals with more cases.

Lastly, there is the argument that already existing employee discontent will both affect patient outcomes (AONE 1994, Sherer 1994, Aiken et al. 1994, Wilson et al. 1990) and spur unionization. Although, this argument has merit, we were unable to obtain data specifically related to union activity, so we were unable to address this problem. This bias would tend to bias the results towards a finding that unionization is bad for patient outcomes, which suggests that our result in the opposite direction may be underestimated. Our results on longstanding versus recent RN unions suggest that the hypothesis of already existing adversarial relations may apply.

Conclusion and New Directions

As of 1998, thirty-five percent of hospitals in California have RN unions. The significant finding in this study is that hospitals in California with RN unions have 6 percent lower mortality rates for AMI after accounting for patient and hospital characteristics. Findings from the study do not rule out the nonrandom selection hypothesis regarding the presence of unions, and this limits the causal interpretation of these results. Our analyses, which examine the association between unions and patient outcomes, may be confounded by several factors, all of which involve nonrandom selection of the union status of hospitals. The direction of bias might go in either direction. For example, unions may be easier to establish in high-quality hospitals, perhaps because these hospitals have the capacity to pay premium wages or otherwise to offer more attractive working conditions. In this case, unions do not "cause" good outcomes but rather choose hospitals with good outcomes. On the other hand, the reverse may be true: unions may thrive in hospitals that have poor outcomes, possibly because morale is low. In the latter case, a correlative analysis would associate unions with bad outcomes, although again the relationship is not causal. In either case, the observed correlation is a biased estimate of a causal relationship, and we have no a priori means of determining the direction of bias. Although the use of instrumental variables techniques (also know as two-stage least squares or 2SLS) would be an econometric solution to this problem, no suitable instrumental variables were available (J. Spetz, personal communication, June 1, 2001). Therefore, we have relied on specification tests of and the Heckman selection correction for nonrandom union selection.

This study demonstrates that there is a positive relationship between patient outcomes and RN unions. Accounting for union status is important when examining organizational factors that influence patient outcomes. Further work exploring the relationship between unions and patient outcomes must be done, including a more exhaustive analysis of the causality in the relationship and more detailed examination of the mechanisms at work. Our findings indicate that there is something beyond wages and number of hours in union hospitals that predict better outcomes; however, type of nursing care delivery system, skill mix, and other organizational factors were not randomized and may be associated with unions. Perhaps having an RN union promotes stability in staff, autonomy, collaboration with MDs and control of practice decisions that have been described as having a positive influence on patient outcomes.

Our findings have implications for the operation of hospitals, organizations whose governance and goals are of a mixed public-private character. So the implications lie between somewhere between recommendations for public policy and suggestions for private, human-resource management.

Hospitals with nonunionized RNs may want to consider facilitation of or, at least, employer neutrality in RN efforts to unionize. Hospitals both with and without unionized RNs may want to identify which of the workplace practices in unionized hospitals affect patient outcomes and encourage those practices which improve patient outcomes. Recall that the net favorable RN-union effect which we identified is potentially an amalgam of outcome-improving practices (retention, quality-selection, and communication) and outcome-harming practices (adversarial workplace relations and personnel budget-squeeze). So the net favorable RN-union effect may not represent the theoretical best-practice.

We will next direct our research to determine which clinical practices by nurses save lives and which—not whether—mechanisms of the unionized hospital workplace facilitate such life-saving practices.

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Table 1: Variable Definitions and Data Sources

Variable	Definition (source)
Dependent Variables	
Risk-Adjusted AMI Mortality rate	30-day death rate adjusted for severity of illness, age, and gender of the AMI patient population (CHOP)
Raw AMI Mortality rate	30-day raw death rate of the AMI patient population (CHOP)
Independent Variables	
RN Union, 1993	Union for RNs present as of 1993 (PDC)
RN Union long	Union for RNs present as of 1987 (PDC)
RN Union new	Union for RNs formed 1988–1993 (PDC)
RN Union coming	Union for RNs formed 1994–1998 (PDC)
Size by staffed beds	Number of staffed beds, three categories: 100 or fewer; 101-300 (reference); 300 or more (AHDR)
Control or ownership	Not-for-profit (reference); For-profit; Public (AHDR)
AMI-related discharges	Number of hospital discharges from AMI-related cost centers (AHDR)
Cardiac Technology Index	Constructed by principal components analysis of all the cardiac services; see text. (AHDR)
RN hours per AMI-related discharge	Number of RN hours per AMI-related discharge (AHDR)
Total hours per AMI-related discharge	Number of total staff hours per AMI-related discharge, excluding RNs (AHDR)
MDs per AMI-related discharge	Number of MDs per AMI-related discharge (AHDR)
Within-HSA RN Wage Premium	RN wage relative to RN wages in other hospitals in the same Health Services Area (AHDR)
Rural Hospital	One of 74 hospitals designated as rural acute care hospitals in California (OSHPD, special correspondence)
Large Metropolitan	Central counties of metropolitan areas with 1 million or more population (ARF)
Medium Metropolitan	Other metropolitan counties (ARF)
Nonmetropolitan	County not in a metropolitan area (ARF)

Sources
PDC: Primary data collection by authors.
CHOP: California Hospital Outcomes Project (CHOP) from the Office of Statewide Health Planning and Development (OSHPD).
AHDR: OSHPD Annual Hospital Disclosure Report.
ARF: Area Resource File, U.S. Bureau of the Health Professions.

Table 2: Summary statistics by 1993 union status

	No RN Union $N = 250$		RN Union $N = 94$	
	Mean	Standard Deviation	Mean	Standard Deviation
Dependent Variables				
Risk-adjusted AMI Mortality Rate	14.96	3.14	13.95	2.63
Raw AMI Mortality Rate	15.13	3.45	14.53	3.04
RN union status				
RN Union new (1988-1993)	0%		9%	
RN Union long (before 1988)	0%		91%	
RN Union coming (1994-1998)	5%		0%	
Other hospital characteristics				
Annual discharges (100s)	58.5	39.9	72.9	52.7
RN hours per discharge	28.8	10.7	28.4	7.5
Total other hours per discharge	26.5	15.6	28.2	12.1
Staff MDs per discharge	0.043	0.117	0.113	0.246
Board Certified MDs per discharge	0.009	0.011	0.010	0.010
Hourly wage of RNs in AMI units	$24.63	$3.04	$28.52	$4.28
Within-HSA RN wage premium	$0.21	$2.71	$1.59	$2.34
Hospital Size by Staffed Beds				
100 or fewer beds	11%		7%	
100-300 beds	48%		38%	
More than 300 beds	41%		55%	
Cardiac treatment services				
Coronary ICU	92%		97%	
Cardiac catheterization	69%		76%	
Echocardiology	91%		93%	
Electrocardiology	96%		96%	
Heart surgery services	60%		62%	
Open heart surgery services	56%		61%	
Mobile cardiac services	8%		8%	
Cardiology clinic	22%		32%	
Cardiac tech index	1.63	0.70	1.74	0.62
Teaching hospital	24%		39%	
Control or ownership				
Not-for-profit hospital	73%		58%	
For-profit hospital	18%		13%	
Public hospital	8%		28%	
Rural hospital	9%		5%	
Location				
Nonmetropolitan	4%		3%	
Medium metropolitan	19%		16%	
Large metropolitan	77%		80%	

Table 3: Main regression results

	[1] bivariate	[2] preferred	[3] RN hours	[4] RN wage	[5] RN wage and hours
RN Union 1993	−1.013 ***	−0.836 **	−0.865 **	−0.734 **	−0.762 **
	(0.3492)	(0.354)	(0.3566)	(0.3649)	(0.3671)
RN hours per discharge			−0.015		−0.016
			(0.0223)		(0.0223)
Within-HSA RN wage premium				−0.071	−0.073
				(0.0622)	(0.0622)
Annual discharges (100s)		−0.009 *	−0.009 *	−0.009 *	−0.009 *
		(0.0049)	(0.0049)	(0.0049)	(0.0049)
100 or fewer beds		−0.948	−0.913	−0.997	−0.962
		(0.6404)	(0.6429)	(0.6416)	(0.6439)
More than 300 beds		−0.265	−0.263	−0.229	−0.226
		(0.4242)	(0.4245)	(0.4251)	(0.4255)
Total other hours per discharge		0.0017	0.008	0.0024	0.009
		(0.011)	(0.0143)	(0.011)	(0.0143)
Cardiac tech index		−0.918 ***	−0.866 ***	−0.903 ***	−0.847 ***
		(0.3047)	(0.3142)	(0.3048)	(0.3144)
Teaching hospital		0.2471	0.3006	0.2298	0.2858
		(0.4171)	(0.4245)	(0.4171)	(0.4245)
Staff MDs per discharge		−0.341	−0.261	−0.428	−0.346
		(1.1386)	(1.1452)	(1.1406)	(1.1469)
For-profit hospital		0.8513 *	0.8857 *	0.8278 *	0.8636 *
		(0.4526)	(0.4557)	(0.4529)	(0.4558)
Public hospital		0.2942	0.2879	0.2295	0.2215
		(0.4762)	(0.4766)	(0.4793)	(0.4797)
Rural hospital		−0.353	−0.374	−0.411	−0.436
		(0.7604)	(0.7617)	(0.7618)	(0.7631)
Nonmetropolitan		0.3284	0.3032	0.4182	0.3934
		(1.0154)	(1.0169)	(1.018)	(1.0192)
Medium metropolitan		−0.37	−0.42	−0.334	−0.386
		(0.4238)	(0.4303)	(0.4247)	(0.431)
Intercept	14.959 ***	17.003 ***	17.207 ***	16.954 ***	17.167 ***
	(0.1944)	(0.6047)	(0.6729)	(0.6059)	(0.6734)
Adjusted R^2	0.0211	0.1038	0.1024	0.1047	0.1034

Notes. The dependent variable in mortality rate, 1991–93. The dependent variable is the risk-adjusted hospital AMI mortality rate, 1991–93. Regressions are weighted by the number of AMI cases at each hospital. Standard errors in parentheses.
*Significant at 10 percent.
**Significant at 5 percent.
***Significant at 1 percent.

Table 4: Specification and Selection Tests

	[1] Raw Mortality	[2] Nonhealthcare union specification	[3] Nonhealthcare union selection	[4] Heckman selection correction	[5] Nonhealthcare union only specification	[6] Future RN union specification
RN Union 1993	−0.186		−0.94 **	−0.74		
	(0.3784)		(0.4494)	(0.6072)		
Nonhealthcare Union		−0.481	0.1988		−0.715	
		(0.4193)	(0.5289)		(1.2312)	
Inverse Mills Ratio				−0.082		
				(0.4204)		
RN Union coming						−1.042
						(0.86)
Annual discharges (100s)	−0.008	−0.009 *	−0.009 *	−0.009 *	−0.013 *	−0.014 **
	(0.0052)	(0.0049)	(0.0049)	(0.0049)	(0.0069)	(0.0068)
100 or fewer beds	−0.851	−0.899	−0.953	−0.943	−0.541	−0.499
	(0.6846)	(0.6442)	(0.6414)	(0.6418)	(0.7741)	(0.7732)
More than 300 beds	0.1285	−0.327	−0.244	−0.265	0.4989	0.4932
	(0.4534)	(0.4287)	(0.4283)	(0.4248)	(0.5435)	(0.5402)
Total other hours per discharge	0.0018	0.0012	0.0017	0.0015	−0.006	−0.005
	(0.0118)	(0.0111)	(0.011)	(0.0111)	(0.0128)	(0.0128)
Cardiac tech index	−1.833 ***	−0.92 ***	−0.923 ***	−0.923 ***	−0.691 *	−0.652 *
	(0.3257)	(0.3069)	(0.3053)	(0.3061)	(0.3684)	(0.3693)
Teaching hospital	0.2775	0.2169	0.2466	0.241	−0.332	−0.333
	(0.4458)	(0.4195)	(0.4176)	(0.4188)	(0.545)	(0.5437)
Staff MDs per discharge	−1.749	−0.325	−0.416	−0.376	−0.595	−0.688
	(1.2171)	(1.1627)	(1.1576)	(1.1543)	(1.9687)	(1.9658)
For-profit hospital	0.3532	0.8055 *	0.8616 *	0.8565 *	1.2479 **	1.1915 **
	(0.4839)	(0.4556)	(0.4541)	(0.4541)	(0.5442)	(0.5453)
Public hospital	−0.43	0.1497	0.271	0.2658	0.819	0.7845
	(0.509)	(0.4797)	(0.4808)	(0.4988)	(0.714)	(0.7128)
Rural hospital	−0.098	−0.313	−0.348	−0.354	−0.189	−0.216
	(0.8129)	(0.7652)	(0.7615)	(0.7616)	(0.8549)	(0.8529)
Nonmetropolitan	0.0283	0.2978	0.3211	0.3264	−0.192	−0.175
	(1.0855)	(1.0221)	(1.0169)	(1.017)	(1.1816)	(1.1777)
Medium metropolitan	−0.867 *	−0.379	−0.364	−0.366	−0.879 *	−0.907 *
	(0.453)	(0.4268)	(0.4247)	(0.4248)	(0.5169)	(0.5127)
Intercept	18.71 ***	16.923 ***	16.997 ***	16.997 ***	16.876 ***	16.873 ***
	(0.6464)	(0.6077)	(0.6057)	(0.6065)	(0.7098)	(0.7079)
Adjusted R^2	0.162	0.0923	0.1015	0.1012	0.079	0.0834

Notes. The dependent variable in column 1 is the raw hospital AMI mortality rate, 1991–93. The dependent variable in columns 2 through 6 is the risk-adjusted hospital AMI mortality rate, 1991–93. Regressions are weighted by the number of AMI cases at each hospital. Standard errors in parentheses.

*Significant at 10 percent.

**Significant at 5 percent.

***Significant at 1 percent.

Table 5: Longstanding union

	[1] bivariate	[2] preferred	[3] RN hours	[4] RN wage	[5] RN wage and hours
RN Union long	−1.122 ***	−0.964 **	−0.991 ***	−0.858 **	−0.884 **
	(0.3614)	(0.3739)	(0.3762)	(0.3872)	(0.3891)
RN Union new	0.0369	0.0974	0.0605	0.1057	0.067
	(0.9628)	(0.9487)	(0.9511)	(0.9486)	(0.9509)
RN hours per discharge			−0.015		−0.016
			(0.0223)		(0.0223)
Within-HSA RN wage premium				−0.066	−0.067
				(0.0624)	(0.0625)
Annual discharges (100s)		−0.009 *	−0.009 *	−0.008 *	−0.009 *
		(0.0049)	(0.0049)	(0.0049)	(0.0049)
100 or fewer beds		−0.912	−0.879	−0.961	−0.926
		(0.6412)	(0.6437)	(0.6428)	(0.6451)
More than 300 beds		−0.267	−0.265	−0.234	−0.23
		(0.4241)	(0.4244)	(0.4252)	(0.4256)
Total other hours per discharge		0.0012	0.0074	0.0019	0.0084
		(0.011)	(0.0143)	(0.011)	(0.0144)
Cardiac tech index		−0.904 ***	−0.853 ***	−0.89 ***	−0.836 ***
		(0.3049)	(0.3144)	(0.3051)	(0.3147)
Teaching hospital		0.2896	0.3414	0.2697	0.3238
		(0.4189)	(0.4262)	(0.4193)	(0.4264)
Staff MDs per discharge		−0.312	−0.235	−0.395	−0.316
		(1.1387)	(1.1453)	(1.1412)	(1.1475)
For-profit hospital		0.902 **	0.935 **	0.8756 *	0.9098 **
		(0.4551)	(0.4581)	(0.4557)	(0.4585)
Public hospital		0.3825	0.3754	0.3147	0.3057
		(0.4833)	(0.4838)	(0.4875)	(0.488)
Rural hospital		−0.37	−0.391	−0.422	−0.446
		(0.7605)	(0.7617)	(0.762)	(0.7633)
Nonmetropolitan		0.2001	0.1767	0.2947	0.2723
		(1.0224)	(1.0238)	(1.0262)	(1.0274)
Medium metropolitan		−0.432	−0.48	−0.392	−0.443
		(0.4276)	(0.4339)	(0.4292)	(0.4352)
Intercept	14.959 ***	16.971 ***	17.17 ***	16.928 ***	17.137 ***
	(0.1943)	(0.6053)	(0.6738)	(0.6066)	(0.6743)
Adjusted R^2	0.0222	0.1042	0.1027	0.1045	0.1031

Notes. The dependent variable is the risk-adjusted hospital AMI mortality rate, 1991–93. Regressions are weighted by the number of AMI cases at each hospital. Standard errors in parentheses.
*Significant at 10 percent.
**Significant at 5 percent.
***Significant at 1 percent.

Table 6: Longstanding union: specification and selection tests

	[1]	[2]
		Nonhealthcare
	Raw	Union
	Mortality	Selection
RN Union long	−0.145	−1.098 **
	(0.4003)	(0.4717)
RN Union new	−0.49	0.0024
	(1.0158)	(0.9717)
Nonhealthcare Union		0.2462
		(0.5305)
Annual discharges (100s)	−0.008	−0.009 *
	(0.0052)	(0.0049)
100 or fewer beds	−0.863	−0.917
	(0.6865)	(0.6421)
More than 300 beds	0.1292	−0.242
	(0.454)	(0.4282)
Total other hours per discharge	0.0019	0.0013
	(0.0118)	(0.011)
Cardiac tech index	−1.838 ***	−0.909 ***
	(0.3265)	(0.3055)
Teaching hospital	0.2636	0.2905
	(0.4485)	(0.4194)
Staff MDs per discharge	−1.758	−0.404
	(1.2191)	(1.1573)
For-profit hospital	0.3367	0.9165 **
	(0.4872)	(0.4567)
Public hospital	−0.459	0.3569
	(0.5174)	(0.487)
Rural hospital	−0.092	−0.365
	(0.8142)	(0.7614)
Nonmetropolitan	0.0701	0.1864
	(1.0946)	(1.0241)
Medium metropolitan	−0.847 *	−0.426
	(0.4578)	(0.4283)
Intercept	18.72 ***	16.962 ***
	(0.6481)	(0.6063)
Adjusted R^2	0.1597	0.102

Notes. The dependent variable in column 1 is the raw hospital AMI mortality rate, 1991–93. The dependent variable in columns 2 and 3 is the risk-adjusted hospital AMI mortality rate, 1991–93. Regressions are weighted by the number of AMI cases at each hospital. Standard errors in parentheses.
*Significant at 10 percent.
**Significant at 5 percent.
***Significant at 1 percent.

Table 7: Inter-year correlation of risk-adjusted AMI mortality, by caseload

Panel A

Inter-year correlation of risk-adjusted AMI mortality for hospitals with 51 or more cases per year.

	1991	1992	1993
1991	1.000		
1992	0.354	1.000	
	[0.000]		
1993	0.373	0.327	1.000
	[0.000]	[0.000]	

Panel B

Inter-year correlation of risk-adjusted AMI mortality for hospitals with 50 or fewer cases per year.

	1991	1992	1993
1991	1.000		
1992	0.016	1.000	
	[0.835]		
1993	0.109	0.171	1.000
	[0.149]	[0.343]	

Top value is Pearson's r. Bottom value in brackets is p-value.