

THE WHITE HOUSE

WASHINGTON

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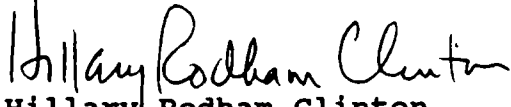
Dear Dr. Lynn:

Thank you for writing to describe your research on the impact of insurance status on mortality. I enjoyed meeting with you in New Hampshire and greatly appreciate your taking the time to send information about the SUPPORT project.

As you noted, I mentioned during our conversation that uninsured patients have higher hospital mortality rates than well-insured patients. I based my comment on two studies reported in the Journal of the American Medical Association (JAMA). The first found that, after adjustment for all other factors that impact upon morbidity, uninsured Americans have an in-hospital death rate up to three times higher than the privately insured. This study, published by Jack Hadley, Earl Steinberg and Judith Feder, is listed on the bibliography that you enclosed. A second study, published by Peter Franks, Carolyn Clancy and Marthe Gold in JAMA in August of 1993, also concluded that the lack of health insurance is associated with an increased risk of subsequent mortality. These studies do not include data on individuals with Medicaid coverage.

I have forwarded your study to the Administration staff members who are working on national health reform. Thank you again for writing.

Sincerely yours,


Hillary Rodham Clinton

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Booklet, "The SUPPORT Study," 13 pages

The SUPPORT Study



*The Program on the
Care of Critically Ill
Hospitalized Adults*



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December 22, 1993

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Dear Mrs. Clinton:

I am writing to follow up on the effects of being uninsured or on Medicaid on mortality for hospitalized patients. As you may recall, in our conversation over lunch at Dartmouth you stated that uninsured and Medicaid patients have some three times the hospital mortality of well insured patients. I mentioned that my research (SUPPORT, The Study to Understand Prognoses and Preferences for Outcomes and Risks of Treatments, funded by the Robert Wood Johnson Foundation) had found no impact of insurance status on mortality. I will here present a brief summary of the results of my study in this regard, along with a brief narrative review of the literature. I will enclose a short bibliography and a brochure about the SUPPORT project. If you or any advisors want to review the methods or analysis, please let me know.

SUPPORT enrolls all adults who live 48 hours after admission to one of five diverse teaching hospitals and who have one of nine very serious diseases. In the first two years of the project, we enrolled 4301 patients, of whom 26% died before discharge and 48% died within six months. We used newly developed statistical techniques to create a formula by which to predict the likelihood of survival to each day over the next six months for each patient like those enrolled in the project. (Previous work addressed survival to discharge or grouped patients into large groups.) The formula works very well, even sustaining its predictive ability when applied to a new group of patients.

This prognostic formula takes into account age, diagnoses, hospital course, and degree of physiological derangement in various lab tests and measures of vital signs. Once these elements are taken into account, none of our measures of socio-economic status showed any correlation with survival time. In other words, having no insurance, or Medicaid only, did not predict *any* difference in expected survival. Likewise, there was no effect of race, income, or education. There was not even an effect of which hospital provided care, and our hospitals included a city hospital, as well as one which is largely private-physician, two large university research and teaching hospitals, and a rural medical center.

Within SUPPORT, then, one can only conclude that very seriously ill patients who are in the hospital for more than two days have the same risks of dying over the ensuing six months without regard to insurance, other measures of socio-economic status, or which of our five hospitals provided care.

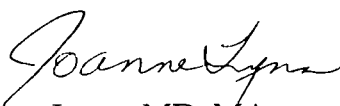
This does not mean that poor, uninsured, and underinsured patients get the same care and the same outcomes as well-insured patients. I would expect that the poor among us come to hospitals later, sicker, and with more complications, tend to present to less capable institutions, and may well have higher mortality in the first 48 hours. The evidence for these claims is frail but suggestive. All of the studies directly addressing hospital mortality drew data from administrative data banks, which adjust for severity only very roughly. Thus, case mix differences of substantial size are expected to persist. In addition, most of the studies have serious unmeasured biases within their definition of the populations to study, or require substantial unjustified interpretative maneuvers to estimate the outcomes of interest (rather than those available). We have included some explanation of the problems in the enclosed summary of the literature.

Nevertheless, Medicaid and uninsured patients probably have more environmental and lifestyle reasons for illness, fewer social resources to aid recuperation or disease management, more barriers to preventive care and early access, more discontinuity in care, and fewer tests and procedures. Specifically on the question of differential hospital care, Hadley et al. (1991) showed that uninsured persons had more hospital deaths, though most odds ratios were between 1.0 and 1.5. Certainly the uninsured get fewer diagnostic procedures (Hadley et al. (1991), Braveman et al. (1991), Wenneker et al. (1990)) and delay care more often (Weissman et al. (1989)). The delays almost certainly have an adverse effect on health and on costs. However, the impact of the reduction in diagnostic tests is uncertain. In general, increased use of hospitalization might have an adverse effect on mortality, at least among the elderly (Fisher et al., under review).

Once established in the hospital, however, there is little evidence that very sick uninsured and Medicaid patients receive care that is deficient in terms of enhancing survival, as compared to those who are well-insured.

I hope this information proves helpful. Let me know if you see any other issue where I might be of help. I greatly appreciated the opportunity to meet and talk with you. Health care reform needs to go farther and faster than Congress yet imagines. I admire your work and wish you well.

Sincerely,


Joanne Lynn, MD, MA

Insurance and Hospital Mortality Narrative Summary of Key Literature

Several studies have examined the effect of insurance status on outcomes and the utilization of hospitals. Although the studies generally report lower utilization and higher mortality among uninsured and Medicaid patients compared to patients with private insurance, the biases in these investigations are substantial and deserve to be acknowledged.

A study by Hadley et al. (1991) examined condition on admission, resource use, and in-hospital mortality by insurance status. The authors found that uninsured patients were 29% to 75% less likely to receive certain high cost or high discretion procedures such as coronary artery bypass graft surgery or total hip replacement. This report excluded Medicaid patients which leaves many questions unanswered. When the population was split by age and race, 5 of 16 groups showed no effect of insurance on hospital mortality, 7 showed effects between 1.0 and 1.6 odds, and 5 showed effects greater than odds of 1.6 (girls of both races and young Black males). The actual number of deaths studied was quite small.

Studies by Wenneker et al. (1990) and Langa et al. (1993) examined utilization of cardiac procedures by insurance status. The Wenneker study (1990) looked at patients admitted to Massachusetts hospitals with circulatory disorders or chest pain. The authors found, after adjusting for demographic, clinical, and hospital factors, that privately insured patients were 80% more likely than the uninsured to receive angiography, 40% more likely to receive bypass grafting, and 28% more likely to receive angioplasty. Medicaid patients had a 48% lower chance of receiving angioplasty compared to the privately insured, but in other respects were similar to the uninsured. The effects on outcome were not studied.

In the study by Langa et al. (1993), the authors found that the rate of increase in frequency of coronary revascularization procedures was much faster among patients with private fee-for-service or HMO insurance than it was for patients on Medicaid between 1983 and 1988. In 1983, fee-for-service and HMO patients were 1.66 and .96 times as likely, respectively, of obtaining revascularization procedures as patients with Medicaid insurance. However, by 1988, patients with private fee-for-service insurance were 2.33 times as likely as Medicaid patients to undergo revascularization and HMO patients were 1.53 times as likely. Young and Cohen (1992) also found that Medicaid patients were less likely to receive certain coronary procedures. Additionally, Weissman and Epstein (1989) reported that among their Boston area sample of hospitalized patients, the uninsured underwent, on average, only 7% fewer procedures than the Blue Cross patients. The differences also varied by type of hospital. These studies use administrative data only and thus have quite inadequate adjustment for severity.

A study by Burns et al. (1993) found no real difference in utilization between patients in Arizona's experimental Medicaid program and privately insured patients. The authors studied patients with 11 different conditions and found that among medical, surgical, and pediatric diagnoses there were "few significant differences" between Medicaid and non-Medicaid patients after controlling for severity of illness, specific hospital used, and several patient, hospital, and physician factors.

Some other studies set out to discover whether the uninsured and Medicaid patients were receiving a lower quality of care than privately insured patients by examining length of hospital stay among the different groups. By looking at all newborns with serious medical problems discharged in 1987 from all California civilian, acute-care hospitals, Braveman et al. (1991) reported that although uninsured and Medicaid babies generally had more complications than the uninsured, they received fewer services. Mean length of stay was 13.2 days for uninsured newborns and 14.8 days for newborns covered by Medicaid, while for privately insured newborns, the mean length of stay was 15.7 days. Complications coded for insurance purposes are not a reliable indication of severity of illness, and the fact that these investigators used this as the only adjustment for severity will have introduced an unknown bias. In addition, sick infants with unsupportive home environments may have been kept in the hospital longer, an effect that could completely explain the results.

Many of the studies comparing quality of care by insurance status also report on in-hospital mortality. The aforementioned Hadley study (1991) found an in-hospital mortality rate 0.9 to 3.2 times higher for uninsured compared to insured patients. In the Young and Cohen study (1992), the risk of mortality was almost twice as high for Medicaid patients as it was for those with private insurance after adjusting for clinical, sociodemographic, and hospital-level factors. In a study of the effect of insurance status among breast cancer patients, Ayanian et al. (1993) found that the risk of mortality in breast cancer, adjusted for age, race, marital status, household income, coexisting diagnoses, and disease stage was 49% higher for uninsured patients and 40% higher for Medicaid patients than for those with private insurance. All of these comparisons were statistically significant. The Burns study (1993) of patients in Arizona, however, did not find a substantial difference in mortality between patients insured by Medicaid and those with private insurance after controlling for severity of illness, specific hospital used, and several patient, hospital, and physician factors.

Only SUPPORT has been able to adjust for severity of disease for hospital patients, but it has the limitations of reflecting only five hospitals, reporting only those who were hospitalized and lived 48 hours, and reporting only very seriously ill adults.

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Comparison of Uninsured and Privately Insured Hospital Patients

Condition on Admission, Resource Use, and Outcome

Jack Hadley, PhD; Earl P. Steinberg, MD, MPP; Judith Feder, PhD

To investigate the association between insurance status and condition on admission, resource use, and in-hospital mortality, we analyzed discharge abstracts for 592 598 patients hospitalized in 1987 in a national sample of hospitals. In 13 of 16 age-sex-race-specific cohorts, the uninsured had a 44% to 124% higher risk of in-hospital mortality at the time of admission than did the privately insured. After controlling for this difference, the actual in-hospital death rate was 1.2 to 3.2 times higher among uninsured patients in 11 of 16 cohorts. The uninsured also were 29% to 75% less likely to undergo each of five high-cost or high-discretion procedures and 50% less likely to have normal results on tissue pathology reports for biopsies performed during five of seven different endoscopic procedures. Our results suggest that insurance status is associated with a broad spectrum of aspects of hospital care.

(JAMA. 1991;265:374-379)

INCREASING evidence suggests that the amount and content of medical care that an individual receives in the United States is related to whether the individual has health insurance.¹ Previous studies, for example, although uncontrolled for differences between groups in clinical status and need for services on admission, have shown that the uninsured have fewer physician visits per year, are less likely to have a hospital stay, and, if hospitalized, have a shorter length of stay than insured individuals.²⁻⁴ Other studies have shown that, even after controlling for differences in admission diagnosis, the services that uninsured individuals receive are different from those that insured individuals receive.⁵⁻⁷ In a recent analysis of 1983 hospital discharge data for 65 032 uninsured, Blue Cross, and Medicare patients hospitalized in the Boston area, for example, Weissman and Epstein⁸

found, after adjusting for case mix, that insured patients had 7% shorter stays and underwent 7% fewer procedures than Blue Cross patients. More recently, Wenneker et al⁹ found, after controlling for clinical and demographic factors, that insurance status was strongly associated with utilization of three expensive cardiac procedures among 28 000 patients admitted to Massachusetts hospitals in 1985 with circulatory disorders or chest pain. Very few data are available regarding differences in health outcomes among insured vs uninsured patients.⁴

The goal of this study is to identify whether there are statistically significant differences between uninsured and privately insured patients in three sets of factors related to hospital care: condition on admission, resources used in the hospital, and outcome. Specifically, we ask whether the uninsured are sicker when admitted to the hospital; whether fewer resources are expended on their care in the hospital, given their condition on admission; and whether they have a poorer outcome, given their condition on admission.

The research reported herein extends prior work in four important

ways. First, we used a large national sample of 592 598 hospital discharges to compare the characteristics of hospital care received by uninsured and privately insured patients. Thus, our findings should be more generalizable to the nation's uninsured than those of studies based on a single, relatively small geographic area, such as Boston or Massachusetts. Second, the only previous study that has examined insured vs uninsured patients' use of high-cost procedures focused exclusively on three cardiac procedures.⁹ In contrast, while controlling for diagnosis, we analyzed the frequency with which a broad spectrum of high-cost invasive and noninvasive procedures are provided to insured vs uninsured hospitalized patients. Third, we used multiple measures of condition on admission and resource use, which allow us to explore these factors in greater depth. Finally, we examined differences between the uninsured and privately insured in in-hospital death rates, an important measure of outcome.

DATA, VARIABLES, AND STATISTICAL METHODS

Data

The data set was derived from discharge abstracts submitted to the Commission on Professional and Hospital Activities, which is a research and data abstracting service for hospitals, by its member hospitals in 1987. The entire Commission on Professional and Hospital Activities 1987 database includes over 10 million discharge abstracts from approximately 1200 hospitals. For this study, we examined only discharges for people between the ages of 1 and 64 years whose primary source of payment was no charge, self-pay, Blue Cross, insurance company, or Medicaid and whose race was white, black, or Hispanic. Discharges for which any of this in-

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formation was unrecorded were deleted from our data set. We excluded newborns, infants under the age of 1 year, and women with a pregnancy-related diagnosis because these discharges account for a much higher proportion of uninsured cases than of privately insured cases.¹⁰ We believed that without such exclusions, aggregate measures of case severity might be misleading.

The data file used in our statistical analyses consisted of all no charge and self-pay discharges plus a 5% random sample of all other discharges. (The sampled records were weighted by the inverse of their sampling probability in all analyses.) There were 592 598 discharges in this file. For some parts of our study, attention was limited to discharges with particular diagnoses or in which particular procedures were performed.

Variables

We defined "uninsured patients" as those having either no charge or self-pay listed as the primary source of payment and "privately insured patients" as those having either Blue Cross or an insurance company listed as the primary source of payment. Our database and statistical analyses also included Medicaid patients as another insured group. However, we only report results associated with uninsured and privately insured patients because it is very difficult to draw conclusions about the insurance effects of Medicaid without state-specific information on Medicaid coverage and payment policies, which vary substantially across states.^{11,12}

The analysis consists of a series of comparisons of specific measures, constructed from the discharge abstract form, which can be interpreted as providing evidence regarding condition on admission, resource use, or outcome. The dependent variables, which we analyzed using multiple regression methods, are summarized in Table 1 and discussed below.

Risk-Adjusted Mortality Index.—The Risk-Adjusted Mortality Index (RAMI), developed by the Commission on Professional and Hospital Activities, is the expected in-hospital mortality rate based on actual in-hospital mortality rates for diagnoses, grouped by their diagnosis related group code, adjusted for patient age, race, sex, the presence of comorbidities (secondary diagnoses at time of admission), and the risk of death associated with comorbidities and the principal operative procedure (if any).¹³ A higher RAMI value can be interpreted as representing a more severely ill patient.¹⁴

Probability of a Weekend Admis-

sion.—The choice of the second measure of condition on admission is based on the hypothesis that a weekend (Friday through Sunday) admission is likely to be more urgent than a weekday admission. In other words, we assumed that elective admissions are more likely to occur during a weekday because of private physicians' schedules, patient preferences, and hospital staffing patterns. Although many scheduled admissions are for serious conditions, we posited that, on average, they are less urgent or immediately life-threatening than the average weekend admission.

Average Length of Stay.—Length of stay is a common measure of resource use. However, even when controlling for diagnosis, length of stay can reflect the severity of a patient's condition as well as the level of resource use for a patient with a given level of severity. We attempted to control for the effects of severity by focusing the analysis of length of stay on diagnoses that have either a high or a low degree of physician discretion with regard to the need for a hospitalization. Using both prior literature^{15,17} and clinical judgment, we initially selected 28 diagnoses thought to differ in the amount of discretion, from the physician's perspective, in the need for a hospital admission. Relatively high-discretion diagnoses may be thought of as being more elective, while relatively low-discretion diagnoses may be considered nonelective.

We then used linear regression models for each of the 28 diagnoses to estimate the RAMI for the uninsured, controlling for age, sex, race, hospital size, ownership, teaching status, and type of community. The 10 diagnoses reported in the analysis are the ones with the five lowest and five highest predicted RAMI scores derived from the regression models. Estimating length of stay equations separately for each diagnosis decreases the likelihood that differences in resource use will be obscured by differences in diagnostic mix between uninsured and privately insured patients.

Probability of a High-Cost and/or High-Discretion Procedure.—If insurance coverage affects resource use, we hypothesized that the uninsured would be less likely to be admitted for procedures that are very costly and/or somewhat discretionary in terms of their clinical necessity. Using these criteria, we selected five procedures—coronary artery bypass graft surgery, total knee replacement, total hip replacement, stapedectomy, and surgical correction of strabismus—and tested whether the probability of undergoing each of these procedures varies with insurance status. We used our entire

sample of discharges, rather than those associated with a limited number of primary diagnoses, for these analyses because we wished to examine the probability that each of these procedures was performed for any reason.

Probability of a Specific Procedure Associated With a Particular Set of Diagnoses.—To better control for the effects on resource use of variations in the mix of diagnoses across groups of patients with different insurance status, we analyzed the probability of a patient's receiving a particular procedure or set of procedures after limiting the sample to discharges with a narrowly defined set of *International Classification of Diseases, Version 9*, codes as the primary discharge diagnosis. Five separate samples of discharges, each with a different set of limited diagnoses, were drawn. (The diagnoses in each group and the related procedure codes are shown in Table 4.) Each of these groups was analyzed separately to determine whether there was a significant difference in the use of the associated procedure between privately insured and uninsured patients.

We also examined two procedures that are often performed in conjunction with another procedure: biopsy, which often occurs with a bronchoscopy, and intraoperative cholangiogram, performed in conjunction with a cholecystectomy. All discharges involving patients who underwent either a bronchoscopy or a cholecystectomy, respectively, are the samples we employed for these two analyses.

Probability of a 'Not Abnormal' Biopsy Result.—The last resource use measure that we examined is the probability of a tissue pathology report of "not abnormal" during hospitalizations in which any of seven different endoscopy procedures (see Table 5) and an associated biopsy were performed. (The other possible findings on the pathology report are abnormal—other, mechanical abnormality, growth alteration, degeneration and necrosis, nonacute inflammation, acute inflammation, nonmalignant neoplasm, and malignant neoplasm.) We hypothesized that a relatively higher probability of a "not abnormal" finding in a tissue pathology report for one group of patients is suggestive of a more aggressive management strategy's being employed in the care of that group of patients compared with a reference group. This hypothesis assumes that the clinical threshold for performing the test and doing a biopsy is lower for the group of patients with the higher probability of "not abnormal" findings in tissue pathology reports than for the other group of patients.

Table 1.—Dependent Variables: Measures of Condition on Admission, Resource Use, and Outcome

Condition on Admission
Risk-Adjusted Mortality Index
Probability of a weekend admission
Resource Use
Average length of stay for each of 10 specific diagnoses
Incidence of selected high-cost or high-discretion procedures
Coronary artery bypass graft surgery
Total knee replacement
Total hip replacement
Stapedectomy
Surgical correction of strabismus
Incidence of selected procedures, controlling for diagnosis
Colonoscopy or proctosigmoidoscopy
Upper endoscopy procedures
Coronary arteriography
Contrast myelogram or CT* of spine
Head CT
Bronchoscopy with biopsy
Cholecystectomy with intraoperative cholangiogram
Tissue pathology of "not abnormal" for biopsy specimens obtained during selected endoscopic procedures
Outcome
In-hospital death rate

*CT indicates computed tomography.

Outcome.—We examined one outcome measure in our analysis, the probability of dying in the hospital. Although this measure has several limitations, it is one of the most frequently used indicators of the outcome of care. This variable was analyzed using the full set of discharges stratified into 16 age-sex-race-specific cohorts.

Statistical Methods

Multiple regression analysis was used to test for statistically significant differences between uninsured and privately insured patients in the various measures of condition on admission, resource use, and outcome. The regression models included as independent variables insurance status and controls for the effects of characteristics not held constant by stratification.

The results reported in Tables 2 through 6 are based on the regression estimates of the coefficients for uninsured and privately insured patients from equations of the following general form:

$$Y = \alpha U + \beta PI + \sum \mu X_i$$

where Y represents a measure of condition on admission, resource use, or outcome; α , the coefficient for uninsured patients; β , the coefficient for privately insured patients; U equals 1 if the patient is uninsured, 0 otherwise; PI equals 1 if the patient is privately insured, 0 otherwise; and X_i , other control variables.

The exact list of X_i variables depends on the particular dependent variable being analyzed and the extent of file strati-

Table 2.—Measures of Condition on Admission: Regression-Adjusted Coefficients for Uninsured Relative to Privately Insured by Age-Sex-Race Cohort*

Sex, Age, and Race	Unweighted N	Relative Regression-Adjusted Coefficient	
		Risk-Adjusted Mortality Index	Probability of a Weekend Admission
Males			
Age 1-17 y			
White	36 527	1.57	1.13†
Black	8813	1.56‡	1.00
Age 18-34 y			
White	74 514	1.82†	1.22†
Black	17 614	2.09†	1.24†
Age 35-49 y			
White	46 429	1.80†	1.17†
Black	11 384	1.59†	1.27†
Age 50-64 y			
White	54 417	1.45†	1.14†
Black	9175	1.44†	1.23†
Females			
Age 1-17 y			
White	29 724	2.00†	1.12†
Black	7036	1.81	1.12‡
Age 18-34 y			
White	72 780	1.80†	1.21†
Black	15 828	1.04†	1.24†
Age 35-49 y			
White	52 371	1.62†	1.16†
Black	11 336	2.24	1.15†
Age 50-64 y			
White	61 611	1.52†	1.11†
Black	10 395	1.86†	1.07

*Null hypothesis is that relative rate or probability equals 1.00.

† $P \leq .01$.

‡ $0.01 < P \leq .05$.

fication. In all analyses, the individual discharge was the unit of observation.

The coefficients α and β can be interpreted as mean values of Y for uninsured and privately insured patients, respectively, adjusting for differences in Y due to the X_i variables. The primary null hypothesis throughout the analysis is that there is no difference in Y between uninsured and privately insured patients, controlling for the effects of other factors, which is equivalent to the null hypothesis $\alpha = \beta$. This hypothesis can be evaluated using a t statistic calculated from the variance-covariance matrix of the regression parameters.²⁸ Note that because the regression model is linear in the parameters, the test of the null hypothesis is independent of the particular values of X_i chosen to standardize the values of α and β .

The tables of results report relative regression-adjusted coefficients, which are defined as the ratio of the adjusted mean of Y for the uninsured to the adjusted mean of Y for the privately insured, again standardized for a particular set of X_i values. Mathematically this is simply $(\alpha + \sum \mu X_i) / (\beta + \sum \mu X_i)$.

For example, if the value of the ratio is 1.10, one can then say, "The rate for the uninsured is 10% greater than for the privately insured, or the probability of the uninsured's receiving a particular

service is 10% greater than for the privately insured."

Controls for Factors Other Than Insurance Coverage

All analyses of differences between uninsured and privately insured patients were adjusted for age, sex, and race, either by stratification or by including age, sex, and race as control variables in the regression models. The RAMI score, the probability of a weekend admission, and the probability of an in-hospital death were analyzed separately for 16 age-sex-race-specific cohorts of discharges. The regression models for these cohort-specific analyses also included control variables for hospital size, ownership (public or private), teaching status (Council of Teaching Hospitals member), and type of community (100 largest cities, other metropolitan community, or nonmetropolitan). In addition, the analysis of the probability of a weekend admission included the RAMI score and the Medicare case-mix index as controls for possible differences in diagnostic mix and severity of illness. The analysis of in-hospital mortality included as additional control variables the RAMI score, the Medicare case-mix index, and an indicator of whether a surgical procedure was performed.

The regression model for average

Table 3.—Length of Stay: Regression-Adjusted Coefficients for Uninsured Relative to Privately Insured, by 10 Specific Diagnoses*

Diagnosis, by Degree of Physician Discretion and Estimated RAMI (ICD-9 Code)	Unweighted N	Estimated RAMI Value†	Relative Regression-Adjusted Average Length of Stay
High Physician Discretion			
Chronic tonsillitis (474.0)	3815	.001	0.88‡
Noninfectious gastroenteritis, NOS (558.9)	8480	.001	0.82‡
Acute bronchitis (466.0)	2884	.003	0.85§
Unilateral inguinal hernia (560.91)	4144	.004	0.82‡
Uterine leiomyoma, NOS (218.9)	3147	.005	0.77§
Low Physician Discretion			
Gastrointestinal hemorrhage, NOS (578.9)	1600	.028	0.89
Malignant neoplasm, bronchus/lung (162.9)	1548	.036	0.97
Congestive heart failure (428.0, 428.1, 428.9)	2768	.051	0.94
Acute myocardial infarction, inferior wall, NEC (410.4)	2512	.066	0.89‡
Acute myocardial infarction, anterior wall (410.0)	1788	.115	0.95

*Null hypothesis is that relative length of stay equals 1.00. RAMI indicates Risk-Adjusted Mortality Index; ICD-9, International Classification of Diseases, Version 9; NOS, not otherwise specified; and NEC, not elsewhere classified.

†Estimated values of the RAMI from a regression model that includes insurance status, age, sex, race, hospital size, ownership, teaching status, and type of community as control variables.

‡ $P \leq .01$.

§ $.01 < P \leq .05$.

length of stay was estimated separately for each of 10 specific diagnoses. The control variables consisted of age, sex, race, hospital size, ownership, teaching status, and community type, as well as the RAMI, the Medicare case-mix index, and an indicator of whether any surgical procedure was performed. The purpose of the latter three variables is to control for differences in severity within the diagnoses analyzed. The regression model for the incidence of selected procedures and for tissue pathology outcomes included control variables for age, sex, race, hospital size, ownership, teaching status, and community type. The analyses related to the five high-cost or high-discretion procedures, as well as those related to condition on admission and in-hospital mortality, were estimated using the entire sample of 592 598 discharges. The analyses of the other procedures, length of stay, and tissue pathology outcomes were limited to discharges with selected diagnoses or procedures.

RESULTS

Condition on Admission

Table 2 reports the results from the analysis of the two measures of condition on admission, the RAMI and the probability of being admitted on a weekend. Regression-adjusted coefficients are reported for the uninsured relative to the privately insured for each of 16 age-sex-race-specific cohorts of discharges. The uninsured have a significantly higher RAMI for 13 of the 16 cohorts and a significantly higher probability of being admitted during the weekend for 14 of the cohorts. The relative coefficients thus strongly suggest

differences between the uninsured patients and privately insured patients and may reflect more serious or urgent conditions among the uninsured.

Resource Use

Table 3 reports the relative lengths of stay for discharges from each of 10 selected high- and low-discretion diagnoses. The uninsured's length of stay is significantly shorter than the length of stay for the privately insured for each of the five high-physician discretion diagnoses, by 12% to 38%. Although the uninsured also have shorter stays than the privately insured for the low-discretion diagnoses, the differences are smaller, and only one is statistically significant.

In Table 4 we present further results pertinent to the question of differences in resource use between the uninsured and insured. Using the entire sample of discharges, we estimated regression models of the probability of undergoing each of five high-cost or high-discretion procedures among the privately insured and the uninsured. The results suggest highly significant and quantitatively large differences between uninsured and privately insured patients in the probability of undergoing these procedures. The uninsured are 29% less likely to have coronary artery bypass graft surgery and almost 75% less likely to have a total knee replacement. Rates for the other three procedures range from between 15% and 56% lower for the uninsured.

In Table 4, we also examine the relative use of other specific procedures, controlling for diagnosis by limiting each analysis to discharges with a spe-

cific diagnosis, group of diagnoses, or associated procedure. These analyses thus provide an opportunity to examine whether the clinical management of specific conditions varies with insurance status. In three of the 11 diagnosis-procedure combinations—colonoscopy or proctosigmoidoscopy, upper endoscopy, and coronary arteriography—there is a statistically significant difference ($P < .01$) in the probability of an uninsured person's undergoing the procedure relative to a privately insured person. In these instances, uninsured patients were from 20% to 44% less likely to have the procedure performed. We found no statistically significant difference ($P > .05$) between the uninsured and the privately insured in the use of myelography or computed tomography of the spine among patients with lumbar disk displacement, in the use of computed tomography of the head among patients with any of five diagnoses, in the incidence of biopsy among patients undergoing a bronchoscopy, or in the incidence of an intraoperative cholangiogram among patients undergoing a cholecystectomy.

In Table 5, we examine tissue pathology findings for biopsies performed during any of seven endoscopic procedures. The sample in each case consists of all discharges that had both a biopsy performed in association with the specific procedure and a tissue pathology report recorded on the discharge abstract. In each case, as well as for all the endoscopic procedures combined, we estimated the likelihood of a finding of "not abnormal," ie, the probability of a completely normal test result. As Table 5 shows, the uninsured are consistently less likely than privately insured patients to have a completely normal tissue pathology result. The differences are statistically significant in three cases. For five of the seven procedures, the relative probabilities are approximately .5 or less, ie, an uninsured person is less than half as likely as a privately insured person to have a biopsy pathology finding that is completely normal.

Outcome

In Table 6, we examine whether there are significant differences between privately insured and uninsured patients in the probability of an in-hospital death. The regression model includes as control variables the Medicare case-mix index, the RAMI, and whether the patient had a procedure, along with hospital characteristics and type of community. In this analysis we divided the full sample of discharges into 16 age-sex-race-specific groups, which were then analyzed separately.

Table 4.—Probability of Selected Procedures: Regression-Adjusted Coefficients for Uninsured Relative to Privately Insured*

Procedure (Procedure Code)	Unweighted N	Relative Regression-Adjusted Coefficient
High cost or high discretion	592 538	
Coronary artery bypass graft surgery (36.10-36.19)		0.71†
Total knee replacement (81.41)		0.26†
Total hip replacement (81.5, 81.51, 81.59)		0.55†
Stapedectomy (19.1-19.2)		0.50†
Surgical correction of strabismus (15.1-15.9)		0.44†
Colonoscopy (45.23-45.26) or proctosigmoidoscopy (48.23)	35 977	0.56†
Upper endoscopy procedures (42.23, 44.13, 45.13)	47 732	0.80†
Coronary arteriography (88.55-88.57)	30 622	0.61†
Contrast myelogram or CT of spine† (87.21, 88.36)	3243	0.95
Head CT (87.03) by diagnosis		
Convulsions (DX code 780.3)	11 582	0.91‡
Syncope and collapse (DX code 780.2)	4363	0.90
Concussion, NOS (DX code 850.9)	3230	0.88
Concussion without loss of consciousness (DX code 850.0)	2268	0.86
Concussion with brief loss of consciousness (DX code 850.1)	2684	1.02
Bronchoscopy (33.22, 33.23) procedures with a biopsy (33.24, 33.25)	3713	0.98
Cholecystoscopies (51.22) with an intraoperative cholangiogram (87.53)	9500	0.94

*Null hypothesis is that relative coefficients equal 1.00. CT indicates computed tomography; DX, discharge diagnosis; and NOS, not otherwise specified.

† $P \leq .01$.

‡All discharges that had any of the following as the primary discharge diagnosis were considered: noninfectious gastroenteritis (DX558.8), abdominal pain (DX789.0), acute pancreatitis (DX577.0), viral enteritis (DX008.8), gastroenteritis hemorrhage (DX578.9).

§All discharges that had any of the following as the primary discharge diagnosis were considered: noninfectious gastroenteritis (DX558.8), abdominal pain (DX789.0), viral enteritis (DX008.8), acute pancreatitis (DX577.0), acute gastritis (DX535.0), gastroenteritis hemorrhage (DX578.9), esophagitis (DX530.1), atrophic gastritis (DX535.1), gastric mucosal hypertrophy (DX535.2), alcoholic gastritis (DX535.3), gastritis/duodenitis (DX535.5), other specified gastritis (DX535.4).

||All discharges that had any of the following as the primary discharge diagnosis were considered: precordial pain (DX785.1), chest pain (DX785.0), angina pectoris (DX413.9), intermediate coronary syndrome (DX411.1), coronary atherosclerosis (DX414.0).

¶All discharges that had lumbar disk displacement (DX722.10) as the primary discharge diagnosis were considered.

‡.05 < $P \leq .10$.

Table 5.—Tissue Pathology for Selected Procedures: Regression-Adjusted Coefficients for Uninsured Relative to Privately Insured*

Procedure (Procedure Code)	Unweighted N	Relative Regression-Adjusted Coefficient for Probability of "Not Abnormal" Tissue Pathology† Uninsured to Privately Insured
Esophagoscopy (42.23)	908	0.26‡
Gastroscopy (44.13)	2478	0.38‡
Proctosigmoidoscopy (48.23)	1120	0.87
Colonoscopy (45.23)	3781	0.52
Cystoscopy (57.32)	3320	0.34
Endoscopy of small intestine (45.13)	5408	0.51§
Bronchoscopy (33.22, 33.23)	2019	0.66
All above endoscopic procedures	18 834	0.64†

*Null hypothesis is that relative coefficients equal 1.00.

†Possible outcomes are not abnormal, abnormal—other mechanical abnormality, degeneration and necrosis, nonacute inflammation, acute inflammation, nonmalignant neoplasm, and malignant neoplasm.

‡.01 < $P \leq .05$.

§.05 < $P \leq .10$.

¶Includes a separate control variable for each endoscopic procedure.

† $P \leq .01$.

In all cases but one (white women 35 to 49 years old), the uninsured have a higher relative probability of in-hospital death. Moreover, for 10 of the 16 age-sex-race-specific cohorts, the difference is statistically significant ($P \leq .05$), with the relative probabilities ranging from 1.20 to 3.20. The relative probability of in-hospital death for uninsured patients compared with privately insured patients is greater for blacks than for

whites in seven of the eight age-sex cohorts.

COMMENT

Our analysis strongly suggests that an individual's condition on admission, use of resources during hospitalization, and likelihood of in-hospital death vary depending on whether the individual has health insurance. In the case of condition on admission, compared with the

Table 6.—In-Hospital Death: Regression-Adjusted Coefficients for Uninsured Relative to Privately Insured by Age-Sex-Race Cohort*

Age, Sex, Race	Unweighted N	Relative Regression-Adjusted Coefficient
Males		
Age 1-17 y		
White	38 527	1.46
Black	8813	1.88
Age 18-34 y		
White	75 514	1.34†
Black	17 614	1.92†
Age 35-49 y		
White	46 429	1.46†
Black	11 384	3.20†
Age 50-64 y		
White	64 417	1.23†
Black	9175	1.08‡
Females		
Age 1-17 y		
White	29 724	2.37†
Black	7098	2.58†
Age 18-34 y		
White	72 780	1.48§
Black	15 828	1.80
Age 35-49 y		
White	52 371	0.82
Black	11 336	1.60†
Age 50-64 y		
White	61 611	1.15§
Black	10 385	1.26‡

*Null hypothesis is that relative coefficient equals 1.00.

† $P \leq .01$.

‡.01 < $P \leq .05$.

§.05 < $P \leq .10$.

privately insured, the uninsured were more likely to be admitted for a condition that has a higher expected risk of death (based on the RAMI score) and appeared to be in more urgent need of care. The latter conclusion is suggested by our finding that, when differences in case mix between the uninsured and privately insured are controlled for with the Medicare case-mix index and the RAMI, the uninsured were significantly more likely to be admitted on a weekend.

Our findings regarding differences in resources expended on the hospitalized uninsured vs the hospitalized privately insured are more mixed, varying with the specific measure of resource use and the method of controlling for diagnostic mix that we employed. In our analysis of length of stay, which was based on hospitalizations for each of 10 diagnoses classified by the degree of physician discretion assumed to be involved in the decision to hospitalize, we found consistently shorter lengths of stay among the uninsured compared with the privately insured. These differences in length of stay were statistically significant for all five of the high-discretion/low-risk of death diagnoses, the category in which physicians' practices would be most likely to differ among insured vs uninsured patients, but for only one of the low-discretion/high-risk of death diagnoses.

We also found evidence of a difference

between the uninsured's and the privately insured's use of some, but not all, of the procedures we examined. For example, although we found a strikingly lower rate of performance of five high-cost and/or high-discretion therapeutic procedures (coronary artery bypass surgery, total knee replacement, total hip replacement, stapedectomy, and surgical correction of strabismus) among the uninsured compared with the privately insured in our entire sample of discharges, we found statistically significant evidence of a lower rate of use of only three (colonoscopy or proctosigmoidoscopy, upper gastrointestinal tract endoscopy, and coronary arteriography) of the seven diagnosis/diagnostic procedure pairs we examined. Our finding of a statistically significant difference in the uninsured's vs the privately insured's rate of use of some, but not all, procedures we examined and of a statistically significant difference in the uninsured's vs the privately insured's average length of stay for high- but not low-discretion hospitalizations suggests that if physicians are discriminating in the resources they expend on the uninsured and the privately insured, they may be doing so on a selective basis to minimize adverse effects on the health of the uninsured.

A similar suggestion emerges from our findings that privately insured patients were consistently more likely than uninsured patients to have completely normal tissue pathology findings on biopsy specimens obtained during any of seven different endoscopic procedures (Table 5). This latter finding strongly suggests that physicians who perform these procedures have a "lower threshold of suspicion" for performing biopsies on privately insured patients than they do on uninsured patients. One cannot infer from our data, however, that physicians' thresholds for performing a biopsy on privately insured patients are too low, or that their thresholds for performing a biopsy on uninsured patients are too high.

Finally, our analysis demonstrates that the uninsured have a higher relative risk of in-hospital death than the privately insured in 15 of 16 age-sex-race-specific cohorts, even when the Medicare case-mix index, the expected risk of death, and whether a procedure was performed are controlled for. In 10 of these 15 cohorts, the difference is statistically significant ($P < .05$). Although it is possible that this observed difference in in-hospital mortality is due to underprovision of needed medical services to hospitalized uninsured patients, the difference also could be due to differences in severity of illness be-

tween the uninsured and privately insured that are not reflected fully in the Medicare case-mix index and the RAMI. It is also possible that privately insured patients are more likely than uninsured patients to be discharged to another facility, such as a nursing home or a hospice, where death might occur shortly after discharge from the hospital.

Our results extend the findings of other studies in several ways. First, we have provided evidence that insurance coverage has an effect on resource use for a broad spectrum of clinical problems in addition to lung cancer,⁴ cardiovascular disease,⁵ and the acquired immunodeficiency syndrome.¹¹ Second, we have found evidence for differences between the uninsured and the privately insured in resource use for some, but not all, conditions, and of differences in use of some, but not all, types of resources. Although our findings of differences in average length of stay for high- but not for low-discretion hospitalizations and of a consistently higher probability of completely normal tissue pathology in biopsy specimens obtained from the privately insured vs the uninsured do not clearly indicate whether differences in the care provided to the uninsured vs the privately insured reflect "the unmet needs of some of our citizens or the overmet needs of others,"¹² they do provide empirical support of the hypothesis that insurance coverage may have a bigger impact on the use of elective or more highly discretionary services than on medically necessary care.¹³

Finally, we have found differences between the uninsured and the privately insured in a very important measure of outcome, namely, in-hospital mortality. Our finding of a higher in-hospital mortality rate among the uninsured, even after controlling for expected risk of death at the time of admission, suggests that we should not conclude that all differences between the care provided to the uninsured vs the privately insured are due to overuse of services by the privately insured.

Clearly, there is a need for further research comparing the care provided to the uninsured vs the privately insured that employs better measures of patients' condition at the time of admission and more refined measures of the process of care as well as measures of the outcomes of care other than simply in-hospital mortality. In addition, studies of the appropriateness of care should treat insurance coverage as an important covariate in assessing both cost and outcomes of the care of specific diagnoses.

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Health Insurance and Mortality

Evidence From a National Cohort

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Objective.—To examine the relationship between lacking health insurance and the risk of subsequent mortality.

Design.—Adults older than 25 years who reported they were uninsured or privately insured in the first National Health and Nutrition Examination Survey, a representative cohort of the US population, were followed prospectively from initial interview in 1971 through 1975 until 1987.

Participants.—Complete baseline and follow-up information was obtained on 4694 (91%) persons of the 5161 who reported not receiving publicly funded insurance at baseline.

Main Outcome Measure.—The relationship between insurance status and subsequent mortality was examined using Cox proportional hazards survival analysis. The analysis adjusted for gender, race, and baseline age, education, income, employment status, the presence of morbidity on examination, self-rated health, smoking status, leisure exercise, alcohol consumption, and obesity. The effects of interactions between insurance and all other baseline variables were also examined.

Results.—By the end of the follow-up period, 9.6% of the insured and 18.4% of the uninsured had died. After adjustment for all other baseline variables, the hazard ratio for lacking insurance was 1.25 (95% confidence interval [CI], 1.00 to 1.55). The effect of insurance on mortality was comparable to that of education, income, and self-rated health. There were no statistically significant ($P < .05$) interactions.

Conclusions.—Lacking health insurance is associated with an increased risk of subsequent mortality, an effect that is evident in all sociodemographic health insurance and mortality groups examined.

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THE NUMBER of uninsured persons has increased throughout the past decade,¹ reaching 35.4 million in 1991.² There are several reasons for this,³ including a changing economy and in-

creased costs of health care and insurance, making both less affordable both for employers and individuals. As a result, there has been widespread recent interest in ensuring health insurance for all Americans. There is limited system-

Medicaid coverage has been shown to reduce this differential.^{5,6,12,13} A lack of health insurance is also associated with a lower prevalence of recommended preventive services.^{14,15} Other studies have found that uninsured persons compared with the insured are more likely to have potentially avoidable hospitalizations,^{16,17} may be sicker at the time of hospital admission,^{18,19} are more likely to experience in-hospital^{20,21} and cancer mortality,²² and are less likely to receive invasive procedures.^{12,23} Many of these studies suggesting adverse outcomes do not include uninsured persons who have not entered the health care system, so that the results may simply represent adverse selection due to delayed access to care.

Some studies do point to health hazards associated with lacking health insurance. Braveman et al²⁴ studied hospital discharge data in California and found that adverse outcomes in newborns were more frequent in the uninsured. The use of hospital discharge data limited the ability of the authors to adjust for potential confounders. Some studies have shown that persons losing health insurance benefits suffer measurable declines in their health.²⁵⁻²⁸ In the Rand Health Insurance Experiment, on average, persons randomized to both the health insurance copayment group and the free care group ended the study with similar levels of health.²⁹ For many

tus among 94 poor patients with financial barriers to care compared with 94 control patients without financial barriers. Patrick et al³¹ found that poor persons in families with all members uninsured compared with those from insured families had lower self-rated health and greater perceived need of services. Most studies have focused on the poor with the implication that the benefits from health insurance are limited to poor persons. Lack of insurance is correlated with low income, which is itself associated with poor health,³²⁻³⁶ and most previous studies do not permit isolation of the effects of being poor from that of being uninsured. Many of the studies are cross-sectional in design, limiting confidence in the direction of the causal pathway. Finally, none of these studies have addressed the relationship between insurance and mortality.

Because of the limitations of studies addressing this important policy issue we examined the relationship between health insurance and mortality, using data from the National Health and Nutrition Examination Survey Epidemiologic Follow-up Study (NHEFS), which followed a representative cohort of the US population for up to 16 years. The rationale and analysis used in the present study was informed by the conceptual framework developed in a US Congress Office of Technology Assessment report.⁴ The availability of health insurance is viewed as an enabling factor facilitating access to care. As noted above, persons lacking insurance have reduced access to care, evidenced by reduced use of a wide range of health services and self-reported forgone and delayed care.^{15,17,11,14,16} In addition, the process of care for uninsured persons may be inferior. For example, the uninsured have an increased risk of suffering medical injury due to substandard care.³⁶ Consequent on reduced access and lower quality of care, we hypothesized that persons lacking insurance would experience increased mortality. Our analysis adjusted for the potentially confounding effect of other factors identified in the Office of Technology Assessment report as affecting health. These factors are categorized as predisposing (age, gender, education, employment, and race), need (perceived health and objective evidence of morbidity), other enabling (income), and individual behaviors (smoking status, alcohol consumption, leisure exercise, and obesity status).

METHODS

The first National Health and Nutrition Examination Survey (NHANES I), conducted between 1971 and 1975, col-

lected sociodemographic, health insurance and utilization, medical history, and clinical and laboratory information from several national probability samples of the civilian noninstitutionalized population.^{37,38} Not all information was collected on all respondents. Detailed information including insurance status was collected on 6913 adults aged 25 to 74 years. Interviewees were also examined by physicians, who assigned up to 15 diagnoses based on history, physical examination, and the results of laboratory investigations. The NHEFS was designed to trace and reinterview respondents aged 25 to 74 years.³⁹ Follow-up surveys in NHEFS were conducted in 1982 through 1984, 1986, and 1987. Data collected comprised interview surveys, medical records from health care facilities, and death certificates for all decedents. The age, race, and sex-specific mortality of the NHEFS cohort is similar to that experienced by the US population.⁴⁰

We analyzed data on adults who reported in NHANES that they did not receive publicly funded insurance. We excluded adults with Medicaid, Veterans Administration insurance, or Medicare to avoid confounding by poor adults whose health status may be compromised by the presence of significant disabilities.⁴¹ Of the 6913 persons asked about their insurance status, 5218 reported that they were either uninsured or had private insurance. Vital status at follow-up was available on 4939 persons (95%). Compared with persons with vital status follow-up information, those unavailable for follow-up were younger and were more likely to be black men and white women.⁴² We also excluded 57 persons whose race (mostly Asian and American Indian) was neither white nor black, because their number was too small to allow reliable analysis. The univariate analysis presented was thus based on a sample of 4882 persons. There were 188 persons (3.8%) with vital status information at follow-up who had incomplete baseline data, mostly missing family income data, so that the multivariate analyses were based on 4694 persons. Compared with persons with complete baseline data, those with incomplete baseline data were more likely to be uninsured (19.9% compared with 14.1%), older (61.8% compared with 52% were over 44 years of age for those with complete data), and have less than 12 years of school (45.7% compared with 34%).

Analyses

The NHANES I used multistage stratified probability samples of clusters of persons. In addition, persons liv-

ing in poverty areas, women of child-bearing age, and elderly persons were oversampled. To accommodate the complex survey design the statistical package SUDAAN⁴³ was used in the analyses reported below. The SUDAAN program uses a Taylor series approximation method to compute variances that allow adjustment for the multistage probability sampling strategy. The revised weights provided on the 1987 NHEFS public use tapes were used to adjust for survey oversampling and non-response to yield population estimates of reported baseline descriptors. The univariate relationships between each baseline variable and health insurance status (dichotomized as reporting having no health insurance or having private health insurance) and mortality were examined using χ^2 tests. To examine the relationship between insurance status and subsequent mortality, proportional hazards survival analyses were used to adjust for other baseline variables, including the following: age (years); gender; race (dichotomized as white or black); education (dichotomized as at least 12 years of school or less); family income at baseline (treated as three dummy variables, income less than \$7000 per year, \$7000 to \$9999 per year, and \$10 000 to \$14 999 per year, using income of \$15 000 and higher per year as a reference); employment status (dichotomized as working most of the previous 3 months or not); morbidity (dichotomized as the presence or absence of evidence of morbidity on medical examination and laboratory testing); self-rated health (treated as three dummy variables, reporting health in general to be very good, good, or fair/poor with excellent as the reference value); smoking status (smoker or not); obesity status (dichotomized as body mass index [weight in kilograms divided by height in meters squared] >27 or not); leisure exercise (dichotomized as reporting little or no exercise compared with moderate or much); and alcohol consumption (dichotomized as consuming at least six alcoholic drinks per week or less). Because the family income variables were not adjusted for household size, the survival analyses also included household size as an interval level variable. In addition to a survival analysis including only main effects, a model with the interaction terms between insurance status and the other independent variables was examined. Interaction terms were retained if they were statistically significant ($P < .05$). To avoid the inefficiency of performing weighted multivariate analyses, we followed the recommendations of Korn and Graubard⁴⁴ to unweighted survival analyses that

Table 1.—Relationships Between Health Insurance Status and Mortality and Selected Characteristics

Characteristic	No. (%)	% Uninsured* (SE)	% Died† (SE)
Vital status			
Dead	593 (10.7)	22.1 (2.4)	...
Alive	4289 (89.3)	11.7 (0.7)	...
Insurance status			
Uninsured	699 (12.8)	...	18.4 (2.0)
Insured	4183 (87.2)	...	9.6 (0.5)
Age group, y			
≥55	1208 (28.8)	17.4 (3.9)	27.1 (1.6)
45-54	1354 (24.9)	11.8 (2.0)	10.9 (0.8)
35-44	1035 (24.6)	10.1 (2.4)	6.7 (0.8)
25-34	1287 (22.8)	13.0 (3.7)	2.9 (0.5)
Sex			
Male	2136 (46.6)	13.8 (1.0)	13.7 (0.8)
Female	2748 (53.4)	11.7 (0.9)	6.1 (0.5)
Race			
Black	617 (9.1)	23.4 (2.5)	14.7 (1.9)
White	4365 (90.9)	11.8 (0.7)	10.3 (0.5)
Education, y			
<12	1694 (31.6)	22.6 (1.4)	18.7 (1.2)
≥12	3108 (68.5)	8.4 (0.6)	7.0 (0.4)
Income, \$			
<7000	950 (18.0)	34.7 (2.2)	19.2 (1.8)
7000-9999	1074 (25.1)	12.9 (1.1)	11.7 (1.0)
10 000-14 999	1283 (27.6)	8.4 (0.7)	7.8 (0.8)
≥15 000	1391 (29.3)	4.6 (0.8)	6.7 (0.7)
Employment status			
Unemployed	1796 (33.6)	20.3 (1.2)	12.0 (1.0)
Employed	3078 (66.4)	9.1 (0.7)	9.6 (0.5)
Self-rated health			
Fair or poor	825 (18.9)	27.6 (1.7)	24.7 (1.6)
Good	1332 (31.4)	13.2 (1.2)	12.4 (0.9)
Very good	1292 (26.5)	12.8 (1.0)	9.7 (0.9)
Excellent	1229 (25.2)	8.4 (0.8)	5.9 (0.6)
Morbidity			
Present	2822 (58.6)	15.7 (1.0)	14.9 (0.9)
Absent	1980 (43.4)	9.1 (1.0)	5.3 (0.6)
Leisure exercise			
Little or none	1865 (39.7)	16.0 (1.1)	12.8 (0.9)
More	2993 (60.3)	11.3 (0.8)	9.4 (0.6)
Smoking status			
Smoker	1968 (42.0)	14.3 (1.0)	13.4 (0.9)
Nonsmoker	2914 (58.0)	11.8 (0.9)	8.8 (0.5)
Alcohol consumption			
≥6 drinks per wk	1299 (27.0)	11.9 (1.3)	13.4 (1.0)
<6	3613 (73.0)	13.2 (0.8)	9.7 (0.6)
Obesity status			
BMI, >27	467 (9.3)	16.9 (1.8)	14.4 (1.7)
BMI, ≤27	4415 (90.7)	12.5 (0.7)	10.3 (0.5)

*Percentage of those with the characteristic who were uninsured at baseline.

†Percentage of those with the characteristic who died during the follow-up period. Percentages are weighted to provide population estimates, and SEs are adjusted for cluster sampling strategy.

BMI indicates body mass index (weight in kilograms divided by height in meters squared).

controlled for the variables used in determining the sample weights (age, gender, race, and income).

RESULTS

Baseline characteristics of the sample are shown in Table 1, which also provides population estimates for each characteristic. Not having health insurance was reported by 699 persons (12.8% of the population). Persons older than 55 years were more likely to be uninsured (χ^2 , 26.1; $P<.001$), as were men (χ^2 , 33.8;

$P<.001$), blacks (χ^2 , 16.7; $P<.001$), those with less than 12 years of school (χ^2 , 73.5; $P<.001$), those with lower family incomes (χ^2 , 114.8; $P<.001$), the unemployed (χ^2 , 78.8; $P<.001$), those reporting lower self-rated health (χ^2 , 59.9; $P<.001$), those with morbidity found on medical examination (χ^2 , 18.4; $P<.001$), and those reporting little or no leisure exercise (χ^2 , 10.1; $P=.002$). There was no statistically significant ($P<.05$) relationship between insurance and smoking status, alcohol consumption, or obesity.

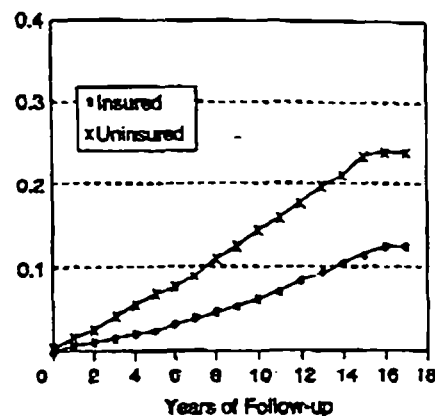


Fig 1.—Cumulative mortality probability by baseline insurance category.

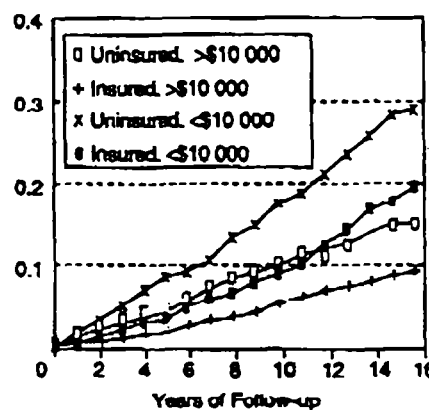


Fig 2.—Cumulative mortality probability by baseline insurance and income categories.

By the end of the follow-up period, 593 persons, representing 10.7% of the population, had died (Table 1). Those who died were nearly twice as likely to have been uninsured at baseline compared with those who survived (χ^2 , 16.5; $P<.001$). Increased mortality was also associated with being older (χ^2 , 228.8; $P<.001$), being a male (χ^2 , 33.8; $P<.001$), black (χ^2 , 8.9; $P<.001$), having less than 12 years of school (χ^2 , 68.7; $P<.001$), lower incomes (χ^2 , 49.4; $P<.001$), unemployment (χ^2 , 9.8; $P=.003$), lower self-rated health (χ^2 , 73.2; $P<.001$), morbidity present on baseline medical examination (χ^2 , 66.4; $P<.001$), reporting little or no leisure exercise (χ^2 , 8.9; $P=.004$), smoking (χ^2 , 20.1; $P<.001$), consuming six or more drinks per week (χ^2 , 11.0; $P=.001$), and obesity (χ^2 , 4.9; $P=.06$). Figures 1 through 3 show the relationship between insurance status at baseline and the probability of mortality over the follow-up period, both unadjusted and stratified by income and morbidity. The adverse association between lacking insurance and mortality was observed in all subgroups.

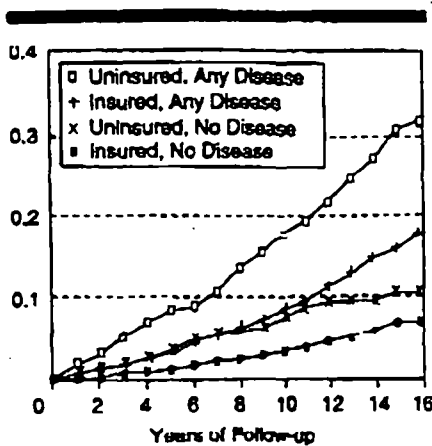


Fig 3.—Cumulative mortality probability by baseline insurance and morbidity categories.

After adjusting for all other baseline characteristics, the proportional hazards survival analysis revealed that lacking insurance at baseline was associated with an increased risk of mortality (hazard ratio, 1.25; 95% confidence interval [CI], 1.00 to 1.55; Table 2). None of the interactions between insurance status and the other baseline characteristics made statistically significant ($P < .05$) contributions to the regression model. When the employment, self-rated health, and morbidity variables were excluded, the adjusted hazard ratio for lacking insurance was 1.34 (95% CI, 1.07 to 1.69).

COMMENT

This analysis of a nationally representative cohort of the US population suggests that the mortality experience of Americans without health insurance is greater than those with insurance. The effect observed is comparable to that observed for education, income, employment status, and self-rated health. In this study and reported in previous studies, lack of insurance is associated with social and medical factors that increase the risk of poor health.²³⁻²⁵ However, lacking insurance is associated with subsequent higher mortality independent of other risk factors. The relationship was observed in those with and without baseline morbidity and in those with higher and lower levels of income and education. The absence of statistically significant interactions in the survival analysis between health insurance and the other baseline characteristics suggests that the benefits of health insurance are not confined to particular subgroups. These results are consistent with a previous study of a nationally representative sample suggesting a beneficial association between insurance and health status in persons both below and above 200% of the federal poverty level.⁴⁵

Table 2.—Proportional Hazards Model for Survival Time Adjusted for Baseline Characteristics (N=4694)*

Baseline Risk Factor	Hazard Ratio (95% Confidence Interval)
No insurance	1.25 (1.00-1.55)
Age	1.07 (1.09-1.05)
Men	1.90 (2.40-1.50)
Black	1.30 (1.72-1.06)
<12 y school	1.32 (1.61-1.09)
Income	
<\$7000	1.26 (1.68-0.98)
\$7000-\$9999	1.27 (1.57-0.97)
\$10 000-\$14 999	1.08 (1.48-0.79)
Unemployment	1.21 (1.59-0.98)
Morbidity present	1.43 (1.74-1.18)
Self-rated health	
Fair/poor	1.62 (2.21-1.18)
Good	1.25 (1.61-0.97)
Very good	1.21 (1.58-0.92)
Little or no exercise	1.14 (1.38-0.96)
Present smoker	1.84 (2.15-1.57)
BMI, >27†	1.40 (1.74-1.13)
≥6 drinks per wk	1.46 (1.78-1.20)

*Analysis also adjusted for household size. Baseline risk factors indicate the value of the baseline characteristics associated with higher mortality. Except where noted, the hazard ratio shows the adjusted hazard with the risk factor present compared with the risk factor absent; for age, the risk factor is a 1-year increment in age; each income group is compared with the income \$15 000 or more reference group; each self-rated health group is compared with excellent self-rated health as the reference group.

†BMI indicates body mass index (weight in kilograms divided by height in meters squared).

Two main causal pathways may explain the results observed in this prospective study. First, the relationship observed between lacking health insurance and increased mortality may be the result of both insurance and mortality being associated with a third unmeasured underlying variable. In particular, persons without health insurance may value health less than those with insurance. Consequently, persons placing less value on health may fail to get insurance and also experience higher mortality because of individual lifestyle factors. There is little evidence to support this hypothesis. Studies on the factors associated with having insurance suggest that most persons lack insurance because they cannot afford it rather than because they are unwilling to buy it.⁴ In the 1987 National Medical Expenditure Survey, 76.9% of uninsured persons were in families with adult workers, and lacking insurance was associated primarily with employment factors such as industry type, size of establishment, and hourly wage.⁴⁶ The present study controlled for several health behaviors that are associated with mortality. These lifestyle factors exhibited only modest associations with the availability of insurance.

Second, the results are consistent with the study hypothesis that a lack of health insurance is causally related to a higher mortality rate, because of decreased access and lower quality of care. This hypothesis is in accordance with the re-

sults of previous studies, and the conclusions of the US Congress Office of Technology Assessment report.⁴ For the most part, previous studies suggesting the benefits of health insurance have either focused on the poor^{27,28,31} or suggested that the benefits of free health care are evident mostly in the poor.³⁹ Although poorer persons had a higher mortality rate, there was no evidence that the adverse effects of lacking insurance was limited to this group. Our analysis compared any level of insurance with no insurance and included a representative sample of adults not covered by publicly funded insurance programs; thus, the analysis may have been more sensitive to the effect of insurance than the Rand Health Insurance Experiment.²⁹

The results obtained probably underestimate the relationship between insurance and mortality. First, many persons in the insured group may be underinsured. Farley found that over a quarter of insured persons under 65 years of age were inadequately protected against the possibility of large medical bills.⁴⁷ Second, insurance status was measured only once, at baseline, and no account could be taken of the impact of changes in insurance status over time. Migration of persons with and without insurance to the other group would tend to bias the observed association toward zero. Most persons older than 50 years at the beginning of the study would have been eligible for Medicare by the end of the follow-up period. Significant shifts in insurance status also occur during the course of a year. In the 1987 National Medical Expenditure Survey, 22.4% of persons younger than 65 years were uninsured at some time during the year, 16.1% to 17.2% were uninsured at any one period during the year, and 11.4% were uninsured throughout the year. Third, the variables included in the survival analysis may have resulted in over-adjusting of the true relationship between insurance status and mortality. Over time, employment status may be a better measure of insurance than self-reported insurance status, since most persons obtain their insurance through employment.⁴⁸ Also, if lacking insurance adversely affects health, then adjusting for baseline health status will tend to reduce the observed relationship between health insurance and mortality. When the baseline employment and health status variables were excluded from the analysis, the relationship between insurance and mortality was increased. Although observational studies cannot exclude the possibility that unmeasured underlying variables exist, the relationship between lacking

health insurance and adverse health outcomes, it is unlikely that a randomized trial will be conducted to address this important policy issue. Our analysis, consistent with previous studies suggest-

ing an adverse impact of lacking insurance on health, provides evidence that lacking insurance also increases mortality in a nationally representative sample of adults. The existing "safety net"

has failed to prevent this increased mortality. The findings support a policy imperative for universal health insurance to reduce both financial barriers to care and the risk of premature mortality.

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To: Hillary Rodham Clinton
Fr: Bob Boorstin and Jason Solomon
Date: January 12, 1994

Re: Health Affairs Introduction

Uwe Reinhardt, who contributed to and helped put together the next issue of Health Affairs, asked Chris Jennings if you might be able to write an introduction to the issue, which contains articles almost all about the Clinton plan. Reinhardt suggested that your introduction remind the academics who tend to read this stuff that people -- not theories and statistics -- are at the center of this debate. He sent along ideas about what your introduction might say. (see attached) And we have attached a draft.

Reinhardt also co-authored an introduction with the editor of the journal. (see attached) As you know, Reinhardt has been very supportive recently, and an introduction by you is not only a way to reach "opinion leaders" on health policy but also to thank him for his continued support.

Included in the issue will be articles from Administration officials such as Walter Zelman, Alice Rivlin, Rick Kronick and Bernie Arons about our proposal. The Table of Contents is attached.

If you could review this draft and give us your comments, that would be great. Thanks.

Introduction

By Hillary Rodham Clinton

Together, we stand at a unique moment in history. In the coming months, we have an opportunity to accomplish what our nation has never done before: provide health security to every American.

Countless experts and academics have thought through the issue of how to control health costs while guaranteeing security to every American. Some of this thinking is reflected in this issue. At the beginning of this Administration, I brought together a Task Force of over 500 policy experts, physicians and other health professionals to consider this challenge. The result of their work -- and our consultation with members of Congress, representatives of different groups with a stake in the debate, and many of you -- is the Health Security Act of 1993.

The articles that follow lay out the dimensions of the crisis that confronts our nation, explain our approach and competing approaches, and state the case for comprehensive reform as proposed in the Health Security Act.

I invite you to read the articles and continue to add meaningful contributions to the final debate. But as we go forward, we must remember that this debate is about more than competing policy options. It is, most fundamentally, about people -- hard-working people who deserve to know that their health care will never be taken away.

Over the past months, I have had the extraordinary opportunity of listening to thousands of Americans talk about health care. I've sat in living rooms talking to farm families. I've stood on loading docks talking to people who have worked for 10, 15, and even

20 years without insurance. I've visited hospitals, talking to doctors and nurses. I have been moved by stories of parents who cannot afford a prescription for a child who is sick and hurting, of families barely hanging on financially and emotionally because of a serious illness. These people would be surprised to learn that some think everything is just fine with our health care system.

I have carried their stories in my mind as we worked long and hard to devise solid answers to tough questions. The President's Health Security Act is a product of all the people who took the time to share their ideas, their research, and their personal experiences with us.

The concerns that were expressed again and again - from those who need care and those who give care - convinced me of one point: although America can still proudly boast the world's finest health professionals and astounding medical advances, our health care system is broken. And if we go on without change, the consequences will be devastating for millions of Americans and disastrous for the nation in human and economic terms.

As experts in health policy, I encourage you to stay involved in the debate and scrutinize the more technical details of the proposal. As you know, the American people need your help in understanding the complex issues and difficult choices that lie behind the design of any comprehensive reform effort.

But I also urge you not to lose sight of the central goal of reform: giving every American the peace of mind that their health care will be there when they need it. Watered-down legislation that does not guarantee comprehensive benefits that can never be taken away is mere tinkering, not real reform. (Note: The previous two paragraphs were adopted from Reinhardt's memo)

As an American concerned about the health of our nation, I stand with you as we confront this challenge that touches all of us. I am convinced that we can achieve lasting, meaningful change -- and with your help, we will.

MEMORANDUM

To: Chris Jennings, The White House
From: Uwe E. Reinhardt, Princeton University
Subject: Intro piece by the First Lady
Date: January 12, 1993

Dear Chris:

The truly important message the First Lady might convey in a foreword or intro piece would be to drive home, once more, the following points which seem to get lost, from time to time, in our nitpicking over technical points.

First, the notion that there is "no health crisis" now being marketed by some would be news to the many Americans who either not have access to proper health care or, if they do, who find themselves near or in bankruptcy after a bout of serious illness.

- o In an article published by Jack Hadley, Jude Feder et al. in JAMA, they showed pretty clearly that, after adjustment for all other factors that impact on morbidity, uninsured Americans die at up to three times the rate from identical illnesses than do similarly situated but insured Americans. Recently, another study of this sort was published in JAMA. I attach a write up. In the face of such statistics, it is cruel and irresponsible to argue that lack of health insurance does not imply lack of health care. Americans must recognize that our health system does ration life years and quality life on the basis of income and price for many low-income Americans. If we want to do this fine; let's acknowledge it. People who deny it, however, we are fit for damnation.
- o Even if the uninsured did have access to proper health care, there remains the problem that they would be hounded by medical bills. It may not be well known that medical bills are one of the most prominent sources of personal bankruptcy in the United States. We are truly unique in that respect in the developed world. *

Second, the argument that most Americans are well insured by the private sector is bogus. It, too, will be news to people who have recently been laid off. Americans not on government programs such as Medicare or Medicaid are not "well insured;" they are "unsured" in the sense that they lose that insurance when they lose their jobs. The notion that COBRA 89 takes care of the problem, because an unemployed can purchase his/her private coverage from the old company at the old group rate *but with the unemployed's own money* is as cruel as it is ridiculous. It is Marie-Antoinetteism at its finest, for it reminds one of Marie Antoinette cute remark "Let them eat cake!" As more and more Americans are being restructured out of their jobs, they need truly portable insurance that is also affordable to them during spells of unemployment.

The First Lady should remind us economists (and others like us) that it is honorable and good to debate the many technical details of the proposal, but that we must never lose sight of the central thrust of the President's endeavor: to give every Americans the peace of mind that proper health care will be available in the hour of need and that the family will not go broke as a result of illness. Technical changes that will deflect us from this goal--e.g., a shift to a form of financing that does not guarantee portable universal coverage--is simply not acceptable and should not be acceptable to economists either. That is tinkering, not health reform. The fundamental question Americans should ask themselves is one I raised at Annapolis: who should bear the risk for our possible failure to control health-care costs: Americans in general, or the uninsured, among them working mothers? It is astounding how many well paid, well insured men would elect the working mother to assume that risk taking. And that in John Wayne country! It is time we move from celluloid decency to real life decency. }

From the Editor

The Policymakers' Dilemma

One need look no further than America's previous failures to reform the health system to recognize that it is a complex undertaking. The challenge is daunting because it must attract broad political support in a nation that has never achieved consensus on an overriding social ethic (universal coverage) to which all other worthwhile goals in health care must take second place. Americans generally profess allegiance to the precept that the highly sophisticated health care of which this nation is justly proud should be available to all citizens, regardless of their ability to pay for it. At the same time, however, broad segments of the public are deeply suspicious of the only stakeholder that can force a truly egalitarian distribution of health care on our highly entrepreneurial delivery system: the government. The result of this ambivalence—open espousal of a very lofty goal, but open hostility to the only means of achieving that goal—has led to the perennial stalemate that is the hallmark of congressional wrangling over universal health insurance. President Bill Clinton's reform proposal can be seen as a bold attempt to shake this nation out of stalemate, literally by forcing us into a vigorous national debate. The plan itself is a highly complex architecture that seeks to bridge the divergent ideologies that drive this nation's debate on health policy. To that end, the plan seeks to enlist the power of market forces, but it attempts to channel these forces, by means of government regulation, toward the goal of portable, universal coverage. Coverage would be financed by a complicated mix of sources in as equitable a manner as the body politic is thought (by the plan's architects) to tolerate. In contemplating these regulations, one must keep in mind that any market system can flourish only within a regulatory and legal framework that makes competition for economic privilege civilized. After all, competition in this context is merely a polite word for *brawl*. Absent the supervision and regulatory powers of a governmental infrastructure, "free markets" almost always degenerate into the untoward and often violent free-for-all now unfolding in the disintegrating former Soviet empire. One of the common themes struck by many of the contributions in this volume is a salute to the Clintons for giving health policy, at long last,

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EDITOR 2

the presidential priority that a sector ~~that~~^{ina} represents one-seventh of our economy deserves. As Congress works its will, the final product will likely resemble the president's initial vision only modestly. What might well be enacted is a version of the president's plan that slouches heavily toward its close cousin—namely, the plan recently introduced by two Republicans—Senator John Chafee of Rhode Island and Representative William Thomas of California. Like the president's plan, the Chafee-Thomas bill features as the chief workhorse for cost control the mechanism of "managed" (really regulated) competition among vertically integrated networks of health care providers. These providers would somehow divide among them a prepaid annual capitation payment per insured family and control their own costs internally through various means, all lumped together under the rubric of "managed" care. Just how such a system would be financed, however, remains the \$64,000 question whose resolution will tax Congress's ingenuity and sorely test its commitment to the goal of universal coverage. The president has proposed a particular financing scheme, centered on an employer mandate and supplemented by a variety of sources, among them some modest sin tax increases and reductions in the growth of Medicare and Medicaid expenditures. While it is easy to question the employer mandate on the basis of pure economic principle, it is not clear whether the typical American voter would greet with equanimity the full implications to him or her of the alternative, the so-called individual mandate. It is surely a matter to which the proponents of the individual mandate must give more thought. Caught between the economic costs inherent in an employer mandate, on the one hand, and the political costs inherent in the individual mandate, on the other, Congress might, of course, retreat once more into its traditional corner of respite: further postponing universal coverage until that hypothetical better day when, it is hoped, health care costs will be under "better" control. For all intents and purposes, such a strategy at this time would spell abandonment of universal coverage as this nation's goal. Another failure in this regard would only serve to once again hold up the United States as the major outlier among civilized nations, all of which provide their citizens with insurance protection against the unpredictable financial consequences of illness. Not one author presenting at the Princeton workshop, nor one discussant who responded, appeared to favor a strategy of further delay.

Uwe E. Reinhardt (Princeton University)

John K. Iglehart (Health Affairs)

side-by-side signatures to sign-off on

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-
- | | |
|-------|---|
| 7 | Rationale behind the Clinton Health Reform Plan
<i>Walter A. Zelman and Ira Magaziner</i> |
| <hr/> | |
| 28 | Financing, Estimation, and Economic Effects
<i>Alice M. Rivlin, David M. Cutler, and Len M. Nichols</i> |
| <hr/> | |
| 48 | Perspective: Alternative Estimate: No Pain, No Gain
<i>John F. Sheils and Lawrence S. Lewin</i> |
| <hr/> | |
| 54 | Economic Analyses of the Clinton Plan: |
| 55 | Sowing the Seeds of Reform in 1994
<i>Henry J. Aaron</i> |
| <hr/> | |
| 67 | Health System Reform: Let's Not Miss Our Chance
<i>Stuart Altman</i> |
| <hr/> | |
| 79 | A Single-Payer System In Jackson Hole Clothing
<i>Alain C. Enthoven and Sara J. Singer</i> |
| <hr/> | |
| 94 | Perspective: A Helping Hand for the Invisible Hand
<i>Richard Kronick</i> |
| <hr/> | |
| 100 | The Clinton Plan: A Researcher Examines Reform
<i>Victor R. Fuchs</i> |
| <hr/> | |
| 113 | The Good, The Bad, And The Bottom Line
<i>Jack Hadley and Stephen Zuckerman</i> |
| <hr/> | |
| 130 | Patients At Risk: Reform And Risk Adjustment
<i>Joseph P. Newhouse</i> |
| <hr/> | |
| 145 | What Happened To The Tough Choices?
<i>Mark V. Pauly</i> |
| <hr/> | |
| 159 | The Clinton Plan: A Salute to American Pluralism
<i>Uwe E. Reinhardt</i> |
| <hr/> | |
| 177 | Health Reform: What Will It Take To Pass?
<i>Gail R. Wilensky</i> |
| <hr/> | |
| 190 | Mental Health and Substance Abuse Coverage
<i>Bernard S. Arons, Richard G. Frank, Howard H. Goldman, Thomas G. McGuire, and Sharman Stephens</i> |
| <hr/> | |
| 204 | Antitrust, Competition, and Health Care Reform
<i>Robert E. Bloch and Donald M. Falk</i> |
| <hr/> | |
| 222 | Eliminating Waste and Inefficiency Can Do Little to Contain Costs
<i>William B. Schwartz and Daniel N. Mendelson</i> |
-

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-
- 237 Managed Care Industry Perspectives:
Why Do We Need Health Alliances?
Robert E. Patricelli
- 241 The Case for Local Budgeting
Jon Glaudemans
- 245 Concerns of an HMO Executive
Roger F. Greaves
- 246 A Workable Framework for Health Reform
James A. Lane
- 248 Quality Measurement and Improvement
William L. Roper
-
- 250 Commentary:
The Vital Role of Professionalism in Health Care Reform
David Blumenthal
- 255 Regional Professional Foundations
John E. Wennberg and Robert Keller
- 262 Pragmatism and Health Care Reform
Donald R. Cohodes
-
- 272 DataWatch:
The Beliefs and Values Shaping Today's Health Reform Debate
Robert J. Blendon, John Martula, John M. Benson, Matthew C. Shelter, Francis J. Connolly, and Tom Kiley
- 283 Public Opinion's Tilt against Private Enterprise
Lawrence R. Jacobs and Robert Y. Shapiro
- 297 Perspective: Should We Rely on Polls?
Marc L. Berk
- 299 Implicit Pooling of Workers from Large and Small Firms
Alan C. Monheit and Jessica Primoff Vismes
- 313 Employer-Based Health Insurance in a Changing Work Force
Deborah Choller
- 325 The Health Insurance Picture in 1993: Some Rare Good News
Jon Gabel, Derek Liston, Gail Jensen, and Jill Marsteller
- 335 Paying for Mental Health and Substance Abuse Care
Richard G. Frank, Thomas G. McGuire, Darrel A. Regier, Ronald Manderscheid, and Albert Woodward
-
- 331 UpDate:
State Report
Will Employer Mandates Really Work? Another Look at Hawaii
Andrew W. Dick
- 338 State Report

HA Winter 1993 SLUG: Contents
Repra — 3 December 1993 Page 4

Small-Employer Health Alliance in California
Emery B. Dowell and Thomas R. Oliver

- 340 Book Review:
*Free for All? Lessons from the RAND Health Insurance Experiment (by
Joseph P. Newhouse and The Insurance Experiment Group)*
Reviewed by Deborah A. Freund
-

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(ATTN: National editors) (Includes optional trims)

Study Finds Race, Income Affects Health Care Quality (Santa Monica)

By Douglas P. Shuit= (c) 1994, Los Angeles Times=

SANTA MONICA, Calif. Adding to the evidence showing that race and income may determine the quality of health care Americans receive, a new study shows that black and poor patients with Medicare cards receive worse hospital care than other acutely ill elderly patients.

The study, to be published Wednesday in the Journal of the American Medical Association, analyzed nearly 10,000 patient records at 297 hospitals in five states. The researchers at RAND, the Santa Monica think tank, found that the black and low-income elderly fared worse than other Medicare patients whether at small rural hospitals or urban teaching hospitals.

Only 47 percent of the black and poor Medicare patients considered to be seriously ill when admitted were put into intensive care units, compared to 70 percent of the more affluent Medicare patients, the study found.

The researchers also said that nurses more frequently checked for breathing difficulties in new pneumonia patients who were affluent than those who were poor and black, and that doctors were less likely to find out what kinds of medicines the poor and black patients were using prior to hospitalization.

Among the other findings of the study, the researchers said that 19 percent of black and poor patients were discharged from hospitals in unstable condition, meaning they were still suffering from at least one serious health problem, compared to 14 percent of the rest of the group. At the same time, the poorer care received by black and low-income patients tended to be minimized since they were nearly twice as likely to be treated at higher-quality urban teaching hospitals.

The researchers said these teaching hospitals, usually located in urban centers, were so extraordinary that they often compensated for the poorer quality of care given to the patients who were black or from low-income neighborhoods.

"Because these patients receive more of their care in better-quality hospitals, there are no overall differences in quality by race and poverty status," the authors said while taking care to point out this was not the case within individual hospitals.

The study found that death rates were about the same for all the patients surveyed. That finding runs counter to several earlier studies which found that blacks and the poor have higher mortality rates for certain diseases and surprised the researchers.

Dr. Katherine Kahn, the study's lead author, said it was possible that her study, even though it examined 9,932 Medicare patients, may not have been broad enough to develop meaningful data on death rates.

The survey has some clear implications for national health care reform, since much of that debate is being driven by a belief that giving everyone a health insurance card will solve many of the nation's health care ills.

(Optional add end)

The study, funded by two federal health agencies, clearly shows that simply possessing a Medicare card does not create an equal health status in the nation's hospitals.

"The message is that health reformers have to pay attention to improving the quality of care blacks and the poor receive as well as to expanding their access," said Kahn, an internist who teaches at the the University of California, Los Angeles Medical School in addition to her research work at RAND.

The hospitals included in the study were not identified. The 297 hospitals were located in 30 urban and rural areas in five states, and comprised a nationally representative database, Kahn said. The hospitals were located in California, Pennsylvania, Texas, Florida and Indiana.

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