The Systematic Assessment of Variations in Medical Practices and their Outcomes

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The staff of Ed Fu Associates, Arlington, VA, carried out the analyses leading to the publication of the "Medicare Hospital Mortality Information," the annual Health Care Financing Administration compilation on which this study was partially based.

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Synopsis

The Health Care Financing Administration of the Department of Health and Human Services has carried out for several years the systematic assessment of variations over time and among geographic locales in patterns of care and patterns of outcomes experienced by Medicare beneficiaries. This routine monitoring focuses principally on hospitalizations and their outcomes (death and readmission) and is based on the Medicare enrollment file and the claims file for inpatient care.

The period 1985–88 has been marked by declining adjusted post-admission risks for mortality (down 4 percent) and readmission (down 6 percent) for Medicare beneficiaries. The downward trend in mortality risks is most evident following hospitalizations for acute myocardial infarction (down 8 percent) and stroke (down 12 percent).

Hospital admission and population mortality rates, adjusted for differences in demographic and socioeconomic characteristics of the populations, vary substantially among areas as large as States and Metropolitan Statistical Areas, as do risk-adjusted post admission probabilities of death among those areas and among hospitals. Thus, if overall admission and mortality rates in the upper three quartiles of Metropolitan Statistical Areas were brought down to the average of the lowest quartile, there would be 20 percent fewer admissions and 12 percent fewer deaths within 180 days of admission for hospitalized patients.

Although favorable trends in the effectiveness of the hospital care received by Medicare beneficiaries appear discernible, the existence of substantial variations suggests that further improvement may be possible.

WIDE VARIATIONS HAVE BEEN NOTED repeatedly in the frequencies with which specific surgical and medical interventions are undertaken in small (*1*) and large (*2–3*) geographic areas. These variations cannot be accounted for by differences in need. They are more probably due to the well-documented (*4–8*) disparities among physicians in their opinion about the utility of interventions. Indeed, the point has been made quite emphatically that the underlying cause is a lack of evidence to substantiate the effectiveness of the great majority of commonly used interventions (*9*).

Decisions to treat, therefore, currently tend to be driven by individually held expectations of a potentially favorable biologic effect rather than by objective demonstrations of likely net benefit to the person being treated. This medical activism is being challenged principally because of its costliness but also on theoretical grounds because the powerful chemical, physical, and surgical interventions now available can do good, but they can also do harm (*10*). Consequently, two components of the U.S. Department of Health and Human Services, the

Health Care Financing Administration (HCFA) and the Agency for Health Care Policy and Research, have embarked on a systematic program of evaluation of medical practices (*11–13*) to provide a more solid empirical basis for patient care and health care management.

The analysis of variations of patterns of practice and of patterns of outcomes is an integral component of this program and serves as a problem-finding tool. The variations investigated are those visible over time and among geographic areas.

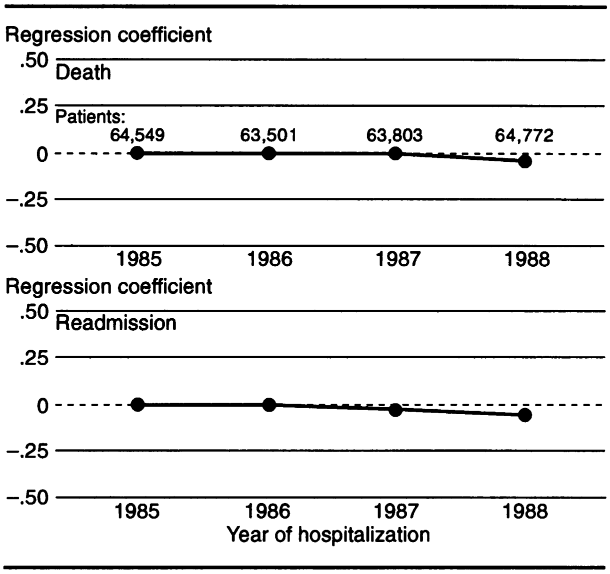
Materials and Methods

The data sets used in this study have been described elsewhere (*12*). Briefly, they consist of the Medicare eligibility or HISKEW file, and claims files, most particularly the MEDPAR or inpatient file that contains the consolidated bills submitted by hospitals for admissions of Medicare beneficiaries. The data contained in the eligibility file include descriptions of the sociodemographic characteristics of the patients, ZIP code of residence, dates of entitlement to coverage, and dates of birth and death. The hospitalization file contains, most importantly, an identification of the hospital and its location, patient demographic data, dates of admission and discharge, the principal diagnosis (the cause of the admission ascertained at discharge) and up to four additional diagnoses (ICD-9-CM codes) representing comorbidities present at the time of admission or complications that occurred during the stay, and up to three ICD-9-CM procedure codes. Only admissions to acute care hospitals were considered in our analyses.

Population-based admission and mortality rates were analyzed at the level of the ZIP code of residence of the Medicare enrollee. All admissions to acute care hospitals and all deaths in fiscal year 1989 were evaluated. The sociodemographic factors taken into account in the multiple regressions using the Poisson Model (*14* and appendix on page 10) included, for each ZIP code area, the average age, the proportion of persons identified as black, the proportion male, as well as the variances and covariances of these proportions. In addition, a measure of the socioeconomic status of the ZIP code area, the Claritas ZQ™ score, was evaluated and found to be a statistically significant contributor to the probability of admission. The average duration of enrollment in health maintenance organizations was also investigated, but it did not have a statistically significant effect at the level of the ZIP code area.

For the evaluation of trends in post-admission mortality rates in the fiscal years 1985 to 1988 (1989 in the case of admissions for acute myocardial infarction), a separate cohort was constructed for each year in the following manner. The hospitalizations of an essentially random 5-percent sample of patients discharged in a given year were placed in disease categories on the basis of the principal diagnosis. If a patient had more than one admission within a category, one was picked at random to serve as the representative admission. Each patient was followed from the day of admission until the end of the next fiscal year to ascertain the date of death, if death had occurred. Thus, followup was for a minimum of 12 months and a maximum of 24, and the patient was withdrawn from the study alive if death had not occurred by the end of the calendar year following the year of admission.

Figure 1. Trends in risk–adjusted mortality and readmission rates following hospitalization for any reason (all admissions), using Cox Proportional Hazards Model



In addition, the hospitalization history of each patient in the 6 months prior to the index admission was characterized. In the analysis of time to readmission, followup began on the day after discharge and continued similarly to the end of the next fiscal year. A person who was not readmitted in this period was withdrawn at its end. A person who died in this interval was withdrawn on the date of death.

Time trends in post-hospitalization mortality and readmission rates were analyzed by means of the Cox Proportional Hazards Model (*15* and appendix). Application of the model produces a convenient quantity, a regression coefficient, that summarizes the effect over the duration of the followup of having been hospitalized in a given subsequent year in comparison to 1985.

Figure 2. Trends in risk–adjusted mortality rates following hospitalization for selected medical conditions, using the Cox Proportional Hazards Model

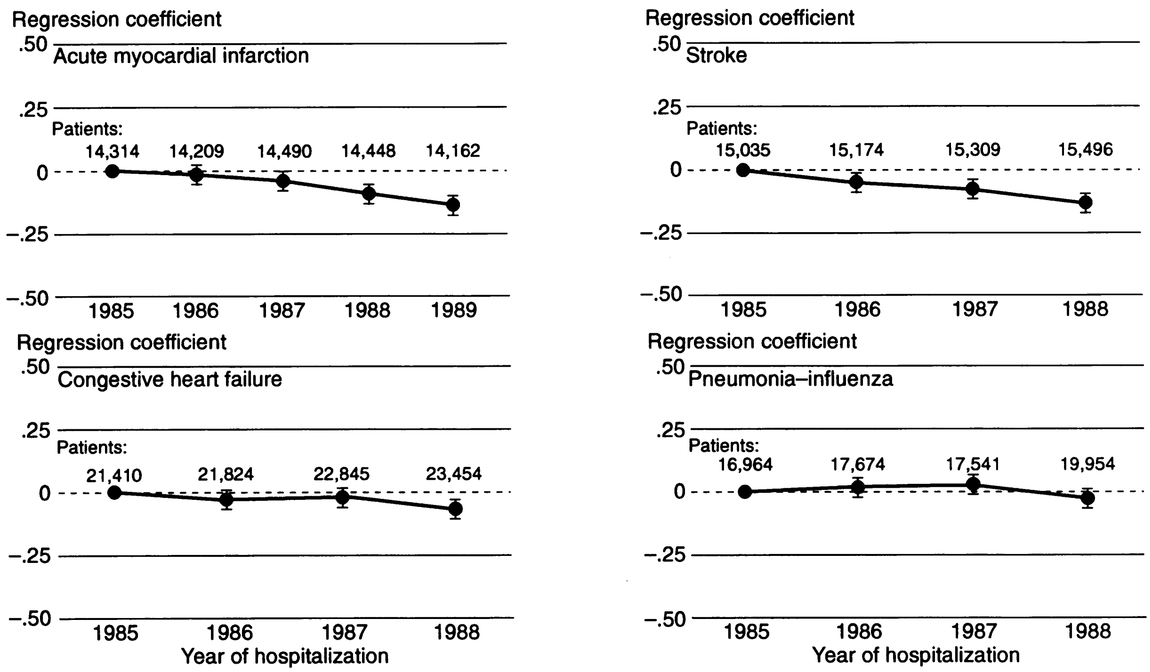
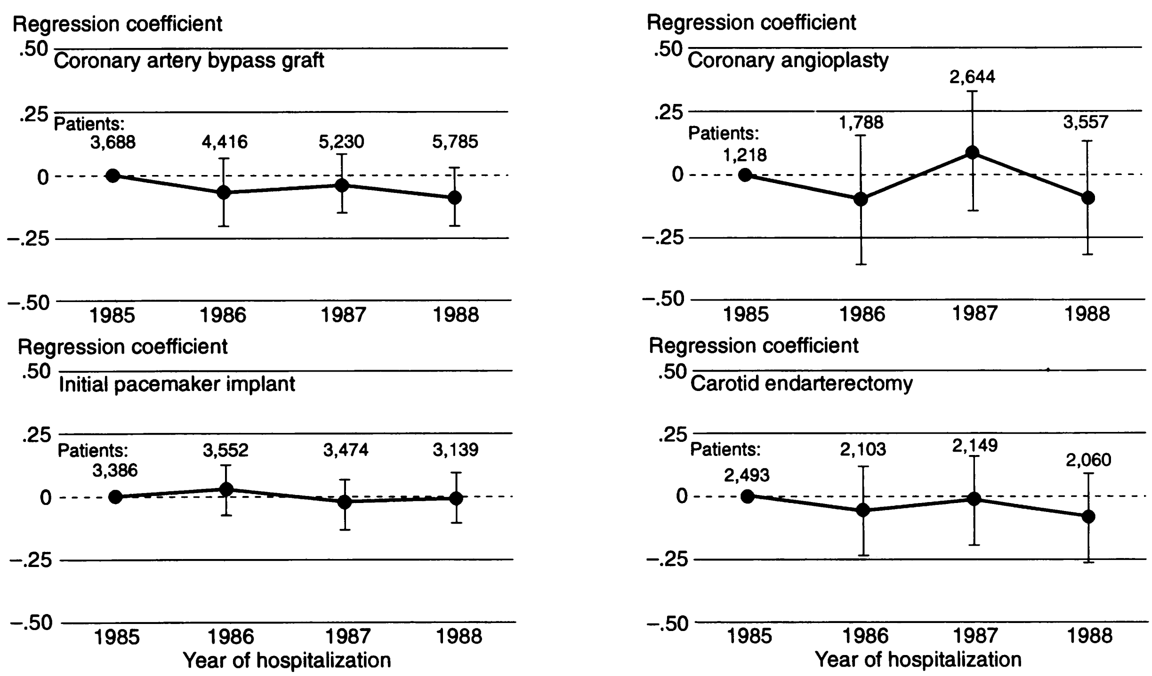


Figure 3. Trends in risk–adjusted mortality rates following hospitalization for selected procedures, using the Cox Proportional Hazards Model



The patient characteristics represented by covariates in this model were age, race, sex, and selected comorbidities. The comorbidities consisted of additional diagnoses beyond the principal that were reported on the claims and were chosen because they are chronic diseases likely to be present at the time of admission, rather than complications that developed during the course of the hospitalization. In addition, whether the current admission was a transfer from another acute care hospital, the number of admissions within the 6 months prior to the current admission, stratified into six groups by risk of death, and the time since the most recent prior admission in each risk group also made up the characteristics profile. For the cohort containing patients not segregated according to reason for admission, 17 additional covariates that classify the patient into risk groups defined by the principal diagnosis were assigned. The cases in this cohort constituted a random 1-percent sample of the persons hospitalized in each year.

The methodology for the analysis of geographic variations in post-admission mortality rates is described in detail elsewhere (*14*). The hospitalizations analyzed were those of persons discharged in fiscal 1989. The analytic approach was similar to that used for the analysis of time trends, but it differs in some details. All hospitalized persons were included, but only one admission per person was selected at random for those with more than one hospitalization in the fiscal year. The selected admission was then assigned to the various condition or procedure categories.

Also, a slightly different set of risk factors was used. Most importantly, the influence of each specific diagnosis (ICD-9-CM code) on the probability of death was estimated and was used as an adjustment in estimating the influence of the hospital or other locale (State, Metropolitan Statistical Area [MSA]) on the probability of the patient's death. Variables characterizing the time since the most recent prior admission were not included in the models. Followup was for a fixed 182 days for all cases. The observed probabilities of death were computed by means of the life-table method (PROC LIFETEST of SAS [*A*]). The time-to-failure model was the Bailey-Makeham (*14* and appendix). The model produces estimates of the influence of individual risk factors on the excess initial hazard of dying that immediately follows the admission, on the persistence of the excess initial risk, and on the long-term risk.

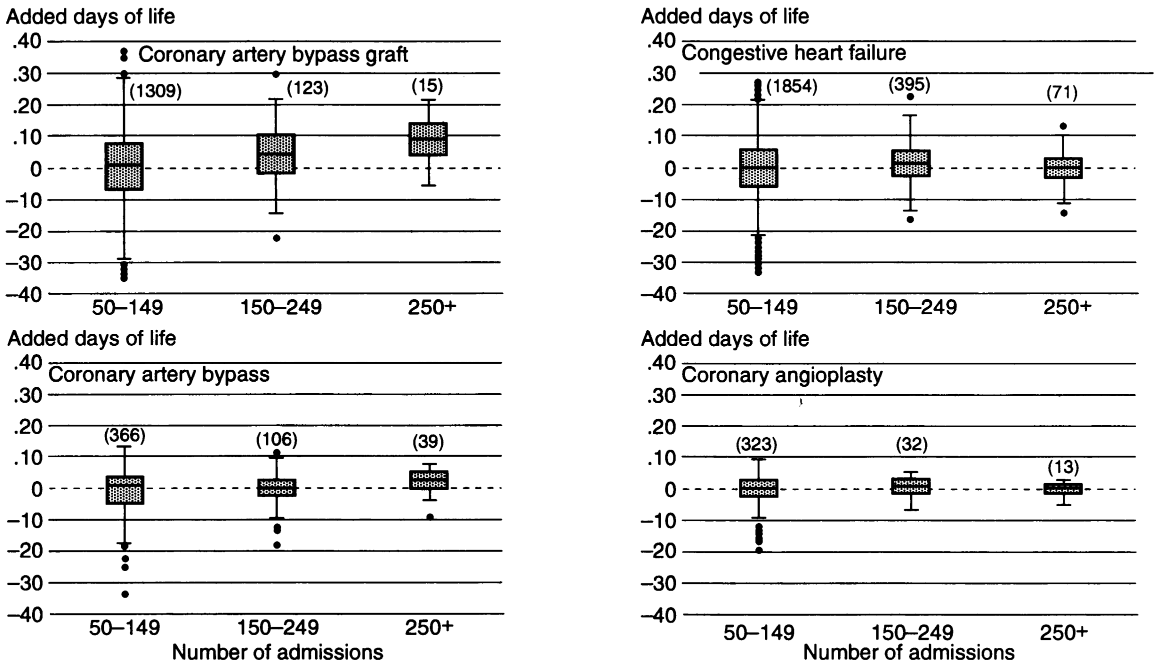
Estimates of the magnitudes of reasonably achievable reductions in admission and in mortality rates for those dying 180 days of admission were obtained in the following way. Case-weighted average rates of admissions or deaths associated with the hospitals or MSAs were first calculated. Then, estimates of these quantities in the lowest quartile were obtained and were applied to the upper three quartiles of the

Table 1. Admission and mortality rates for members of the Medicare population, 1989

		Mortality		*Admissions*	
State	*Enrollees*	*Observed (percent)*	*Relative[1] rate*	*Observed (percent)*	*Relative[1] rate*
United States	33,937,624	4.64	1.00	28.30	1.00
Alaska	22,781	4.01	1.03	23.66	.98
Alabama	552,212	4.75	.98	34.02	1.08
Arkansas	499,764	4.10	.96	25.87	.97
Arizona	405,118	4.74	1.00	33.31	1.07
California	3,389,081	4.52	1.00	25.30	.96
Colorado	332,214	4.29	.94	23.82	.88
Connecticut	357,913	4.51	.99	23.91	.92
District of Columbia	84,306	5.14	.95	28.98	1.09
Delaware	90,297	4.81	1.07	29.09	1.05
Florida	2,272,416	4.25	.96	24.81	.91
Georgia	753,466	4.74	1.00	34.03	1.10
Hawaii	128,225	3.47	.86	18.98	.80
Iowa	472,782	4.72	.97	27.44	.97
Idaho	134,869	4.32	.98	24.21	.89
Illinois	1,564,972	4.80	1.02	28.03	1.02
Indiana	788,178	4.79	1.03	30.65	1.07
Kansas	375,123	4.69	.97	31.04	1.09
Kentucky	554,411	4.86	1.03	34.14	1.10
Louisiana	551,853	4.82	.95	35.40	1.14
Massachusetts	905,413	4.68	.97	27.69	.99
Maryland	506,058	4.78	1.04	31.70	1.14
Maine	189,416	4.70	1.00	27.36	.94
Michigan	1,273,145	4.64	1.02	27.94	1.01
Minnesota	608,117	4.57	.95	24.18	.88
Missouri	811,534	4.85	1.00	28.50	.97
Mississippi	381,353	4.86	.96	40.20	1.23
Montana	120,561	4.30	.94	32.08	1.16
North Carolina	911,974	4.65	1.01	27.45	.88
North Dakota	101,839	4.45	.93	32.11	1.15
Nebraska	247,170	4.82	.99	26.27	.93
New Hampshire	140,945	4.70	1.02	24.90	.90
New Jersey	1,004,969	4.77	1.05	29.37	1.08
New Mexico	181,243	4.08	.91	25.94	.92
Nevada	135,655	4.19	1.06	24.56	.95
New York	2,633,484	4.76	1.00	25.70	.92
Ohio	1,524,779	4.75	1.03	29.23	1.03
Oklahoma	459,695	4.85	1.02	32.75	1.12
Oregon	432,901	4.53	1.00	22.83	.85
Pennsylvania	2,032,102	4.80	1.05	30.06	1.05
Rhode Island	165,040	4.63	.99	26.81	.94
South Carolina	452,836	4.61	1.01	27.97	.90
South Dakota	114,224	4.64	.95	30.43	1.04
Tennessee	713,817	4.80	1.02	34.29	1.11
Texas	1,872,731	4.68	1.00	29.20	1.00
Utah	160,454	4.12	.94	22.26	.84
Virginia	740,870	4.62	1.00	31.13	1.04
Vermont	77,353	4.83	1.02	24.51	.86
Washington	628,765	4.38	.97	23.43	.88
Wisconsin	706,264	4.62	.98	28.47	1.04
West Virginia	320,409	4.79	1.02	33.56	1.09
Wyoming	52,529	4.52	1.01	32.25	1.22

[1]Relative admission and mortality rates are ratios of the observed to the predicted rates.

Figure 4. Variations among hospitals in risk-adjusted mortality rates , using the Bailey-Makeham Model



hospitals or MSAs. Variations in case-mix or severity of illness were corrected for by multiplying the revised numbers of admissions or deaths in the upper three quartiles by the ratio of the predicted admission rate or probability of death in each hospital or MSA to the national average predicted admission rate or probability of death. The projected reduction is the ratio of the difference between the actual and the projected numbers of admissions to the actual number of admissions or deaths. Only hospitals or MSAs with 50 or more cases were used in the computation of the reductions in mortality rates.

Results

Figure 1 illustrates the analysis of variations over time (1985–88) in the probability of death after hospitalization and the probability of readmission after discharge alive. On the vertical axis are the coefficients for the covariates identifying the year of the hospitalization obtained in a regression employing the Cox Proportional Hazards Model and containing an array of patient risk factors derived from the MEDPAR claims file.

The regression coefficients in this figure are approximately equal to the proportionate change in the hazard of dying following hospitalization or of readmission in 1986, 1987, or 1988 compared with 1985. They represent an effect averaged over the duration of followup of 1–2 years. The numbers running horizontally across the upper panel are the number of patients drawn from each annual cohort (a 1-percent random sample of persons hospitalized). Their vertical position reflects the trend over time in the numbers of persons hospitalized. The error bars represent ± 2 standard errors of the coefficient estimate. Therefore, the declines in both the probabilities of readmission and of death do attain a high level of statistical significance ($P < 0.05$). Analysis by means of the Bailey-Makeham Model indicates that the lower risk of death in 1988 compared with 1985 reported by the Cox Model regression results predominantly from a reduction in the long-term risk (data not shown).

Figure 2 presents for 5-percent annual samples a more detailed assessment of the trends over time in the numbers of persons hospitalized for selected medical conditions or procedures and in the probability of their death after hospitalization. Rates of admission for acute myocardial infarction (AMI) or stroke appeared to remain rather stable over the years 1985–88, but the risk-adjusted mortality rates declined perceptibly and statistically significantly ($P < 0.05$). In the case of congestive heart failure, admissions appeared to be rising slightly and deaths were stable. In the admissions for pneumonia or influenza, the admission rate was more variable and the mortality rates revealed no clear trend.

Figure 3 presents analogous data for persons admitted for selected surgical interventions. The increases between 1985 and 1988 in admissions for coronary artery bypass (up 57 percent) and for coronary balloon angioplasty (up 192 percent) are striking, particularly in the face of essentially stable risk-adjusted hazards of dying.

Variations among the States in population-based admission and mortality rates are summarized in tables 1, 2, and 3. Table 1 presents the observed and the ratios of the observed to predicted (relative) admission frequencies irrespective of cause and the observed and the ratio of the observed to the predicted overall frequency of death for Medicare enrollees in fiscal 1989. The predictions are based on assessments of the influence of demographic and socioeconomic factors in national analyses and their application to the particular demographic and socioeconomic population mix in each State. The experience of Hawaii is quite distinctive, with strikingly low observed-to-predicted admission (0.80) and mortality (0.86) ratios. Rather greater variability is observed in the frequencies of admission than of mortality.

Table 2 presents data on variations in frequencies of admissions for the four selected medical conditions shown in fig. 2, and table 3 shows data for the selected procedures shown in fig. 3. Table 4 presents a summary assessment of the magnitude of the variation among MSAs in admissions irrespective of cause and for two cardiac conditions (acute myocardial infarction and congestive heart failure) and two cardiac procedures (coronary angioplasty and bypass). The variations are quite substantial, with the frequency of admission for any reason, adjusted for demographic and socioeconomic factors, in the uppermost quartile of the MSAs about one-third higher than in the lowest quartile. The disparity in admissions for coronary angioplasty is a factor of nearly 3.

The consequences of the variations in admissions and in post-admission (not population-based) mortality rates are summarized in table 5. It illustrates the reductions in risk-adjusted admission and mortality rates that might be achieved if the rates in the three upper quartiles of MSAs and, in the case of mortality, also of hospitals were those experienced on average in the lowest quartile.

Figure 4 illustrates the analysis of variations among hospitals in risk-adjusted mortality following admission, as represented by the summary statistic "Added Days of Life," for two medical conditions and two surgical procedures. This statistic, computed from results obtained with the Bailey-Makeham Model, is the cumulative difference per patient between the observed probability of death over the period of observation (180 days) and that predicted from the national experience for the patients of the hospital. Positive quantities indicate that the observed probability of death is lower than that predicted.

'Decisions to treat, therefore, currently tend to be driven by individually held expectations of a potentially favorable biologic effect rather than by objective demonstrations of likely net benefit to the person being treated.'

The hospitals in figure 4 were grouped into the three categories indicated according to the number of admissions for the specified condition or procedure. The points made by figure 4 are (*a*) the risk-adjusted mortality rates of hospitals vary widely, (*b*) the variation tends to be smaller for the larger hospitals, and (*c*) there is an inconsistent relationship between the volume of cases and risk-adjusted mortality. For example, in the cases of acute myocardial infarction and coronary artery bypass, patients of the hospitals with a larger number of admissions for that condition or procedure enjoy a larger number of added days of life than those of hospitals with fewer admissions. Indeed, the trends are statistically significant. Such a positive trend is not observed in the case of congestive heart failure or of coronary angioplasty. Whatever the trends in risk-adjusted mortality, however, they are very small compared with the variations among hospitals.

Discussion

The examples presented in this paper illustrate results obtained in a program of epidemiologic oversight carried out by HCFA in the discharge of its responsibility for assessing the quality of the medical care provided to Medicare beneficiaries (*12*). Two components of that program are monitoring of time trends and of geographic variations in rates of mortality, of morbidity, of disability (assessed through a surrogate measure), and of health care expenditures.

The data given in figures 1–3 indicate that trends in risk-adjusted mortality and readmission rates are clearly discernible. The meaning of such trends hinges on the adequacy and correctness of the risk-

Table 2. Admission rates for members of the Medicare population by selected medical conditions, 1989

State	Enrollees	Acute myocardial infarction Observed (percent)	Acute myocardial infarction Relative rate	Congestive heart failure Observed (percent)	Congestive heart failure Relative rate	Pneumonia, influenza Observed (percent)	Pneumonia, influenza Relative rate	Stroke Observed (percent)	Stroke Relative rate
United States	33,937,624	.89	1.00	1.70	1.00	1.23	1.00	.94	1.00
Alaska	22,781	.53	.81	1.18	.95	1.10	1.13	.69	.98
Alabama	552,212	.93	1.00	1.86	1.00	1.49	1.07	1.14	1.07
Arkansas	499,764	.81	.91	1.27	.82	1.08	.94	.76	.88
Arizona	405,118	.98	.99	1.91	1.00	1.62	1.08	1.08	1.04
California	3,389,081	.74	.90	1.53	.98	1.19	1.10	.90	1.03
Colorado	332,214	.74	.85	1.05	.66	1.10	.90	.66	.77
Connecticut	357,913	.84	1.04	1.52	.99	1.01	.96	.79	.91
District of Columbia	84,306	.55	.95	2.22	1.16	.91	1.15	1.01	.93
Delaware	90,297	1.10	1.31	1.86	1.10	.99	.91	.91	.97
Florida	2,272,416	.84	.93	1.41	.86	.79	.69	.84	.91
Georgia	753,466	.90	1.01	1.76	.89	1.53	1.20	1.08	1.02
Hawaii	128,225	.58	.74	1.01	.79	.73	.74	.84	1.08
Iowa	472,782	.96	1.04	1.40	.83	1.30	.98	.92	1.00
Idaho	134,869	.88	.95	1.07	.68	1.17	.94	.74	.84
Illinois	1,564,972	.83	.99	1.81	1.08	1.19	1.05	.94	1.01
Indiana	788,178	.92	1.00	1.95	1.12	1.47	1.15	1.01	1.08
Kansas	375,123	.94	1.03	1.74	1.02	1.67	1.25	.94	1.01
Kentucky	554,411	.99	.99	1.94	1.05	1.93	1.24	1.07	1.08
Louisiana	551,853	.92	1.07	2.37	1.22	1.58	1.12	1.11	1.05
Massachusetts	905,413	1.02	1.15	1.81	1.08	1.23	.99	.85	.93
Maryland	506,058	1.03	1.24	1.96	1.15	1.15	1.07	1.02	1.07
Maine	189,416	1.11	1.13	1.63	.94	1.17	.81	.78	.83
Michigan	1,273,145	.92	1.09	1.81	1.09	1.09	.94	.97	1.06
Minnesota	608,117	.85	.93	1.31	.80	1.18	.89	.77	.86
Missouri	811,534	.91	.96	1.73	.96	1.49	1.09	1.02	1.05
Mississippi	381,353	.97	1.08	2.30	1.12	1.82	1.23	1.26	1.13
Montana	120,561	.86	.93	1.45	.92	1.39	1.04	.89	1.01
North Carolina	911,974	1.07	1.16	1.63	.83	1.17	.92	1.12	1.06
North Dakota	101,839	1.02	1.00	1.90	1.16	1.64	1.16	.97	1.10
Nebraska	247,170	.91	.97	1.41	.83	1.44	1.06	.86	.92
New Hampshire	140,945	1.07	1.17	1.54	.95	1.03	.84	.81	.91
New Jersey	1,004,969	.86	1.03	1.98	1.21	.98	.92	.98	1.06
New Mexico	181,243	.77	.81	1.15	.71	1.58	1.15	.68	.77
Nevada	135,655	.79	.91	1.32	.89	.98	.97	.74	.89
New York	2,633,484	.85	.98	1.68	.98	1.06	.89	.85	.90
Ohio	1,524,779	.87	.99	1.95	1.14	1.29	1.06	.92	.99
Oklahoma	459,695	1.05	1.09	1.72	.97	1.63	1.17	1.10	1.14
Oregon	432,901	.78	.87	1.23	.79	.97	.80	.82	1.00
Pennsylvania	2,032,102	.98	1.06	1.99	1.14	1.12	.93	1.01	1.06
Rhode Island	165,040	1.18	1.26	1.89	1.09	.96	.75	.88	.94
South Carolina	452,836	.91	1.02	1.74	.89	1.08	.88	1.06	.99
South Dakota	114,224	1.24	1.23	1.57	.91	1.89	1.23	.89	.98
Tennessee	713,817	.98	1.02	2.01	1.06	1.67	1.20	1.14	1.12
Texas	1,872,731	.85	.92	1.72	.97	1.36	1.07	.96	.98
Utah	160,454	.78	.88	.94	.60	.98	.82	.68	.79
Virginia	740,870	.93	1.03	1.76	.96	1.26	1.03	1.04	1.02
Vermont	77,353	.87	.94	1.48	.88	1.26	.93	.73	.81
Washington	628,765	.82	.95	1.22	.79	.97	.83	.82	.96
Wisconsin	706,264	.92	1.03	1.67	1.02	1.10	.87	.96	1.07
West Virginia	320,409	1.15	1.14	2.05	1.12	1.61	1.05	1.02	1.05
Wyoming	52,529	.98	1.13	1.43	.95	1.63	1.39	.70	.86

adjustment procedure which, in the examples shown, is contingent on the accuracy of the data submitted on the claims by the hospitals to the HCFA. The limitations in the accuracy of the ICD-9-CM codes have been repeatedly investigated (*16,17*) and are clearly a cause for concern. Thus, the trends might be the result of changes in coding practices such as an increasingly extensive reporting of comorbidities. The internal evidence of the data, that is, the presence of a clear downward trend in mortality following admission for acute myocardial infarction and stroke and its absence in the case of admissions for congestive heart failure or for the pulmonary diagnoses, suggests, however, that is not the likely

Table 3. Percentage admission rates for members of the Medicare population by selected surgical procedures, 1989

State	Enrollees	Coronary angioplasty		Coronary artery bypass surgery		Initial pacemaker implant		Carotid endarterectomy	
		Observed	Relative	Observed	Relative	Observed	Relative	Observed	Relative
United States	33,937,624	.24	1.00	.35	1.00	.19	1.00	.12	1.00
Alaska	22,781	.13	.57	.33	.90	.08	.58	.09	.59
Alabama	552,212	.31	1.38	.44	1.37	.20	1.07	.16	1.30
Arkansas	499,764	.28	1.02	.33	.84	.17	.88	.10	.71
Arizona	405,118	.26	1.14	.52	1.44	.22	1.14	.17	1.31
California	3,389,081	.30	1.16	.35	.95	.20	1.02	.13	1.01
Colorado	332,214	.20	.85	.25	.71	.13	.69	.07	.56
Connecticut	357,913	.20	.86	.35	1.00	.19	.99	.07	.64
District of Columbia	84,306	.14	1.04	.18	.93	.15	.78	.04	.91
Delaware	90,297	.25	.97	.36	1.01	.16	.84	.16	1.30
Florida	2,272,416	.27	.98	.39	.99	.25	1.24	.14	.98
Georgia	753,466	.26	1.23	.37	1.16	.22	1.19	.14	1.31
Hawaii	128,225	.19	.67	.29	.68	.13	.64	.08	.54
Iowa	472,782	.29	1.30	.35	1.04	.15	.77	.09	.81
Idaho	134,869	.24	.85	.42	1.02	.17	.85	.12	.89
Illinois	1,564,972	.23	.95	.34	.99	.18	.93	.10	.86
Indiana	788,178	.33	1.40	.40	1.17	.20	1.05	.13	1.09
Kansas	375,123	.38	1.75	.37	1.02	.21	1.06	.19	1.66
Kentucky	554,411	.18	.79	.39	1.16	.21	1.15	.15	1.21
Louisiana	551,853	.32	1.40	.37	1.16	.24	1.31	.22	1.00
Massachusetts	905,413	.14	.64	.28	.82	.18	.92	.07	.63
Maryland	506,058	.24	1.00	.36	1.06	.15	.79	.13	1.13
Maine	189,416	.18	.81	.30	.86	.16	.81	.08	.66
Michigan	1,273,145	.25	1.04	.40	1.17	.19	.99	.18	1.45
Minnesota	608,117	.21	.90	.30	.84	.14	.70	.06	.48
Missouri	811,534	.35	1.53	.41	1.19	.21	1.06	.13	1.03
Mississippi	381,353	.14	.70	.30	.98	.24	1.32	.16	1.41
Montana	120,561	.45	1.79	.42	1.09	.16	.86	.11	.90
North Carolina	911,974	.23	1.03	.35	1.04	.19	.96	.11	.89
North Dakota	101,839	.22	.96	.42	1.14	.18	.86	.09	.76
Nebraska	247,170	.20	.92	.29	.82	.19	.93	.13	1.14
New Hampshire	140,945	.19	.82	.35	.96	.14	.75	.06	.42
New Jersey	1,004,969	.19	.79	.35	1.00	.22	1.13	.10	.90
New Mexico	181,243	.24	.93	.26	.70	.11	.60	.09	.68
Nevada	135,655	.33	1.05	.42	1.00	.16	.88	.15	.88
New York	2,633,484	.11	.47	.26	.77	.17	.89	.06	.56
Ohio	1,524,779	.23	.97	.36	1.05	.20	1.05	.16	1.28
Oklahoma	459,695	.28	1.24	.36	1.02	.22	1.11	.11	.95
Oregon	432,901	.22	.83	.34	.89	.12	.64	.18	1.31
Pennsylvania	2,032,102	.21	.88	.37	1.02	.20	1.01	.11	.80
Rhode Island	165,040	.13	.56	.24	.70	.16	.80	.08	.64
South Carolina	452,836	.24	1.08	.34	1.04	.21	1.15	.09	.80
South Dakota	114,224	.29	1.28	.41	1.07	.19	.88	.10	.92
Tennessee	713,817	.19	.86	.40	1.17	.18	.97	.13	1.02
Texas	1,872,731	.27	1.07	.36	.98	.21	1.06	.15	1.13
Utah	160,454	.26	1.00	.39	1.03	.20	.98	.08	.59
Virginia	740,870	.21	.92	.34	.99	.17	.86	.11	1.00
Vermont	77,353	.18	.80	.30	.81	.18	.93	.05	.47
Washington	628,765	.25	.92	.40	1.07	.13	.67	.18	1.32
Wisconsin	706,264	.30	1.32	.38	1.10	.17	.93	.11	1.03
West Virginia	320,409	.24	1.02	.36	1.02	.19	1.02	.15	1.14
Wyoming	52,529	.28	1.10	.40	1.05	.15	.80	.14	1.27

explanation. Except in isolated instances of changes in coding rules, there is no clear indication that the observed trends could be due to systematic changes in reporting practices.

Variations among geographic areas and hospitals in reporting practices may also confound analyses of variations in admission and mortality rates. Of particular concern are the Medicare beneficiaries enrolled in health maintenance organizations (HMOs), because they tend to be concentrated in certain States and MSAs. Although Medicare regulations require notification of hospitalizations through submission of "no pay" bills, the degree of conformance with this requirement is unclear. We, therefore, evaluated the

Technical Appendix

The Cox Proportional Hazards Model: In this model, the hazard of dying, the probability per unit time of a person alive at the beginning of an interval of dying in that interval, is

$$h(t) = h_o(t) \exp(\Sigma_i \beta_i z_i) \quad (1)$$

where $h_o(t)$ is the underlying nonparametrized hazard pattern characteristic of the condition, z_i is the covariate representing a particular patient characteristic that affects the probability of death, β_i is the coefficient estimated in the analysis that reports the magnitude and direction of the effect of z_i on the hazard of death, and t is the time since admission (*15*).

To estimate the trend over time in the hazard of death following hospitalization, the annual cohorts were assigned a dichotomous (0,1) covariate to indicate their year of discharge and then assembled into a conglomerate data set. The reference year was 1985 and the trend in risk-adjusted mortality was obtained by estimating the coefficients of the indicator covariates that labeled each of the years 1986–88.

The proportionality assumption in the Cox model requires that the coefficients be independent of time. It is known, however, that both patient characteristics and medical interventions may differentially affect early or late risks of death.

The Bailey-Makeham Model: In this model (*14*), the hazard of dying is formulated in terms of a constant long-term component (δ) and of a component seen early (α, the excess initial risk) which decays exponentially with a rate constant, τ, the reciprocal of the persistence:

$$h(t) = \alpha \exp(-\tau t) + \delta \quad (2)$$

where $\alpha = \exp(\alpha_o + \Sigma_i \alpha_i z_i)$, $\tau = \exp(\tau_o + \Sigma_i \tau_i z_i)$, $\delta = \exp(\delta_o + \Sigma_i \delta_i z_i)$.

Unlike the Cox, the Bailey-Makeham is fully parametric and not subject to the restriction that the hazards be proportional.

The Poisson Model: This model was used for the computation of the number of events, y (admissions or deaths), in a population, N, in a ZIP code area (*14*). The probability of the number of events, $p(y)$, is given by:

$$p(y) = \exp(-\phi) \quad \text{if } y = 0 \quad (3a)$$

$$\text{and } p(y) = R \frac{\mu^y \exp(-\mu)}{y!} \quad \text{if } y > 0, \quad (3b)$$

where $R = (1 - \exp(-\phi)) \div (1 - \exp(-\mu))$ and $\mu = N \exp(\mu_o + \Sigma_i \mu_i z_i)$, $\phi = N \exp(\phi_o + \phi_1 \Sigma_i \mu_i z_i)$.

Table 4. Variations in admission rates among Metropolitan Statistical Areas in fiscal year 1989

Condition	*Ratio of relative admission rates*
All	1.33
Acute myocardial infarction	1.50
Congestive heart failure	1.58
Coronary artery bypass	1.68
Coronary angioplasty	2.70

NOTE: These quantities are ratios of population-weighted observed to predicted admission ratios in the top quartile to the bottom quartile of the Metropolitan Statistical Areas.

extent to which the proportion of time the residents of a ZIP code area were enrolled in HMOs affected admission rates in those areas. HMO membership was expressed in this way because the enrollment of Medicare beneficiaries in HMOs is unstable. Those who were enrolled at any time in a year were members of an HMO, on average, for only half a year. No statistically significant effect of this measure of HMO enrollment was found ($P > 0.05$). It was, therefore, not retained in the models.

The decline over the years in the risk-adjusted hazards of death or readmission is quite encouraging, particularly because concerns had been raised that the introduction of the Prospective Payment System might have resulted in a reduction in the quality of care, with patients being discharged "quicker and sicker." A previous examination of this matter through a comparison of mortality rates of patients hospitalized in 1982 and 1986 also failed to detect a deterioration (*18*).

The declines in mortality rates in the case of hospitalizations for acute myocardial infarction and stroke suggest that the major technologic advances of recent years in the management of cardiovascular disease, such as pharmacological and surgical revascularization, may be bearing fruit. On the other hand, the continuing rapid increase in the frequency of coronary artery bypass surgery and balloon angioplasty in the Medicare population in the absence

Notes for Figures

Figure 1. A 1-percent sample was used, whose size in each year is shown above the data points on the graph. The trend in admission rates is indicated by the vertical positioning of these numbers. The reference year was 1985. In that year, the probability of death within 30 days of admission was 8.7 percent and 16.8 percent within 180 days. The probabilities of readmission were 9.1 percent and 26.3 percent at 30 and 180 days following discharge alive. The error bars represent ± 2 standard errors of the estimate of the regression coefficient (the risk-adjusted hazard) for the indicated year. The model used is specified by equation (1). Additional variables in the multiple regression include age, race, sex, reasons for admission (grouped principal diagnoses), selected comorbidities, type and source of admission, and numbers of admissions and time since most recent prior admission within the 6 months preceding the current admission, stratified by risk. (The lines connecting the data points were added as an aid to the eye. The horizontal line at 0 for the regression coefficient is for reference only.)

Figure 2. The model was as in figure 1, except that the cases analyzed constituted 5-percent samples, the additional variables in the model did not include groups of reasons for admission, and the 30- and 180-day mortality rates in 1985 were for AMI—24.8 percent and 33.8 percent; Stroke— 20.5 percent and 34.5 percent; Pneumonia-influenza—15.3 percent and 28.4 percent; Congestive heart failure—14.9 percent and 32.1 percent.

Figure 3. The model was as in figure 2. The 30- and 180-day mortality rates in 1985 were for coronary artery bypass grafts—6.6 percent and 10.7 percent; Coronary angioplasty—3.5 percent and 5.7 percent; Initial pacemaker implant—3.3 percent and 12.5 percent; Carotid endarterectomy—2.1 percent and 5.7 percent.

Figure 4. The model was the Bailey-Makeham. "Added Days of Life" is the cumulative difference per person between the observed probability of death and that predicted from the national experience for the patients admitted to the hospital for the specified condition over the period of observation of 180 days. The upper and lower boundaries of the boxes identify the values of this parameter between which fall the middle 50 percent of the hospitals in the relevant size category, the interquartile (25–75 percent) range. The line through the middle of the box is the median value. The horizontal bars linked to the box by the vertical lines ("whiskers") mark off distances equal to 1.5 the interquartile range above and below the box. The symbols outside the "whiskers" represent hospitals whose "added days of life" parameter fall beyond those extremes. The quantities in parentheses are the numbers of hospitals within each size category.

of evidence of increasing effectiveness (the risk-adjusted hazards of death are unchanged) is problematic. Of course, it may be that an increasing proficiency in the performance of these procedures is masked by an increasingly adverse case-mix, the severity of illness of the patients not having been adequately captured by the claims data.

An approach to the characterization of changes over time in the manner in which patients with particular conditions are managed and of the effects of interventions on outcomes was presented elsewhere (*12*). In our analyses, however, variables representing treatments were deliberately excluded from the risk-adjustment models because our objective was to detect, if possible, trends and variations in the effectiveness or quality of care. Inclusion of treatment variables in the models would have masked such trends and variations if they were due to changes over time or to differences among areas in the use of the treatments.

Tables 1–4 and figure 4 describe the very substantial variations in admission and mortality rates at geographic levels as broad as the State and as narrow as the hospital. Since the data currently available suggest that risk-adjustment models based on claims data reveal patterns of variation quite consistent with those based on detailed clinical data (*19*), our results strongly imply that the health care system, at least in the Medicare environment, is not functioning efficiently or, at least, sufficiently effectively. Table 5 indicates the magnitude of the reductions in admission and in post-admission mortality rates that might be achieved.

The essential conclusion to be drawn from our material is that the systematic assessment of variations in patterns of care and patterns of outcomes, as a problem-finding component of a program of epidemiologic oversight, is entirely feasible and is capable of yielding clinically and administratively useful information. Major efforts are currently under way (*13,19,20*) to establish mechanisms to correct the detected problems by providing to patients, clinicians,

Table 5. Consequences of variations in patterns of practice and patterns of outcomes

Condition	Reductions achievable (percent)[1] Admissions	Mortality
Analysis by hospital:		
All	...	17
Analysis by Metropolitan Statistical Area:		
All	20	12
Acute myocardial infarction	21	17
Congestive heart failure	31	7
Coronary artery bypass	21	[2]35
Coronary angioplasty	46	[2]45

[1]Reductions achieved in the number of deaths within 180 days of admission by bringing the rates in the top 3 quartiles down to the average rate in the lowest quartile. Analyses of mortality rates were limited to hospitals and Metropolitan Statatistical Areas with 50 or more cases.

[2]The frequencies of admissions and the mortality rates in these categories are relatively low (see figure 3). In such cases, rather small reductions in the number of deaths may result in relatively large percentage changes. These estimates must, therefore, be viewed with caution.

and managers the data and decision-support tools to enable them to make choices in the selection of health care strategies based more clearly on their probable impacts on the health of patients.

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