Analysis of underwriting factors for AAPCC

by Kenneth G. Manton and Eric Stallard

The adjusted average per capita cost (AAPCC) formula is used to determine payment to health maintenance organizations (HMOs) by Medicare. The four original underwriting factors (i.e., age, sex, institutional status, and welfare status) for the AAPCC were calibrated from the Current Medicare Surveys for 1974-76. Those factors have been updated by various actuarial adjustments.

Revised calculations of the AAPCC underwriting factors are presented using survey data from the 1984 National Long-Term Care Survey and expenditure data from the Medicare Part A and Part B bill files. Also examined is the effect on the underwriting factors of chronic functional disability, defined as having one or more chronic limitations in activities of daily living. Comparison of alternative underwriting factors is conducted by simulating the dollar impact on payment to HMOs for select enrollee populations.

Introduction

Currently, the adjusted average per capita cost (AAPCC) formula is used to pay health maintenance organizations (HMOs) for medical services provided to enrolled Medicare beneficiaries. This formula is based on four underwriting factors (i.e., age, sex, institutional status, and for non-institutionalized persons, their welfare status) that adjust payment for variables thought to affect the costs of providing health services to Medicare beneficiaries (Kunkel and Powell, 1981). In addition, ratio adjustments are used to scale payments from national to local area averages to control for geographic variation in cost.

In this article, we analyze the AAPCC as currently used by calculating revised AAPCC factors from a more recent survey, i.e., the 1984 National Long-Term Care Survey (NLTCS), and comparing the payments implied by the original AAPCC and those suggested by the AAPCC revised with the 1984 NLTCS data. An important feature in this analysis is that we have the individual Medicare Part A and Part B payment records and can compare the actual Medicare expenses in the fee-for-service (FFS) sector with the payment amounts from either AAPCC schedule. Having the individual data permits us to examine additional underwriting factors defined from any variable deemed to be relevant from the NLTCS.

The goals of this analysis are threefold:

- To demonstrate the effect of updating the AAPCC factors (from the period of 1974-76 to 1984) to reflect more current underwriting experience.
- To evaluate the effect of adding an additional underwriting factor based on chronic disability.
- To use simulation methods to demonstrate the dollar impact of these two changes on payments to HMOs and the implications for HMO solvency.

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Reprint requests: Kenneth G. Manton, Ph.D., Duke University, Center for Demographic Studies, 2117 Campus Drive, Box 90408, Durham, North Carolina 27708-0408.

Although our revised and modified estimates of AAPCC factors are not intended to replace the current AAPCC factors, the analyses presented represent the likely effects of such replacement.

Background

Several studies (e.g., Beebe, Lubitz, and Eggers, 1985; Lubitz, Beebe, and Riley, 1985) have assessed the ability of the AAPCC underwriting factors to explain cost variation. The amount of cost variation explained by the original four factors was low—only 0.5 percent of the variance in a recent analysis for cross-sectional data (Ash et al., 1989; Anderson, 1983; Beebe, Lubitz, and Eggers, 1985; Lubitz, Beebe, and Riley, 1985). This means that the formula may not accurately estimate the payments that HMO enrollees would have received had they remained in the FFS sector (Eggers, 1980; Eggers and Prihoda, 1982). As a consequence, there has been a search for additional underwriting factors that could better describe cost variation.

Underwriting factors, in addition to explaining payments and service use, must be based on easily measured factors and not be easily manipulated by providers. Prior service use (e.g., Beebe, Lubitz, and Eggers, 1985), disability status, and, most recently, diagnostic cost groups (DCG) have been considered as additional underwriting factors (Epstein and Cumella, 1988; Ash et al., 1989). None of these have explained more than 10 percent of the variance of Medicare costs. Usually less than 5 percent is explained (Ash et al., 1989). As a result of events unforseeable by either the HMO or the individual, the maximum achievable R^2 has been argued to be less than 100 percent—possibly closer to 15 percent to 20 percent (an estimate of 14.5 percent is provided in Newhouse et al., 1989; also see Welch, 1985).

In addition, searches for improved underwriting factors have been motivated by recent changes in Medicare payment practice and policy. The original AAPCC underwriting factors (Kunkel and Powell, 1981) were used unchanged through 1984. The first modification was in 1985, and since then the factors have been periodically revised. The existing underwriting factors have been updated using current Medicare expenditure data but not new survey data.

These revisions, however, may not fully reflect changes in the relation of the underwriting factors caused by alterations in either provider or consumer behavior as a result of changes such as the introduction of the prospective payment system (PPS) to pay for acute hospitalization. The effects of PPS on all types of Medicare-paid service consumptions, not just hospital admission rates and length of stay (LOS), have been documented to vary both by types of service (e.g., use of home health services; Manton and Liu, 1990) and by sex and age groups (Sager et al., 1989; Manton, Vertrees, and Wrigley, 1990). Thus, without revisions that reflect these changes (e.g., in the mix of individual types of Part A services for different demographic and welfare groups) the AAPCC formula may not accurately estimate the Medicare payments that HMO enrollees would have received had they remained in a changing FFS market.

We present revised estimates of the underwriting factors using the experience of 22,674 Medicare beneficiaries in the 1984 NLTCS and their linked Medicare Part A and Part B bills for April 1, 1984, to March 31, 1985. We used the basic logic of the original AAPCC calculation to determine how new data and additional factors would affect existing ratios.

In Table 1, modified AAPCC factors are presented where a single ratio adjustment (scale factor = 1.03846) was applied to the original AAPCC factors (Kunkel and Powell, 1981) to control for demographic shifts '(primarily aging) and sampling variability in the1984 NLTCS. These AAPCC factors as modified provide estimates of the payment that would be received by an HMO if an NLTCS sample person were enrolled.

For Part A, the basic monthly payment of $104.13 is multiplied by the Part A factor in Table 1 appropriate to the person's sex, age group, institutional status, and if non-institutionalized, his or her welfare status. For example, for an 82-year-old woman living in the community and receiving no welfare payments, Medicare would pay an HMO $118.92 (i.e., $104.13 x 1.142) per month. For Part B, the basic monthly payment of $50.34 is multiplied by the appropriate Part B factor in Table 1.

The basic monthly costs ($104.13 and $50.34) were derived from Medicare records linked to members of the 1984 NLTCS sample for services delivered during the period April 1, 1984, to March 31, 1985. Additional adjustments for geographic variation and for within-county classification of HMO versus non-HMO populations are performed in practice (e.g., see Kunkel and Powell, 1981) but not in this article.

Thus, one set of tables was calculated to reproduce the AAPCC demographic underwriting factors used by the Health Care Financing Administration (HCFA) in 1984 to pay Medicare HMO programs. This represents Schedule 1. The original AAPCC factors were based on 3 years (1974-76) of the Current Medicare Survey (CMS), representing approximately 20,000 beneficiary-years of observation for both the disabled and the

Table 1

Modified AAPCC payment factors for the aged, by institutional status, type of coverage, sex, and age[1]

Type of coverage, sex, and age	Institutional	Non-institutional: Welfare	Non-institutional: Non-welfare
Part A, hospital insurance			
Males:			
65-69 years	2.129	1.402	0.727
70-74 years	2.233	1.610	0.831
75-79 years	2.440	2.025	1.038
80-84 years	2.440	2.388	1.246
85 years or over	2.440	2.700	1.402
Females:			
65-69 years	1.713	0.935	0.623
70-74 years	1.973	1.194	0.727
75-79 years	2.285	1.558	0.935
80-84 years	2.285	1.869	1.142
85 years or over	2.285	2.233	1.298
Part B, supplementary medical insurance			
Males:			
65-69 years	1.788	1.226	0.869
70-74 years	1.942	1.431	1.022
75-79 years	1.942	1.584	1.124
80-84 years	1.942	1.737	1.175
85 years or over	1.942	1.737	1.175
Females:			
65-69 years	1.584	1.124	0.715
70-74 years	1.635	1.175	0.818
75-79 years	1.737	1.277	0.971
80-84 years	1.737	1.277	1.022
85 years or over	1.737	1.277	1.073

[1]Single ratio adjustment (scale factor = 1.03846) was applied. Entries in this table define payment Schedule 1.

NOTE: AAPCC is adjusted average per capita cost.

SOURCE: Adapted from Kunkel and Powell (1981, Table 1).

elderly (65 years of age or over) Medicare beneficiaries (Kunkel and Powell, 1981).

In contrast, the 1984 NLTCS has more than 22,000 person-years of experience during the 1-year followup for elderly (age 65 or over) Medicare beneficiaries. The 1984 NLTCS does not represent disabled beneficiaries under age 65 and, consequently, underwriting factors cannot be determined for that group. However, because of this sample restriction, the NLTCS yields more precise estimates of the demographic factors for the elderly—especially for the oldest-old (85 years or over) population with its high per capita service use.

The 1984 NLTCS is approximately 9 years more current than the CMS and reflects experience after a number of Medicare policy changes (e.g., the introduction of PPS; changes in regulations for Medicare home health services). Not only did regulatory and reimbursement changes occur from 1974 to 1984 but, in response to those changes, there have been significant changes in the consumption behavior of Medicare beneficiaries and the marketing and service provision strategies of health care providers. For example, changes have occurred in the use of acute care hospitals, i.e., lower rates of hospitalization, shorter LOS, and more acute case mix (Eggers, 1987; Liu and

Manton, 1988) and home health services—large increases in consumption, especially among non-married persons, and in the utilization of Medicare skilled nursing facilities (SNFs) (Sager et al., 1989; Vertrees, Manton, and Mitchell, 1989). As a consequence, the 1984 NLTCS may better represent the current patterns of consumption behavior of Medicare beneficiaries within the various underwriting classes as originally defined. Additionally, the NLTCS was replicated in 1989, and a further followup is being planned for 1993, so that data reflecting further consumer and provider adjustments to PPS and other changes will be available to update and to validate the underwriting factors as the Medicare program, and responses by consumers and providers, evolve.

In addition to replicating the four original underwriting factors with the 1984 NLTCS (Schedule 2), we examined disability measures that could serve as additional underwriting factors. Disability has considerable power to predict objective health measures, health service use, health service costs, and mortality (Manton, 1988a, b). For example, disability, as reported in the 1982 and 1984 NLTCS, predicted changes in Part A service use, e.g., a 20-percent longer hospital stay for the chronically disabled in 1984 (Liu and Manton, 1988; Manton and Liu, 1990). In Schedule 3, we examine how a simple dichotomous underwriting factor reflecting the presence or absence of chronic disability affected the prediction of service use. Other more sophisticated, but complicated, measures of chronic disability that have been found to be even more potent predictors of service use and mortality (e.g., Manton, Stallard, and Woodbury, 1991) were not used in the analysis. We are attempting to demonstrate whether disability has a significant effect on expenditures, not to determine the optimal measure of disability to be used as an underwriting factor.

Moreover, clinical, epidemiological, and gerontological research have demonstrated that disability, defined in terms of activities of daily living (ADL) impairments, is reliably and objectively measurable from survey data (e.g., Wiener and Hanley, 1989). The richness of the 1984 NLTCS survey data could be used to examine a number of additional underwriting factors for the AAPCC.

We also restricted our attention to stratification on the simple presence or absence of chronic disability, both because it was featured in recent congressional legislative proposals regarding expansion of Medicare benefits to include long-term care (LTC) services where eligibility was based on functional disability (defined by limitations in ADL requiring personal or mechanical assistance) and because it is being explored in various HCFA demonstration projects on capitated organizations (e.g., social health maintenance organization [S/HMO]; Program for All-inclusive Care for the Elderly [PACE]) as a factor in determining eligibility for LTC services. The possibility of future modifications to Medicare benefits based on disability as an eligibility criterion makes it prudent to examine the effects of disability on the AAPCC formula.

Data

The 1984 NLTCS is the second cycle (first cycle was in 1982; the third cycle was completed in 1990) of a complex multistage sample survey of disabled Medicare beneficiaries (age 65 or over) living in the United States on April 1, 1984. Details of the survey design, instrumentation, and methods for imputing characteristics of non-respondents based on known information about these persons are presented in Manton, Stallard, and Woodbury (1991). There were 22,674 sample persons alive on April 1, 1984, of whom 8,825 were male and 13,849 were female. Sample weights are provided that are proportional to the inverse of the case selection probabilities. These range from 0.63 to 1.38, with mean 1.00, on a relative scale.

Linked to the individual survey record is a complete set of Medicare Part A and Part B bills for the period April 1, 1984, to March 31, 1985. These bills were categorized according to month of service, except for service episodes spanning 2 or more months. In the latter case, the bills were prorated by days of service in each month.

One feature of the NLTCS design that is important for the AAPCC calculations is that it is based on a list sample drawn from HCFA administrative records on all Medicare-eligible persons age 65 or over. This means that for all persons drawn in the 1984 NLTCS sample there is a 100-percent followup of both Medicare Part A and Part B service use and mortality. No cases are lost to followup because persons had to be in the Medicare eligibility files to be initially sampled.

It is estimated by HCFA, Office of the Actuary, that about 97 percent of all persons in the United States age 65 or over are in these files (Ruther and Reilly, 1989). Obviously, the files contain 100 percent of Medicare-eligible persons. The response rates for the detailed surveys are very high—about 95 percent in both years (1984 and 1982), after adjusting for mortality (Macken, 1984; Manton, 1988a, b). Though this response rate is high, it is consistent with the response rates for other health-related surveys (e.g., the National Health Interview Survey) where the field work was conducted (as for all three waves, 1982, 1984, and 1989 NLTCS) by the U.S. Bureau of the Census (1985).

The 95-percent response rate is calculated by defining two types of non-interview (Type A and C). Type A refers to people who leave the sample universe, e.g., those who die before the survey date or who move beyond the geographic area of the primary sampling units (PSUs). A person leaving the sampling universe is not counted as a non-respondent because there is no possibility of an interview. Type C refers to persons who, for example, refused the interview, had residences in the geographic area of the PSU but could not be contacted, required a proxy respondent and none could be found, or persons dying during the survey operation prior to being contacted for an interview. A Type C non-interview is counted as a non-response after adjusting for mortality.

The NLTCS was conducted in two stages, a screening stage and a personal interview stage. The sample was frozen on April 1, 1984, with 22,674 persons. Of these, 21,876 persons were screened for disability (or contacted for address verification if disabled in 1982)—a response rate of 97.5 percent. As there were 376 persons who died during the screening operation, who could not possibly be screened, the adjusted response rate is 98.1 percent.

In 1984, the personal interview stage involved either a detailed community interview or an institutional interview. Of 6,264 persons identified as eligible for a detailed community interview, 5,934 persons (94.7 percent) were interviewed. Of 1,836 persons identified as eligible for the institutional survey, 1,763 persons (96.0 percent) were interviewed. The combined community/institutional response rate was 95.0 percent. There were 154 persons who died during the personal interview stage, who could not possibly be interviewed, yielding an adjusted response rate of 96.9 percent. The adjusted response rate for both stages was 95.1 percent. Thus, the combination of the use of a Medicare list sample and excellent field work by the U.S. Bureau of the Census make the NLTCS an excellent vehicle for examining AAPCC issues.

Despite the high response rate, one cannot assume that remaining non-response bias is insignificant. A study in 1966 found that non-respondents, who were only 5 percent of the sample, used 15 percent of hospital services, as verified by hospital records (National Center for Health Statistics, 1966). Similarly, the 5-percent of the 1984 NLTCS sample who were non-respondents had high levels of Part A and Part B service use and mortality. Because we had a list sample, and Medicare service use and mortality for non-respondents is known, adjustment for this source of non-response bias was possible with the NLTCS design. More difficult were adjustments for death and changes in health and functioning occurring during the interview period—especially for the very frail and the oldest old (age 85 or over) (Manton and Suzman, 1992). We examined several ways of making these adjustments and selected a maximum likelihood-based imputation model (Manton, Stallard, and Woodbury, 1991). Without such adjustments, and the existence of a list sample for their implementation, there is risk of serious bias in calculating cell-specific payment rates.

To replicate the AAPCC factors in Table 1, we deleted five cases who were in the end stage renal disease program. All other persons were classified according to their status on five variables on the first of each month. This yielded 264,591 person-months of observation. There were 7,437 person-months of observation lost as a result of the death of approximately 1,240 persons in the followup period of April 1, 1984, to March 31, 1985. The distribution of these person-months of followup on the five variables is presented in Table 2. The columns headed with "Both" correspond to the AAPCC categories in Table 1. We present additional detail showing a breakout of institutional persons by welfare status and a breakout of non-institutional persons by disability status. Next, we discuss the definition of each factor.

Sex and age

Sex and date of birth were determined from HCFA administrative records. Age was determined from the difference between current date and date of birth and classified into one of five groups at the start of each month.

Table 2

Distribution of person-months of followup by underwriting factors for the revised AAPCC payment factors, by institutional status, sex, and age

	Institutional			Non-institutional					
				Welfare			Non-welfare		
Sex and age	Welfare	Non-welfare	Both[1]	Disabled	Non-disabled	Both[1]	Disabled	Non-disabled	Both[1]
Males									
Total	1,813	3,257	5,070	2,389	3,076	5,465	13,131	78,269	91,400
65-69 years	269	285	554	333	882	1,215	2,768	36,782	39,550
70-74 years	318	512	830	666	809	1,475	3,086	20,096	23,182
75-79 years	313	778	1,091	575	749	1,324	2,700	12,628	15,328
80-84 years	432	577	1,009	361	464	825	2,282	6,473	8,755
85 years or over	481	1,105	1,586	454	172	626	2,295	2,290	4,585
Females									
Total	6,229	9,715	15,944	6,564	8,551	15,115	23,989	106,908	130,897
65-69 years	386	413	799	699	2,390	3,089	3,502	44,112	47,614
70-74 years	638	767	1,405	1,131	2,230	3,361	4,338	26,490	30,828
75-79 years	853	1,459	2,312	1,514	1,841	3,355	4,468	19,338	23,806
80-84 years	1,377	2,102	3,479	1,311	1,273	2,584	5,292	10,725	16,017
85 years or over	2,975	4,974	7,949	1,909	817	2,726	6,389	6,243	12,632

[1]These data columns correspond to the AAPCC categories in Table 1.

NOTE: AAPCC is adjusted average per capita cost.

SOURCE: Data from the 1984 National Long-Term Care Survey; data tabulated at the Center for Demographic Studies, Duke University.

Institutional status

Two major classifications of living quarters are *non-institutional units and institutional units*. Non-institutional units refer to single family homes, apartments, rooms in boarding houses, and similar residential arrangements. Institutional units refer to short-stay and long-stay hospitals, mental hospitals, nursing homes, convalescent homes, rest homes, homes for the aged, needy, or infirm, and correctional institutions. Estimates based on this definition of institutionalization are more comprehensive than estimates based on the 1977 and 1985 National Nursing Home Surveys (NNHS), which are restricted to registered nursing homes with three or more people (Hing, Sekscenski, and Strahan, 1989). That is, the NLTCS institutional estimates include persons in board and care facilities and, combined with the community resident sample, represent 100 percent of Medicare-eligible persons.

Among the Medicare population in the 1984 NLTCS, approximately 77 percent of institutional persons resided in nursing homes (compared with Hing, Sekscenski, and Strahan, 1989). Thus, there is potential for discrepancy between our payments for institutional persons (i.e., Schedule 2, where welfare status is used to stratify the community, but not the institutional population; and Schedule 3, where the non-institutional population is stratified by both welfare and disability status) and those of Kunkel and Powell (1981) (Schedule 1; 1981). Kunkel and Powell did not indicate the type of units included in their definition of institutionalization. Thus, to the extent that their definition had a higher proportion of nursing home residents with higher average costs, their average payment rates will be higher.

We used the U.S. Bureau of the Census definition of institutionalization both because it exhaustively decomposes the entire Medicare-eligible population into community versus institutional populations, and because this definition is used in many published Federal population statistics. Furthermore, it appears that the number of elderly residents in board and care homes is increasing and becoming a significant policy concern. It seems reasonable to include them in the institutional population.

The potential for discrepancies resulting from different definitions of institutional residence is substantially lower for the non-institutional population. This is because only 6.5 percent of the population age 65 or over was institutionalized in 1984 (with 5.0 percent in nursing homes). Because HMOs paid by the AAPCC tended to enroll non-institutional persons almost exclusively, evaluations for the first few years after enrollment need not consider issues of discrepant definitions. More generally, our payment schedules (Schedules 2 and 3) give unbiased results when used in conjunction with our definitions.

In the NLTCS, 1,953 persons were determined to be institutionalized on April 1, 1984. Of these, 1,836 were determined to be in institutions by the institutional survey instrument delivered after April 1 at a time determined by the pace of field operations. In addition, 117 persons who died between April 1, 1984, and the date of the initial effort to conduct the interview were assigned to institutional residence based on an analysis of relative mortality differentials between institutional and non-institutional persons, using a maximum likelihood model of the probability of non-response (Manton, Stallard, and Woodbury, 1991).

Institutional status was not updated on a monthly basis because of data limitations. This means that our results will differ from an HMO payment system having updating. Updating with the existing information requires a more complex analysis where institutional transition rates between, say, 1982 and 1984 are applied to change the proportions in institutions on a monthly basis. This analysis would require rates to be stable between points of observation. A similar assumption is required for the point prevalence rates to apply over time.

Welfare status

Two sources of information were used to determine welfare status. The first is based on the State buy-in code for Medicaid recorded in the Third Party Master (TPM) file for the followup period. A person was classified as on welfare if, on the first day of the month, the person was eligible for Medicaid. This is consistent with HMO payment schemes using monthly updating of welfare status. Unfortunately, this determination is only about 89 percent complete because several States do not have buy-in programs; and some States with buy-in programs do not buy in for their medically needy beneficiaries.

To reduce bias, a second determination was made of current Medicaid eligibility from the 1984 NLTCS, for both institutional persons and non-institutional disabled persons. For non-institutional persons, the broadest definition of eligibility was used to ensure finding the maximum number of welfare cases. Including these people increased our weighted population estimates of Medicaid-eligible persons from 89 percent in the TPM file to from 95 to 98 percent of the Medicaid-eligible population estimated by the U.S. Bureau of the Census (1985). When additional cases were ascertained only from the NLTCS, they were assumed to be on welfare for the entire 12-month followup period. Thus, the monthly welfare prevalence counts should be even closer to the true monthly values. Therefore, no imputation of welfare status was made. In the original AAPCC, institutional status was not broken out by welfare status. This is consistent with the small cell sizes for welfare status within the institutional population in Table 2.

Our definition of welfare as "currently Medicaid eligible" is consistent with the definition used by Kunkel and Powell (1981). In the 1985 version of the AAPCC (*Federal Register*, 1985), the term "welfare" was replaced by the term "Medicaid" in the headings of the payment schedule. Thus, in this article the two terms are viewed as equivalent.

Disability status

Disability status was determined from the NLTCS for the date April 1, 1984. This variable was not updated on a monthly basis because of data limitations. Again, it would have been possible, but more complex, to use disability transition rate estimates (e.g., Manton, 1988a, b) to simulate monthly changes. However, in analyses of disability transition rates estimated from the NLTCS for 1982 to 1984, only 10 percent of non-disabled persons become disabled over a 12-month period. Thus, the rate of change in disability status in a single month is small.

A person was classified as disabled if he or she currently received either personal or mechanical help in at least one of six ADL, i.e., eating, getting in or out of bed, getting around inside, dressing, bathing, and getting to the bathroom or using the toilet. In 1984 there were 4,157 persons so classified—3,502 on the basis of a community interview, 308 persons on the basis of a screening interview, and 347 persons imputed to be disabled by the maximum likelihood procedure. This procedure, applied to the latter two groups, was based on an analysis of mortality differentials between each group and corresponding groups with complete information (Manton, Stallard, and Woodbury, 1991).

Our choice of the ADL criterion for defining disability is consistent with HCFA's original intent in designing the 1982 NLTCS. It is a criterion that is easily applied and can be objectively verified. Nonetheless, we considered two alternatives.

The first was to expand the list of ADL and to include instrumental activities of daily living (IADL), which were obtained in the NLTCS screening instrument. The use of IADL limitations such as having difficulty shopping, cooking, doing laundry, or lifting heavy weights might introduce a sex, cultural, or size of place bias (e.g., living in a large metropolitan area in high rise apartment buildings could make outside mobility more difficult).

The second was to base the classification on original reason for entitlement to Social Security. One reason for entitlement is disability. However, this disability definition is more appropriate for workers under age 65 because it emphasizes medical condition (e.g., heart disease) instead of functioning. In addition, beneficiaries under Social Security Disability Insurance did not become entitled to Medicare until 1973. Consequently, in 1984 the oldest beneficiaries who were formerly disabled would have been about age 75, i.e., age 64 in 1973. Over time, the upper age limit for such formerly disabled Medicare beneficiaries would increase, but for the current analysis, the lack of such data is critical.

Adjusted average per capita cost factors

Our first goal was to replicate as closely as possible the original AAPCC underwriting factors with the 1984 NLTCS data. We did this independently for each of the 30 cells in Table 1. The cells in Table 1 (which were only modified by a simple scale factor) define payment Schedule 1. No attempt is made to smooth these adjusted factors, or the factors to be estimated from the 1984 NLTCS, over age or to pool cells with small numbers of cases. The original AAPCC factors were smoothed, rounded to the nearest .05, and not allowed to decrease with increasing age (Kunkel and Powell, 1981). We did not do this with our NLTCS estimates because our elderly sample was substantially larger than the CMSs for 1974-76. To utilize the NLTCS tables for ratesetting, the question of smoothing would have to be addressed.

Kunkel and Powell (1981) suggest that the AAPCC factors can be used for both prospective payment and retrospective reimbursement contracts. In developing the NLTCS-based estimates, we assumed a prospective payment contract in which all persons enrolled on the first of the month are 100 percent covered for the entire month for either Part A or Part B (or both). Thus, an HMO using our factors could enroll Medicare beneficiaries without regard to the demographic underwriting factors and be assured of fair payment. Our estimates do not provide for end stage renal disease. If an enrollee should die during the month, the HMO would not have to return part of the payment.

The 1984 NLTCS-based estimates are in Table 3. To facilitate comparisons, ratios of the entries in Table 1 to those in Table 3 are presented in Table 4. The ratios for both parts are based on weighted averages of Part A and Part B ratios, with the weights equal to the monthly payments for each part (i.e., for Part A, $104.13; for Part B, $50.34).

To test the statistical significance of differences between the two sets of underwriting factors, we used the 1984 NLTCS to compute the loss (i.e., actual Medicare cost minus AAPCC payment) for each person for the period April 1, 1984, to March 31, 1985. The AAPCC payment was the weighted average sample cost multiplied by the appropriate factor from Tables 1 or 3. To establish an initial payment, the baseline underwriting factor was set to 1.0. At any constant level of fiscal risk (probability of ruin), contingency reserves (money held aside for adverse experience) can be reduced by a fraction R^2 when an improved payment schedule is introduced; R^2 is the proportion of variance explained by the improved schedule. To calculate R^2, we conducted a weighted analysis of variance of the original AAPCC factors in Table 1. For Part A, $F(29; 22{,}639) = 4.03, p < .01, R^2 = .0051$. Although the original AAPCC factors are significant, they account for only one-half a percent of the variance in costs. For Part B, the variance explained is again substantively small ($R^2 = .0102$) even though significant ($p < .01$) because of the large sample size (i.e., the F statistic of 8.03 is based on 29 and 22,639 degrees of freedom). This level of explained variance is too small to meaningfully affect the fiscal risks of a provider.

To test the 1984 NLTCS-revised AAPCC in Table 3, we used Table 1 as the baseline. Thus, we tested residuals from the first analysis. To maintain comparability, welfare status was used as a stratum for non-institutional persons, but not for institutional persons, in payment Schedule 2. For Part A, $F(30;$

Table 3
Revised AAPCC payment factors for the aged, by sex, type of coverage, and age

	Males			Females		
Type of coverage and age	Welfare	Non-welfare	Both	Welfare	Non-welfare	Both
Institutional, Part A						
Total	1.998	2.344	2.226	1.307	1.802	1.615
65-69 years	4.334	2.348	[1]3.369	1.466	3.549	[1]2.551
70-74 years	1.088	2.071	[1]1.723	2.001	2.764	[2]2.436
75-79 years	1.442	2.277	[1]2.045	1.297	1.719	[2]1.565
80-84 years	1.690	2.867	[1]2.398	1.181	1.886	[2]1.618
85 years or over	1.776	2.236	[1]2.109	1.198	1.493	[2]1.386
Non-institutional, Part A						
Total	1.457	1.038	1.058	1.313	0.843	0.885
65-69 years	[2]1.414	[2]0.756	0.774	[2]1.097	[2]0.558	0.589
70-74 years	[2]1.675	[2]1.016	1.047	[2]0.874	[2]0.789	0.796
75-79 years	[2]1.250	[2]1.257	1.257	[2]1.275	[2]0.874	0.914
80-84 years	[2]1.371	[2]1.434	1.430	[2]1.586	[2]1.152	1.204
85 years or over	[2]1.542	[2]1.458	1.466	[2]1.990	[2]1.451	1.532
Institutional, Part B						
Total	1.662	1.894	1.816	1.777	1.714	1.738
65-69 years	2.605	1.673	[1]2.152	2.273	2.510	[1]2.397
70-74 years	1.664	1.779	[1]1.738	1.752	1.934	[1]1.856
75-79 years	1.445	2.584	[1]2.268	3.556	1.945	[1]2.532
80-84 years	1.308	1.923	[1]1.678	1.617	2.199	[1]1.978
85 years or over	1.531	1.492	[1]1.502	1.243	1.333	[1]1.301
Non-institutional, Part B						
Total	1.814	0.993	1.033	1.494	0.845	0.903
65-69 years	[2]1.585	[2]0.825	0.846	[2]1.330	[2]0.667	0.705
70-74 years	[2]2.291	[2]1.002	1.064	[2]1.479	[2]0.862	0.912
75-79 years	[2]1.411	[2]1.095	1.115	[2]1.631	[2]0.883	0.956
80-84 years	[2]2.076	[2]1.204	1.263	[2]1.642	[2]1.031	1.104
85 years or over	[2]1.532	[2]1.265	1.291	[2]1.353	[2]1.036	1.084

[1]Entries used in payment Schedules 2 and 3.
[2]Entries used in payment Schedule 2 only.

NOTE: AAPCC is adjusted average per capita cost.

SOURCE: Data from the 1984 National Long-Term Care Survey; data tabulated at the Center for Demographic Studies, Duke University.

22,639) = 4.90, $p < .01$, $R^2 = .0065$; for Part B, $F(30; 22{,}639) = 2.64$, $p < .01$, $R^2 = .0035$. Thus, the revision is significantly better than the original, though the total percentage of explained variance is small: 1.16 percent for Part A, 1.36 percent for Part B. In interpreting the F statistics, the reader is cautioned that the loss distribution is skewed to the right and is non-normal. If the skewness is the same from cell to cell, the effect on the F test is reduced. Later, we examine the issue of normality in detail.

There are several differences between Tables 1 and 3. In Table 3, the institutional population is subdivided on welfare status. This was not done in the original AAPCC which had less experience for the aged population. There are differences in institutional Medicare expenditures depending upon welfare status. For example, there is a 15-percent decrease in Medicare Part A expenses and a 12-percent decrease in Part B expenses over all ages for males in institutional residence on welfare. For females, there is a 27-percent decrease for Medicare Part A and a 4-percent increase for Part B.

Analyses of nursing home use suggest that this is not surprising in that shorter term institutional stays are often funded by Medicare (under the SNF benefit), or are privately paid. Short-stay persons would probably have higher Medicare expenses because of being more medically acute (Liu, Manton, and Liu, 1990). Medicaid nursing home residents have longer stays and, often, less medically intensive needs. Although the Medicaid payments are not part of the monies available to the HMOs, such persons have particular health characteristics and, consequently, acute care costs for Medicare under their dual eligibility. We would expect persons in Medicaid nursing homes to be less expensive in terms of acute medical care use (Hing, Sekscenski, and Strahan, 1989), e.g., they would leave the nursing home less often for an acute hospital stay. Ignoring such factors loses information on the acute medical service use of institutional persons. Failure to recognize the higher medical acuity of certain institutional populations could provide disincentives to acute and rehabilitative use of nursing homes.

Tables 1 and 3 exhibit different trends of payments with age for persons in institutions (Table 4). In Table 1, payments are relatively constant over age. In contrast, in Table 3 these are lower than average cost factors for those age 85 or over for Parts A and B for both sexes. Table 4 shows that the original AAPCC rate schedule now appears to overcompensate for the institutional population age 85 or over. It also seems to underpay for institutional persons at earlier ages,

Table 4

Ratio of original AAPCC payment factors in Table 1 to revised AAPCC payment factors in Table 3, by institutional status, type of coverage, sex, and age

Type of coverage, sex, and age	Institutional	Non-institutional	
		Welfare	Non-welfare
Part A, hospital insurance			
Males:			
65-69 years	0.632	0.991	0.962
70-74 years	1.295	0.961	0.818
75-79 years	1.193	1.620	0.826
80-84 years	1.018	1.742	0.869
85 years or over	1.157	1.751	0.961
Females:			
65-69 years	0.672	0.852	1.116
70-74 years	0.810	1.367	0.921
75-79 years	1.459	1.222	1.069
80-84 years	1.412	1.178	0.992
85 years or over	1.648	1.122	0.895
Part B, supplementary medical insurance			
Males:			
65-69 years	0.831	0.774	1.053
70-74 years	1.117	0.625	1.020
75-79 years	0.856	1.122	1.027
80-84 years	1.157	0.837	0.976
85 years or over	1.292	1.134	0.929
Females:			
65-69 years	0.661	0.845	1.073
70-74 years	0.881	0.794	0.949
75-79 years	0.686	0.783	1.100
80-84 years	0.878	0.778	0.991
85 years or over	1.336	0.944	1.035
Parts A and B			
Males:			
65-69 years	0.697	0.920	0.992
70-74 years	1.237	0.851	0.884
75-79 years	1.083	1.458	0.891
80-84 years	1.063	1.447	0.904
85 years or over	1.201	1.550	0.951
Females:			
65-69 years	0.668	0.849	1.102
79-74 years	0.833	1.180	0.930
75-79 years	1.207	1.079	1.079
80-84 years	1.238	1.048	0.991
85 years or over	1.546	1.064	0.941

NOTE: AAPCC is adjusted average per capita cost.

SOURCE: Center for Demographic Studies, Duke University.

especially for females, where medical interventions, rehabilitation, and deinstitutionalization are more likely than at advanced ages (Manton, 1988a, b).

The cross classification of institutional persons by age, sex, and welfare status produces the smallest cell sizes in Table 2, especially for younger males. These cells exhibit the greatest variation from one age to the next, likely a result, in part, of small cell sizes. Consequently, in subsequent analyses (using Schedules 2 and 3, discussed later) the institutional population is not stratified by welfare status. This is consistent with the original AAPCC (Table 1) and is a reasonable way to increase cell sizes. It does not mean that welfare status is unimportant in predicting costs for institutional persons.

In examining the non-institutional population, we find that certain trends hold in Tables 1 and 3. In particular, for both sexes the payment ratios for persons not on welfare show a rough similarity for both Parts A and B, with the largest discrepancies for males age 70-84 for Part A (Table 4).

For non-institutional males age 75 or over on welfare, the Part A payment levels are substantially higher in the original than in the 1984 NLTCS AAPCC, whereas they are closer for Part B. This may reflect effects of PPS on Medicare in shortening hospital stays and increased use of Part B, post-acute care (e.g., home health), and outpatient services to compensate (Ruther and Reilly, 1989). The differences for females are consistent with this interpretation. For non-institutional females age 70 or over on welfare, the payments in the original AAPCC are from 12 percent to 37 percent higher for Part A but 6 percent to 22 percent lower for Part B. This is also consistent with increases in home health services use from 1982 to 1984.

Thus, changes in the patterns of service use are consistent with documented trends in the pattern of Medicare service use induced by PPS and other Medicare policy changes.

Disability factors

In the second set of calculations the non-institutional population is divided into disabled and non-disabled groups based on chronic limitations (90 days or more) in one or more ADL; i.e., individuals with only one (or more) of six ADL limitations are considered disabled.

The estimated underwriting factors with disability included as a factor for the non-institutional population are in Table 5. Payment Schedule 3 is defined by using these factors to replace the non-institutional factors in Schedule 2.

To test the statistical significance of this simple disability dichotomy, we used Schedule 2 (with institutional status not broken out by welfare status) as the baseline. Thus, we are testing residuals from Table 3. For Part A, $F(20; 22{,}619) = 27.97, p < .01$, $R^2 = .0241$; for Part B, $F(20; 22{,}619) = 18.65, p < .01$, $R^2 = .0162$. The total explained variance is 3.54 percent for Part A and 2.96 percent for Part B. We conclude that the disability factor for non-institutional persons is significant even given the other four factors.

In Table 5, there is more than a threefold difference in Medicare Part A payments over disability status (e.g., for males the total ratio is $3.15 = 2.648 \div 0.840$; for females it is $3.74 = 2.336 \div 0.625$). The differentials are larger for Part A than for Part B (e.g., about 2.36 for males and 2.44 for females). Differentials hold for males and females. The differential is greater at younger ages and is generally larger for non-welfare persons than for welfare persons. As there are many more non-disabled than disabled persons (about 5.7 to 1) the exclusion of a small number

Table 5

Extended AAPCC payment factors with disability included as an underwriting factor for the non-institutional population, by sex, type of coverage, and age

	Males			Females		
Type of coverage and age	Welfare[1]	Non-welfare[1]	Both	Welfare[1]	Non-welfare[1]	Both[1]
Non-institutional disabled, Part A						
Total	2.428	2.688	2.648	2.141	2.385	2.336
65-69 years	2.401	2.523	2.509	1.874	2.303	2.234
70-74 years	3.310	2.459	2.611	1.060	2.724	2.399
75-79 years	2.033	3.001	2.847	2.256	2.424	2.385
80-84 years	2.128	2.984	2.877	2.579	2.166	2.244
85 years or over	1.866	2.486	2.377	2.500	2.345	2.378
Non-institutional non-disabled, Part A						
Total	0.870	0.838	0.840	0.881	0.605	0.625
65-69 years	1.067	0.645	0.655	0.914	0.445	0.468
70-74 years	0.704	0.880	0.874	0.814	0.601	0.616
75-79 years	0.842	1.013	1.005	0.755	0.654	0.661
80-84 years	0.939	1.076	1.068	0.936	0.847	0.855
85 years or over	0.873	0.943	0.939	1.203	0.866	0.901
Non-institutional disabled, Part B						
Total	2.818	1.960	2.090	1.952	1.767	1.805
65-69 years	2.382	2.022	2.062	2.188	1.950	1.989
70-74 years	3.741	2.245	2.513	2.283	2.167	2.190
75-79 years	2.015	1.867	1.890	2.151	1.713	1.815
80-84 years	4.015	1.911	2.173	1.787	1.712	1.726
85 years or over	1.876	1.669	1.706	1.620	1.482	1.512
Non-institutional non-disabled, Part B						
Total	1.206	0.876	0.887	1.255	0.703	0.741
65-69 years	1.304	0.750	0.762	1.128	0.584	0.611
70-74 years	1.430	0.885	0.903	1.220	0.735	0.768
75-79 years	1.098	0.987	0.992	1.355	0.765	0.809
80-84 years	0.970	1.040	1.036	1.547	0.826	0.896
85 years or over	0.822	1.062	1.050	0.942	0.744	0.765

[1]Entries in these columns used in payment Schedule 3, along with entries indicated in Table 3.

NOTE: AAPCC is adjusted average per capita cost.

SOURCE: Center for Demographic Studies, Duke University.

of disabled persons could greatly affect the payment to an HMO (Table 7).

The original AAPCC does not adjust for health status, on the grounds that there is no generally used insurance system for assessing health status that could be used in the AAPCC (although with the emergence of private LTC insurance this may no longer be the case); and, there could be substantial costs associated with the administration of such a system. The statistical tests of the ADL disability measure indicate that a significant improvement in payment can be made using a very simple and inexpensive indicator. Given that the R^2 statistics appear "low," however, we need to assess them.

Tolley and Manton (1985) indicate that approximately 40 percent of Medicare costs are attributable to mortality processes, with about 30 percent incurred in the last year of life. To translate these into R^2 statistics, we classified our sample into survivors and non-survivors according to their vital status on April 1, 1985. Using payment Schedule 3, we calculated annual Part A losses of $4,002.60 per capita for non-survivors and profits (negative losses) of $219.11 per capita for survivors. Accounting for within-group variances, these differentials yielded $R^2 = .0711$. For Part B, the corresponding loss was $922.00 for non-survivors and profit of $50.47 for survivors, with $R^2 = .0212$. On the basis of the size of these estimates (i.e., 7.1 percent for Part A; 2.1 percent for Part B) one could conclude that R^2 values in the range of 2 percent to 7 percent are "high," i.e., the R^2 values associated with the test for Table 5 are substantively meaningful, and not statistically significant simply because of the large sample size of the 1984 NLTCS.

Simulations

It is difficult to assess the differences between the payment factors in Tables 1, 3, and 5 because of the number of factors involved and because the sample size associated with each factor is different. To deal with this, we conducted computer simulations to demonstrate the dollar impact of the alternative schedules on the distribution of HMO losses, under the assumption that HMO enrollment is open and

Table 6

Simulated HMO loss distributions based on original AAPCC underwriting factors in Schedule 1 and Table 1

Power of 2	HMO population size	Mean	Standard deviation	50th percentile	95th percentile	99th percentile	Percent Pr(loss ≤ 0)	Percent Kolmogorov D_n
Part A, hospital insurance								
0	1	$0.00	$3,566.65	-$912.79	$6,347.12	$16,271.62	80.832	36.148
1	2	0.00	2,521.77	-912.04	4,946.68	10,645.30	73.009	24.957
2	4	0.00	1,783.08	-661.23	3,594.62	6,901.34	66.755	17.078
3	8	0.00	1,260.80	-351.49	2,501.21	4,345.43	61.751	11.849
4	16	0.00	891.50	-181.19	1,702.70	2,814.23	58.363	8.393
5	32	0.00	630.38	-91.62	1,164.11	1,858.27	55.881	5.895
6	64	0.00	445.74	-46.15	801.80	1,242.24	54.172	4.182
7	128	0.00	315.18	-23.38	554.91	837.54	52.979	2.986
8	256	0.00	222.87	-11.80	385.56	570.90	52.118	2.123
9	512	0.00	157.59	-5.93	268.92	392.86	51.504	1.506
10	1,024	0.00	111.43	-2.98	188.22	272.38	51.065	1.071
11	2,048	0.00	78.79	-1.49	132.09	189.87	50.753	0.754
12	4,096	0.00	55.72	-0.74	92.90	132.90	50.539	0.543
13	8,192	0.00	39.40	-0.37	65.44	93.29	50.377	0.379
14	16,384	0.00	27.86	-0.19	46.14	65.63	50.267	0.273
15	32,768	0.00	19.70	-0.09	32.56	46.23	50.188	0.194
16	65,536	0.00	13.93	-0.05	22.99	32.61	50.135	0.139
17	131,072	0.00	9.85	-0.02	16.24	23.01	50.099	0.103
18	262,144	0.00	6.96	-0.01	11.48	16.25	50.069	0.079
19	524,288	0.00	4.92	-0.01	8.11	11.48	50.050	0.053
20	1,048,576	0.00	3.48	0.00	5.73	8.11	50.038	0.045
Part B, supplementary medical insurance								
0	1	$0.00	$1,498.03	-$431.29	$2,274.74	$5,406.46	76.219	27.641
1	2	0.00	1,058.23	-335.88	1,717.77	3,587.11	70.233	22.730
2	4	0.00	747.78	-210.17	1,211.11	2,382.10	65.069	18.162
3	8	0.00	528.68	-117.90	850.39	1,622.30	61.259	13.225
4	16	0.00	373.81	-66.57	597.66	1,091.24	58.782	9.462
5	32	0.00	264.31	-39.13	420.14	778.59	57.118	7.952
6	64	0.00	186.89	-23.76	299.19	590.21	55.975	6.857
7	128	0.00	132.15	-14.94	217.95	432.03	55.197	6.007
8	256	0.00	93.44	-9.56	159.77	305.07	54.570	5.182
9	512	0.00	66.07	-5.99	116.51	198.01	53.935	4.289
10	1,024	0.00	46.72	-3.56	82.42	129.58	53.191	3.351
11	2,048	0.00	33.04	-1.95	57.44	86.90	52.413	2.463
12	4,096	0.00	23.36	-1.01	40.05	59.25	51.734	1.767
13	8,192	0.00	16.52	-0.51	28.00	40.83	51.248	1.252
14	16,384	0.00	11.68	-0.26	19.64	28.35	50.876	0.892
15	32,768	0.00	8.26	-0.13	13.80	19.80	50.623	0.630
16	65,536	0.00	5.84	-0.06	9.71	13.88	50.442	0.449
17	131,072	0.00	4.13	-0.03	6.85	9.75	50.307	0.317
18	262,144	0.00	2.92	-0.02	4.83	6.86	50.217	0.231
19	524,288	0.00	2.06	-0.01	3.41	4.84	50.155	0.160
20	1,048,576	0.00	1.46	0.00	2.41	3.41	50.109	0.116

See footnotes at end of table.

representative. This latter assumption is central to the validity of the AAPCC (Kunkel and Powell, 1981).

We use three sets of payment factors: Schedule 1 = Table 1; Schedule 2 = Table 3, with welfare status stratifying the non-institutional population, but not the institutional population; and Schedule 3 = Table 5, with welfare status and disability status stratifying the non-institutional population, and with institutional population factors the same as for Schedule 2. These are the same payment schedules discussed previously.

For each population subgroup, we independently simulated effects for each payment schedule. First, we computed the loss, i.e., actual Medicare cost minus AAPCC paid cost, for each person in that subgroup for the period April 1, 1984, to March 31, 1985. Second, we sorted these losses by their size and attached a probability proportional to the sample weight to each loss. Third, we grouped the losses into class intervals of $10 each for Part A, $5 each for Part B, and $10 each for both parts (i.e., the sum of Parts A and B), with each interval represented by the within-interval probability-weighted average loss. This procedure yielded an empirical distribution function (EDF) for each type of loss. For Part A there were more than 1,400 class intervals; for Part B, about 1,200 class intervals; and for both parts, nearly 1,800 class intervals.

Table 6—Continued

Simulated HMO loss distributions based on original AAPCC underwriting factors in Schedule 1 and Table 1

Power of 2	HMO population size	Mean	Standard deviation	50th percentile	95th percentile	99th percentile	Percent Pr(loss $\leq$ 0)	Percent Kolmogorov D_n
Parts A and B								
0	1	$0.00	$4,555.64	−$1,359.44	$8,290.44	$20,903.92	78.522	30.128
1	2	0.00	3,221.21	−1,231.60	6,329.85	13,315.37	71.483	21.847
2	4	0.00	2,277.70	−783.56	4,580.86	8,501.37	65.875	16.087
3	8	0.00	1,610.55	−425.16	3,148.43	5,457.39	61.212	11.257
4	16	0.00	1,138.82	−220.40	2,148.60	3,594.85	57.971	7.990
5	32	0.00	805.26	−111.92	1,480.51	2,381.03	55.654	5.657
6	64	0.00	569.40	−57.31	1,022.91	1,584.40	54.066	4.067
7	128	0.00	402.62	−29.17	708.05	1,067.46	52.913	2.915
8	256	0.00	284.70	−14.74	492.02	727.84	52.074	2.074
9	512	0.00	201.31	−7.39	343.24	501.09	51.464	1.475
10	1,024	0.00	142.35	−3.70	240.28	347.54	51.044	1.045
11	2,048	0.00	100.65	−1.85	168.65	242.35	50.737	0.738
12	4,096	0.00	71.17	−0.92	118.63	169.66	50.520	0.527
13	8,192	0.00	50.33	−0.46	83.56	119.12	50.366	0.371
14	16,384	0.00	35.59	−0.23	58.93	83.81	50.260	0.268
15	32,768	0.00	25.16	−0.12	41.59	59.05	50.183	0.187
16	65,536	0.00	17.79	−0.06	29.37	41.65	50.130	0.136
17	131,072	0.00	12.58	−0.03	20.74	29.40	50.094	0.096
18	262,144	0.00	8.90	−0.01	14.66	20.76	50.066	0.071
19	524,288	0.00	6.29	−0.01	10.36	14.67	50.052	0.050
20	1,048,576	0.00	4.45	0.00	7.32	10.36	50.034	0.041

NOTES: HMO is health maintenance organization. AAPCC is adjusted average per capita cost.

SOURCE: Center for Demographic Studies, Duke University.

EDFs of HMO losses were generated for different size HMO enrollments from individual level EDFs obtained from the 1984 NLTCS. This was done by cumulating the discrete distribution functions.

Table 6 illustrates the simulations using Schedule 1. The extent of these tables means that we had to select small sections of the simulation to compare in Table 7. All results are normalized to a per capita basis and HMO sizes increase as powers of 2, e.g., $2^{10} = 1,024$. By using standard variance formulas, one can get exact results for any size HMO from 1 to 2,097,151 persons. In addition, the Kolmogorov D_n statistic (Kendall and Stuart, 1973) indicates the maximum discrepancy between each distribution and an approximating normal distribution with the same mean and standard deviation. For Part A and total costs, the discrepancy is less than 1 percent for HMOs with 2,048 persons or more. For Part B, the 1-percent discrepancy requires an HMO with 16,384 or more persons.

In each case, the mean loss is zero dollars, and the standard deviation decreases by $\sqrt{2}$ as one increases HMO size. The 50th percentile starts negative and converges rapidly to zero. The 95th and 99th percentiles are useful for determining security loadings (i.e., the amount of additional premium necessary to compensate for losses) at these levels of confidence. For example, the average total payment is $1,853.64 (i.e., 12 x $104.13 + 12 x $50.34). For both Part A and Part B, for HMOs with 16,384 clients, the 99th percentile is $83.81, implying a 4.52-percent security loading. For HMOs with larger numbers of clients, where the distributions are approximately normal, the locations of the 95th and 99th percentiles decrease by $\sqrt{2}$ as the HMO population size increases.

The standard deviations are also useful for developing approximate standard errors for payment factors in Tables 3 and 5. This is because the standard deviations for the underwriting classes in Tables 3 and 5 are roughly equivalent to the standard deviations for the same sized cells in Table 6. This can be seen in Table 7 by comparing each entry with the "Total" entry in the standard deviation column. Hence, from Table 6, for HMOs with 128 clients, the standard deviation for Part A is $315.18, which yields a coefficient of variation of 25.2 percent (i.e., $315.18/12 x $104.13). Because 128 clients corresponds to 1,536 person-months in Table 2, estimates in Tables 3 and 5 for that size cell should have standard errors of approximately 25 percent of their values. The Kolmogorov D_n statistic is 3.0 percent, indicating that estimates for cells this size or larger are close to normally distributed. These standard errors decrease by $\sqrt{2}$ as the cell size in Table 2 doubles. For smaller cells, normality fails to hold and standard errors can be misleading.

Tables (not shown) similar to Table 6 were computed for payment Schedules 2 and 3. In both cases, the mean loss was always zero dollars. The standard deviations were reduced from 1 percent to 2 percent, as expected from the R^2 statistics previously mentioned.

Schedules 2 and 3 have the property that the mean loss for each cell formed from the underwriting factors is zero. This means that any mixture of cells can be combined to form a synthetic enrolled population, with the appropriate payment provided by the associated schedule. This is true as long as there is representative and open enrollment within each cell.

The same property of unbiasedness is not true for Schedule 1. This is because our initial rescaling of the

Kunkel and Powell (1981) factors ensures unbiasedness for the entire 1984 NLTCS sample, but not for subgroups. The ratios in Table 4 indicate that substantial discrepancies may exist for any given cell. To investigate this, we decomposed the loss distributions for both Parts A and B in Table 6 (and similar distributions for the other two payment schedules) by sex, age, and institutional status. Within the non-institutional population, we further decomposed these loss distributions by welfare status and disability status. For the original AAPCC factors, we also classified the 30 underwriting cells into low, medium, and high payment classes (10 cells each) based on the ratios for both Parts A and B in Table 4. In all cases, the classification was based on a person's status on April 1, 1984. For those underwriting factors that change during the year (age and welfare status), the unbiased mean loss may differ from zero because of transfer to more or less favorable payment classes. The results of these calculations for HMOs with 16,384 clients (a medium size HMO; U.S. Bureau of the Census, 1986) are presented in Table 7. Results for other HMO sizes can be obtained using the $\sqrt{2}$ adjustment indicated above, after subtracting out the mean.

Table 7 confirms our expectation of bias in Schedule 1. In these results, a negative loss represents an HMO profit. For four underwriting classes (females, age 85 or over, institutional, and non-institutional welfare), the distribution is biased with even the 99th percentile being negative. This implies that if HMOs could enroll only those persons, they would almost never suffer a loss. For two other underwriting classes (males and age 70-74), the distribution is biased in the opposite direction, i.e., HMOs that enrolled only those persons would almost always suffer a loss. The most extreme bias among current factors is for institutional persons (the mean profit is $563.34 per person year). For all of these underwriting factors, the probability of loss diverges toward either 0 percent or 100 percent. These results can be contrasted with the unbiased results from Schedules 2 (except for disability status) and 3. This suggests that differentials in the mixture of these different groups in an HMO do not have to be very large to create a relatively large loss or profit.

One underwriting factor of particular interest, because of its size, is non-institutional non-welfare persons. Persons in the cells of this factor tend to be overrepresented in the enrollment of HMOs of the group or staff variety (e.g., Brown, 1988) and would appear to be the desired audience for marketing HMO services. Surprisingly, for Schedule 1, the cells in this class have an average per capita loss of $59.36 per year, with a 99th percentile maximum of $140.68. The probability of a loss is 95.9 percent. The mean security loading rate (denominator = $1,853.64) is 3.2 percent and the 99-percent security loading rate is 7.6 percent. For HMOs of other sizes, the mean security loading rate is constant, but the 99-percent rate changes. With 4,096 clients, the 99-percent rate is 12.1 percent; with 262,144 clients, the 99-percent rate is 4.3 percent. At no point will the rate fall below the mean rate of 3.2 percent because of the bias in the payment schedule. Of course, the effect of this loss is modified if there is differential enrollment of specific cells associated with this factor. For example, males have a $107.59 loss. So, the total loss for a non-institutional, non-welfare male is much higher than the average. Females show a $70.22 profit that would more than balance the $59.36 loss for the factor. Losses occur only for persons age 70 to 74 with profits manifest for both older and younger enrollees. Thus, the loss for this group could be accommodated if younger or older persons were differentially enrolled.

Biases also arise for the cells assigned to the low, medium, or high payment groups of ratios in Table 4. These results illustrate the maximum cost effects that could occur with systematic enrollment biases. If an HMO could enroll exclusively from the 10 high payment cells, this would yield an average profit of $305.26 per person year. The 10 low payment cells yield an average loss of $197.04 per person year. Of course, recruitment from only the top or bottom 10 cells is unlikely for an HMO but, given the size of the differences, even recruitment of only several percent more than the proportion in the populations of the high payment group could lead to fairly large financial savings for the HMO.

The contrast between disabled and non-disabled non-institutional statuses is, strictly speaking, not a bias because disability is not an underwriting factor in Schedules 1 or 2. It is a characteristic of enrollees that can have major cost consequences. If only non-disabled (non-institutional) persons enrolled, the HMO would have an expected profit of $300.14 per person year, only $5.12 less than for the high payment cell selection. For HMOs there is a strong incentive not to enroll the non-institutional disabled because the loss associated with such a person is by far the highest in the table, i.e., $1,956.91. Thus, an HMO would only have to overenroll or underenroll a small number in this group to have a sizable financial effect.

HMOs operating under the Tax Equity and Fiscal Responsibility Act of 1982 (Public Law 97-248) are not allowed to screen enrollees (except in rare instances in certain demonstrations; Brown, 1988) so that, with geographic adjustments that account for unusual demographics, it would not be possible for an HMO to consciously use these biases to their advantage by law. However, it is permissible for an HMO to select one of several organizational forms that may lead to a "passive" selection bias. For example, Brown (1988) showed (as did earlier demonstrations) that HMOs based on an individual practice association (IPA) model tended to enroll an unbiased group of clients, but staff or group practice HMOs tended to enroll a population that was strongly biased toward being healthy. This is because persons with diagnosed ailments, and with established relations with caregivers, are unlikely to drop their current providers to join a staff or group HMO. Because the physician and his existing practice are recruited in IPA plans, IPA HMOs are not subject to this selection bias (Brown, 1988). In staff or group HMOs, individuals are recruited, and they tend to be differentially healthier because they are not yet

Table 7

Simulated HMO loss distributions for Medicare Parts A and B for HMOs with 16,384 clients drawn from select underwriting classes, under three alternative payment schedules

Schedule and underwriting class	Mean	Standard deviation	50th percentile	95th percentile	99th percentile	Percent Pr(loss ≤ 0)	Percent Kolmogorov D_n
Schedule 1							
Total	$0.00	$35.59	−$0.23	$58.93	$83.81	50.260	0.268
Males	107.59	37.69	107.36	169.97	196.26	0.188	0.238
Females	−70.22	34.13	−70.46	−13.68	10.22	97.910	0.287
Age 65-69	−21.61	31.64	−21.85	30.86	53.08	75.390	0.313
Age 70-74	147.81	37.39	147.56	209.73	235.89	0.000	0.270
Age 75-79	−34.61	35.53	−34.80	24.16	48.88	83.507	0.222
Age 80-84	−20.91	39.61	−21.18	44.71	72.45	70.292	0.279
Age 85 or over	−232.16	37.58	−232.33	−170.07	−144.01	100.000	0.186
Institutional	−563.34	44.12	−563.58	−490.37	−459.65	100.000	0.222
Non-institutional	36.86	34.94	36.63	94.72	119.15	14.552	0.275
Welfare	−252.81	39.95	−253.02	−186.74	−158.94	100.000	0.217
Non-welfare	59.36	34.51	59.13	116.53	140.68	4.147	0.272
Disabled[1]	1,956.91	53.59	1,956.64	2,045.51	2,082.75	0.000	0.203
Non-disabled[1]	−300.14	29.74	−300.34	−250.88	−230.07	100.000	0.272
Payment level:							
Low[2]	197.04	37.44	196.82	259.00	285.11	0.000	0.243
Medium[2]	0.24	36.37	0.00	60.48	85.92	50.002	0.272
High[2]	−305.26	30.85	−305.48	−254.14	−232.52	100.000	0.295
Schedule 2							
Total	$0.00	$35.50	−$0.23	$58.79	$83.62	50.260	0.266
Males	0.00	37.66	−0.23	62.31	88.58	50.236	0.243
Females	0.00	34.02	−0.24	56.37	80.21	50.278	0.289
Age 65-69	−14.74	31.62	−14.98	37.68	59.89	68.151	0.311
Age 70-74	17.26	37.39	17.01	79.19	105.34	32.409	0.276
Age 75-79	−7.61	35.48	−7.80	51.07	75.77	58.692	0.223
Age 80-84	−8.02	39.53	−8.29	57.46	85.14	58.297	0.276
Age 85 or over	27.12	37.13	26.95	88.48	114.24	23.319	0.190
Institutional	0.00	43.49	−0.25	71.93	102.22	50.220	0.223
Non-institutional	0.00	34.92	−0.23	57.83	82.25	50.264	0.265
Welfare	−86.81	39.70	−87.02	−21.14	6.50	98.486	0.221
Non-welfare	6.75	34.51	6.51	63.92	88.07	42.515	0.278
Disabled[1]	1,937.68	53.51	1,937.42	2,026.16	2,063.35	0.000	0.204
Non-disabled[1]	−340.09	29.72	−340.29	−290.87	−270.07	100.000	0.273
Schedule 3							
Total	$0.00	$35.03	−$0.22	$58.01	$82.49	50.246	0.264
Males	0.00	37.26	−0.21	61.66	87.64	50.226	0.236
Females	0.00	33.50	−0.23	55.49	78.95	50.276	0.280
Age 65-69	−14.38	31.22	−14.61	37.36	59.27	67.958	0.306
Age 70-74	19.08	36.90	18.84	80.18	105.98	30.424	0.267
Age 75-79	−1.01	35.01	−1.20	56.89	81.25	51.369	0.218
Age 80-84	−10.02	38.99	−10.28	54.57	81.87	60.382	0.278
Age 85 or over	10.38	36.61	10.22	70.88	96.27	38.993	0.186
Institutional	0.00	43.49	−0.24	71.94	102.22	50.221	0.218
Non-institutional	0.00	34.41	−0.22	56.97	81.02	50.254	0.260
Welfare	−101.72	38.90	−101.92	−37.39	−10.33	99.513	0.210
Non-welfare	7.90	34.03	7.68	64.26	88.06	41.053	0.271
Disabled	0.00	54.06	−0.26	89.36	126.91	50.191	0.197
Non-disabled	0.00	29.64	−0.20	49.11	69.85	50.269	0.282

[1]Not an underwriting factor in this schedule.

[2]Cuts across all underwriting factors in this schedule, but not itself an underwriting factor.

NOTE: HMO is health maintenance organization.

SOURCE: Center for Demographic Studies, Duke University.

associated with a specific caregiver. Of equal concern was that Brown (1988) showed that there were disenrollment biases that were in the same direction as enrollment biases, so that any "regression to the mean" effect over time was mitigated.

Other possible "passive" sources of selection bias involve the mixture of services and medical specialties that the HMO chooses to offer and how those services are marketed and advertised. Again, this is a passive source of selection because it is the clients' choice to enroll or not to enroll. A final source of possible biased selection can occur when an HMO opens in an area with an existing high market penetration of HMOs, e.g., Minneapolis. In this case, the selection biases may be adverse, i.e., the remaining pool of potential clients may be less healthy on average.

The implications of these results are threefold. First, for new HMOs that, for some reason, over-enroll non-institutional, non-disabled persons, there will be substantial unearned profits in the first few years of operation if the enrollment is large enough to minimize the risk of random loss. These profits are "unearned" because they are not related to the HMO's level of operating efficiency. Second, as the enrolled population increases in average length of enrollment, it will tend to take on the characteristics of the general population (except for sex, which is readily adjustable to an unbiased payment). The unearned profits will decrease and, even if the operating efficiency of the HMO improves, it may not be to a level that compensates for the loss of unearned profits. Depending on the rapidity and extent of these changes, and the component of profit or loss resulting from random variation in costs, this process could result in a loss of confidence in the management of the HMO and threaten its financial stability. Third, this latter effect of regression to the mean may be partly or wholly countered by disenrollment bias of the same direction as the selection biases in staff model HMOs (Brown, 1988).

Summary

The findings demonstrate that the use of more recent underwriting factors from the 1984 NLTCS (as compared with the 1974-76 CMS), and of an underwriting factor for disability, can explain variations in Medicare costs more effectively. For example, we found, as documented in other studies, a shift in expenditures to Medicare Part B (e.g., use of outpatient surgery for cataract operations). This shift is apparently a result of the effects of PPS in decreasing hospitalization (and Part A expenditures). The shift to Part B has fostered research on systems to pay for ambulatory care on a prospective basis (Averill et al., 1990).

We found that there were considerable differences in acute care expenditures over a welfare factor defined for persons in institutions—that persons not on Medicaid have higher Medicare expenditures. This is consistent with the presence of long- and short-stay nursing home populations and their different levels of payment for acute health service needs (Keeler, Kane, and Soloman, 1981; Liu and Manton, 1990).

Not surprisingly, we also found large differences in expenditures over disability status. Using the measures in Table 5, this differential was approximately 3 to 1 with, interestingly, the oldest old having the lowest differential between the disabled and non-disabled groups. Thus, if HMOs have passively induced selection biases against disability (e.g., by selecting a staff model for their HMO) they may have undue profits. Furthermore, in constructing the underwriting factors, a number of methodological issues became apparent (e.g., non-response bias; date of measurement of characteristics; degree of statistical precision of underwriting factors) that could adversely affect the underwriting factors if unadjusted in the calculations.

In the AAPCC calculations we restricted ourselves, in representing health and functional status, to a single dichotomous categorical variable. If the threshold for disability were set higher (e.g., in certain congressional proposals eligibility for LTC benefits required impairment in two of five selected ADL), then the cost differences between disabled and non-disabled persons would be greater, and it would presumably be more difficult for a person to incorrectly claim disability (i.e., the impaired status would be more physically manifest). Alternatively, one could use a risk scoring system (Cummins et al., 1983) where continuous variation in risk is represented.

The use of "fuzzy sets" to conduct risk scoring, an area of current research by the Society of Actuaries (1991), could greatly increase the predictability of costs and reduce risks (Tolley and Manton, 1991). The explanatory power of such risk scoring using a combination of health and functional factors has been demonstrated both in analyses of the 1984 NLTCS (Manton, Stallard, and Woodbury, 1991) and in preliminary analyses of data from the S/HMO evaluations. The ability to manipulate such scores requires changes in multiple objectively verifiable characteristics that can be reviewed. It has been shown (Weiner and Hanley, 1989) that such disability items can be objectively and reliably determined in surveys (e.g., National Medical Expenditure Survey (NMES) and NLTCS for community populations; NNHS, NMES, and NLTCS for institutional populations). It was found feasible in the operation of the S/HMO to conduct such health screening at entry to the program. Health screening is used to determine eligibility for LTC services (e.g., if "nursing home certifiable") in the S/HMO and other programs such as PACE. Thus, research on the measurement of health and functional status of elderly persons, and on payment mechanisms that are adjusted for those characteristics, is important in developing payment systems in a number of areas.

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