Use of Outpatient Drugs as Death Approaches

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This article explores changes in outpatient prescription drug use up to 72 months prior to death and relates the findings to trends in Medicare-covered services during the same life stage. The study sample comprises 5,261 decedents who, prior to their deaths, had enrolled in the Pennsylvania Pharmaceutical Assistance Contract for the Elderly (PACE) program. Descriptive time-series show steady increases in both outpatient drug use and physician contacts in the final 36 months of life. However, multivariate analysis shows that impending death is associated with significant reductions in the probability of using outpatient drugs. Only in the final 12 months of life is this effect offset by rising numbers of drug claims by prescription users.

INTRODUCTION

The use of health services in the year or two preceding death has been the subject of numerous studies. Most rely on Medicare claims data, thereby limiting the focus to services and expenditures covered by the Medicare program (Lubitz and Prihoda, 1984; McCall, 1984; Riley et al., 1987; Riley and Lubitz, 1989; Gaumer and Stavins, 1992). Although use of nursing home care—arguably Medicare's most significant excluded benefit—has received limited attention in the use-preceding-death literature (Roos, Montgomery, and Roos, 1987; Scitovsky, 1988; Temkin-Greener et al., 1992), other Medicare exclusions such as outpatient prescription drugs have not been studied in this context.

Here we investigate patterns in outpatient pharmaceutical utilization for aged individuals during their final months of life and compare the results with utilization of Medicare-covered services during the same period. The research is driven by two aims. First is to fill an obvious gap in the literature. Despite the Medicare exclusion, more than 80 percent of the elderly fill at least one prescription per year (Moeller and Mathiowetz, 1989). This represents a higher prevalence of use than for any Medicare-covered service, including physician services (Helbing, 1993). By studying drug use prior to death, we add a potentially important element to our understanding of life-stage health care utilization patterns. By comparing prescription use with other health care services we open new avenues of research into the complex interrelationships that govern Medicare expenditures. The link between ambulatory physician contacts and legend drugs is an obvious one in this regard, but other Medicare Part A and Part B services may also influence (or be influenced by) outpatient drugs either as service complements or substitutes.[1]

A second aim of the study is to investigate utilization patterns with analytical tools that control for differences among Medicare beneficiaries as they approach

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[1]For example, inpatient hospital care may be a complement to outpatient drug use if the condition which necessitated the admission also requires continuing post-hospital drug treatment. On the other hand, hospitalization may be a substitute for outpatient prescriptions if the drugs needed in the for post-hospitalization phase are provided via an inpatient prescription.

death. Most previous work in this area is descriptive and therefore subject to confounding by factors that may be correlated with dying (age and widowhood for example). We employ both a cohort design in which utilization rates for a small group of beneficiaries ($n = 758$) are tracked for 36 months up to their deaths, and an interval cross-sectional (ICX) design in which a larger group of decedents ($n = 5{,}261$) is tracked for up to 72 months prior to death. The cohort design provides strong controls for interpersonal variation, whereas the ICX design controls for temporal differences. The combination of relatively long pre-death exposure periods with a unit of analysis as short as the month enables us to detect effects of impending death with a much higher degree of exactitude than in previous studies.

DATA

We selected for study a random 3-percent sample of Medicare beneficiaries who had enrolled in the PACE program at some time between July 1984 and November 1988. PACE is the largest pharmaceutical assistance program in the Nation. Since its inauguration in 1984, the program has provided prescription benefits to more than 850,000 elderly State residents. In June 1993, there were 341,361 enrollees (Pennsylvania Department of Aging, 1993). PACE eligibility is restricted to residents 65 years of age or over with incomes under $13,000 if single and $16,200 if married.[2] The program covers all Federal legend drugs, insulin, and insulin syringes dispensed either on an outpatient basis or to enrolled nursing home residents. Beneficiaries are responsible for a modest copayment per prescription or prescription refill ($4 up to July 1991, $6 thereafter).

We had access to the complete PACE enrollment and claims history (1984-88) for the sampled individuals. From the annual PACE application form, we obtained information on age, gender, race, prior-year income, marital status, and type of residence. PACE beneficiaries' Social Security numbers were used to link to Medicare Health Insurance Skeleton Write-Off files and then to the Medicare Part A Data Retrieval System (MADRS) and Part B Medicare Annual Data (BMAD) file claims records. We were able to link PACE and Medicare records for approximately 83 percent of the individuals selected, yielding a final sample size of 18,278. Annual MADRS and BMAD records from January 1984 to December 1988 were obtained for each of these beneficiaries. Finally, we screened death records maintained by the Health Care Financing Administration (HCFA) and the Pennsylvania Department of Health to determine date of death for any sample member who died between August 1984 and December 1989. This article focuses on the 5,261 decedents identified through the screening process.

Four broad-based utilization measures were selected to represent drug use and the other service categories: (1) counts of prescriptions and prescription refills recorded in PACE claims files,[3] (2) counts of ambulatory physician visits,[4] and (3) Part B charges, both from BMAD records, and finally (4) Part A charges from MADRS. Measures of counts and charges are based

[2] It is estimated that about one-half of the State's 1.7 million elderly met these PACE income guidelines in 1992.

[3] We also investigated prescription charges per month. The results for analyses based on prescription counts and pharmacy charges were virtually identical. For this reason, we only report results for prescription counts.

[4] Ambulatory physician services were defined to include the following Medicare procedure codes from the HCFA Common Procedure Coding System: office and other outpatient medical services (90000-90080), home visits (90100-90170), nursing home visits (90300-90370), boarding home visits (90400-90470), emergency room visits (90500-90580), outpatient psychiatric services (90830-90853), and outpatient ophthalmology services (92002-92012).

on all non-duplicate claims submitted to PACE or Medicare, irrespective of payment status. This preserved records of services that might not have been paid because of deductible or copayment exclusions and other administrative restrictions. To avoid confounding expenditure trends by inflation, we deflated Part A and Part B charges to 1990 dollar terms using the monthly Consumer Price Indexes for hospital and physician services, respectively. An unknown but presumably small number of services are missed in cases where providers (or patients in the case of non-assigned Medicare claims) fail to file a claim, through either neglect or advance knowledge that it will be rejected.

We wished to maximize the opportunity to observe short-term changes in the use of health services over time and consequently chose the person-month as the unit of analysis. Previous studies have shown that most of the increased utilization of health services in the final year of life takes place in the 2 to 3 months immediately prior to demise (Lubitz and Prihoda, 1984; Long et al., 1984). The choice of the person-month permits us to determine whether this is also true for outpatient prescription drug use.

The resulting panel data set contains more than 15 million records representing from 1 to 60 monthly observations per person for each study variable. The panel contains approximately 2.7 million records for individuals who died during the study period.

ANALYTIC FRAMEWORK

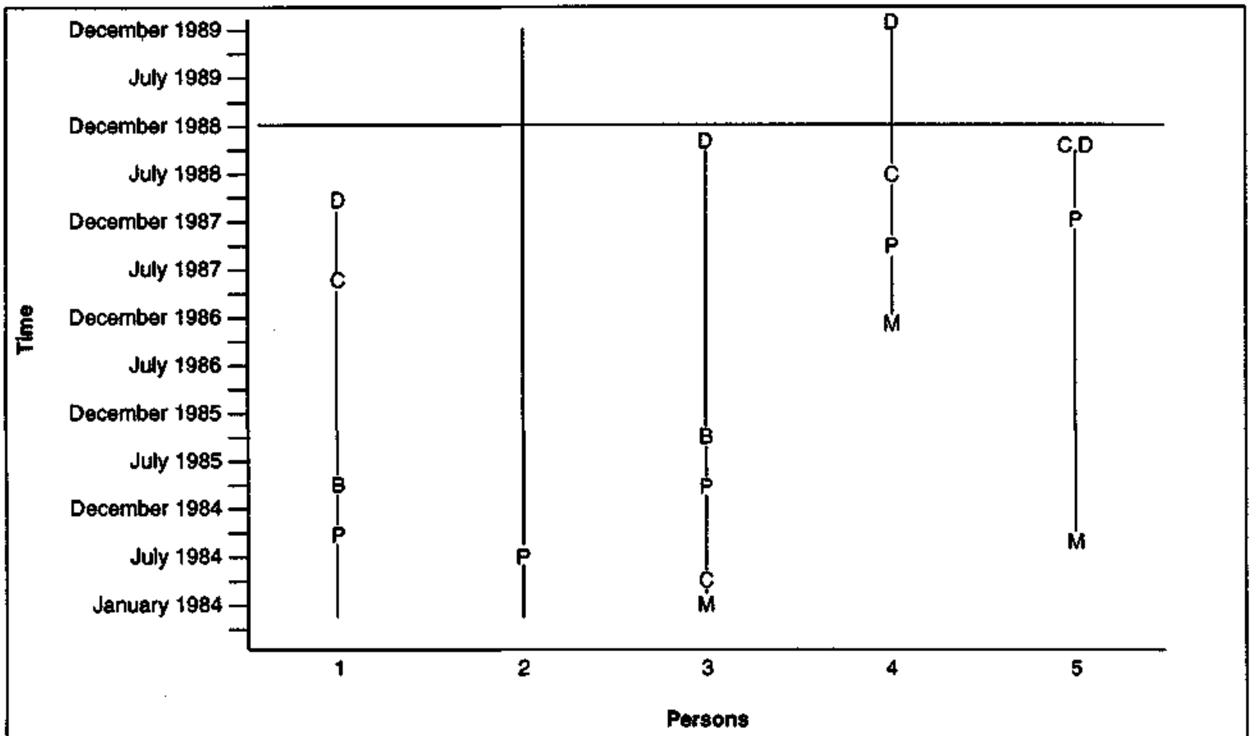
Traditional cohort analyses assume a rectangular panel—that is, a comparable (and complete) time series of observations for each person in the sample. Our sample does not share this attribute. Although each decedent in the panel is represented with at least one set of monthly observations prior to death, the number of observations declines as the death month becomes more remote. For example, the panel contains 12 (or more) months of observations prior to death for 81 percent of decedents, 24 months for 70 percent of decedents, and 36 months for 52 percent of decedents. Restricting the decedent sample to persons continuously enrolled in PACE up to their month of death—a necessary condition for time-series analysis of changes in prescription drug use—reduces these percentages considerably. In fact, just over 4 percent of the sample (758 decedents) had at least 36 months of continuous PACE enrollment prior to their deaths.[5]

Given the relatively small number of decedents in our sample, we sought analytic methods that would maximize the length of observation period while maintaining as much sample variability as possible. The solution was to conduct two parallel analyses: One was a cohort study based on 36 pre-death-month observations for the 758 sample decedents who were continuously enrolled in PACE during their final 3 years of life, and the other was a cross-sectional analysis of a single randomly drawn monthly observation for every decedent in the sample (n = 5,261). Each method has unique strengths. The cohort approach permits detailed study of individual-level changes in utilization behavior as death approaches for a small subset of decedents. The ICX design assures that some information from every eventual decedent is incorporated in the sample and permits analysis of time factors that cannot be readily modeled in traditional panel designs.

Figure 1 illustrates how each study sample was selected. The horizontal axis

[5]The number is small because only individuals who died after July 1987 could possibly have 3 years of PACE enrollment, given that the program began operation in July 1984.

Figure 1
Structure of the Data Set



NOTES: M is date of Medicare enrollment. P is date of enrollment in the Pennsylvania Pharmaceutical Assistance Contract for the Elderly. D is date of death. B is begin date for persons in decedent cohort. C is person-month selected for decedent interval cross-section.

SOURCE: Stuart, B., and Coulson, N.E., Pennsylvania State University, 1993.

represents persons included in the original sample of 18,278 Pennsylvania PACE beneficiaries. The vertical axis reflects the study timeframe. Vertical lines indicate the observation periods available for analysis for each individual sampled. These periods begin with the first month of data collection (January 1984 for Medicare data; July 1984 for prescription data) or enrollment in Medicare (M), whichever comes later. They end with death (D) or the end of the study period (December 1988). Because death records were tracked through December 1989, dates of death are recorded up through that month even though service data collection ended in December 1988. The purpose of tracking death dates beyond December 1988 was to expand the decedent sample and to enable us to determine the remaining lifetime (through 1989) for persons observed during the study period. For example, individual number 4 in Figure 1 died in December 1989. Had we observed that individual in July 1988, we would have known (in retrospect) that that person had 17 months to live.

The subsample of decedents selected for the cohort study was restricted to individuals who died between July 1987 and December 31, 1988, and were continuously enrolled in PACE and Medicare Parts A and B for the 36 months prior to their deaths. Although these conditions are necessary to assure that the time series for all four utilization variables are of equal length, they have important implications for the study design. First, note that only two of the four decedent time paths illustrated in Figure 1 meet these conditions (persons 1 and 3). The selection criteria systematically exclude two groups: "younger" elderly who reached 65 years of age (and Medicare entitlement) after the second year of observation (person 4) and individuals who joined PACE relatively close to their demise (person 5). If excluded and included persons differ in their use of health services prior to death, then results gained from the cohort analysis cannot be generalized to the host population. This is not a problem unique to our data set. Selection bias is a generic concern for cohort studies, particularly those with many repeat measures.

A second point worth noting is that our decedent panel is not rectangular in the conventional sense that each individual is observed over an identical timeframe. The fact that the individual time series is out of phase actually strengthens the design in that it reduces the possibility that utilization patterns will be incorrectly attributed to progression to death when they are actually because of changes in the external environment. Nonetheless, the phasing is close enough that deterministic time trends and other factors, like length of PACE enrollment, cannot be included in the multivariate analysis because of near-perfect collinearity to month-to-death measures.

The ICX design avoids these two major problems. It involves a two-stage sampling procedure in which persons are selected in the first stage and observation periods (months in our case) in the second. Random sampling in both stages assures that the final sample is representative of the original population in both cross-sectional and longitudinal dimensions (Stuart and Coulson, 1993).[6]

The principal advantage of the ICX approach is that we can model temporal variables without the necessity of repeat measures for each individual in the data set. Returning to Figure 1, we note that individual 1 enrolled in PACE in October 1984 and died in October 1988. Assume that the ICX-sampling procedure selected July 1987 as the one observation for that individual (the selected month is illustrated by C in Figure 1). With the information available from the original panel data set, we can construct values for three temporal measures: calendar date (July 1987, coded as "42," which represents the number of elapsed months since the beginning of the data collection period in January 1984), months of remaining life (17), and months of continuous PACE enrollment (33). For individual number 5, at observation month C, the values for these three measures are 57, 0, and 10, respectively. In this manner, the ICX design maintains temporal variation in the data set without the need to restrict the sample to persons with long and coterminous observation periods

The ICX sample generated for this study represents one randomly selected month for each of the 18,278 individuals in the original sample including the 5,261 decedents. The original sampling frame was constructed so that Medicare information would be available for periods predating PACE enrollment for most sample members and postdating it for a few. To avoid potential selection bias in the Medicare observations contained in the ICX sample, we chose not to restrict the sample generator to PACE-enrolled months. As a result, approximately three-fifths of the observations in the ICX samples represent PACE-enrolled months and two-fifths

[6]Technically, the ICX is representative of the population unweighted by longevity. From an aggregate life-year perspective, this sample selection procedure underrepresents individuals with long lifespans and overrepresents those with short lifespans. However, this has no effect on the properties of regression estimates based on such samples as long as age and month to death are included as conditioning variables.

Table 1

Individual Characteristics, by Sample Configuration

Characteristic	Full-Sample Cross-Section (n = 18,278)	Decedent Cross-Section (n = 5,261)	Decedent Cohort (n = 758)
Observation Dates			
Median Observation	December 1986	October 1985	September 1986
Range of Observations	July 1984–December 1988	August 1984–December 1988	August 1984–December 1988
Sociodemographic Characteristics			
Age in Years	76	79	80
Percent:			
Males	38	45	40
Race Other Than White	7	6	4
Single or Divorced	19	16	14
Widowed	45	50	58
Nursing Home Resident	4	9	6
Annual Income			
Percent in Class:			
Less Than $3,000	6	6	5
$3,000–$5,999	24	28	28
$6,000–$8,999	31	33	36
$9,000–$11,999	25	23	22
$12,000–$15,000	14	10	9
Mortality Characteristics			
Percent Deceased by December 1989	29	100	100
For Decedents: Month From Observation to Death	23	23	18
PACE Enrollment Characteristics			
Percent in Class:			
PACE Participants Enrolled in Observation Month	62	59	100
Original PACE Enrollment Date: July 1984–September 1984	42	73	76
Future PACE Dropout	15	19	0
Months of PACE Enrollment	17	16	20

NOTE: PACE is Pennsylvania Pharmaceutical Assistance Contract for the Elderly.

SOURCES: Data for 1984-89 are from the Pennsylvania Department of Aging, Bureau of Pharmaceutical Assistance; the Health Care Financing Administration; and the Pennsylvania Department of Health.

non-enrolled months. Naturally, analyses of prescription use are limited to the PACE-enrolled months (n = 3,091).

The characteristics of the study population are shown in Table 1. The first column presents characteristics for the entire study sample of 18,278 individuals including survivors as well as decedents. The second column presents means for the 5,261 decedents. The third summarizes characteristics of the subsample of 758 decedents selected for the cohort analysis. The "vintage" of the three sets of observations varies by nearly a year, with the decedent cross-section reflecting a slightly earlier period than either the full-sample cross-section or the decedent cohort. The sociodemographic characteristics reflect expected differences between population-based and mortality-based samples. The average age for PACE beneficiaries during the 5 years of the study is 76. Ages are 3 to 4 years higher in the decedent samples. Males, widows, and nursing home residents are represented in greater proportion in the decedent samples. Black persons, single persons, and persons with higher annual incomes are represented in smaller proportions compared with the host population. Twenty-nine percent of the full sample died by December 31, 1989. To capture time-to-death effects for these individuals, we created both a continuous variable and a set of dummy variables representing the difference in months between the observation and the death month. The ICX observations for decedents-to-be reflect periods that average 23 months prior to death.

Our previous work (Stuart et al., 1991; Stuart et al., 1992; Stuart and Coulson, 1993) has shown that PACE enrollees' utilization patterns are sensitive to enrollment date, how long they have been program beneficiaries (PACE exposure), and whether survivors maintain their PACE enrollment status (PACE dropout). We have included variables to capture each of these time dependencies. The variable "original PACE enrollment date: July 1984 through September 1984" is a dichotomous indicator of whether the individual joined PACE during the first 3 months of program operation. (Our prior work shows that these beneficiaries fill significantly fewer prescriptions compared with later program entrants). The average period of PACE exposure prior to the observation month was 17 and 16 months, respectively, in the full-sample and decedent cross-sections and 20 months for the decedent cohort. Fifteen percent of the full sample dropped out of PACE at some point prior to December 31, 1989.[7] The dropout rate is higher for the decedent cross-section (19 percent), perhaps reflecting transfers to Medicaid. Sample selection criteria for the decedent panel assure a zero dropout rate for this group.

[7]Between 2 and 5 percent of dropouts subsequently reapply for PACE benefits.

ESTIMATION PROCEDURES

We employed various regression procedures to ascertain the relationship between month to death and the four utilization variables. The simplest are unconditional models that include only dummy variables representing month to death. These produce descriptive results of utilization changes relative to some baseline (excluded) period together with significance tests (*t*-statistics) for each month in the time series. We also estimated a number of conditional equations in which use is modeled as a function of other individual and time-specific variables in addition to month to death. The purpose of these regressions is to determine if the descriptive findings capture the "true" relationship of impending death on health care utilization or confound it with other dynamic changes at the individual or aggregate level.

The cohort sample was analyzed using fixed-effects models. Individual-specific heterogeneity is known to be the leading cause of variation in health care utilization (Newhouse et al., 1989), but it is of little policy interest. We eliminated it by subtracting the individual-specific mean (computed across the 36 months) from each of the individual's 36 monthly observations for every variable. This implies (as is always the case in the fixed-effect framework) that some demographic and other attributes that are fixed over time are also eliminated as covariates. The conditional fixed-effects models included age and dummy variables for month to death, nursing home residence, marital status, and income. Near perfect multicollinearity precluded adding the observation month (our measure of time trend) and PACE exposure variables to these models.

We had considerably more flexibility in modeling time-to-death effects in the ICX

Table 2
Mean Utilization Rates, by Sample Configuration

Monthly Use	Full-Sample Cross-Section (n = 18,278)	Decedent Cross-Section (n = 5,261)	Decedent Cohort (n = 758)
Prescription Claims	2.0	2.6	2.7
Medicare Variables			
Ambulatory Visits	0.5	0.6	0.6
Part A Charges	$290	$478	$505
Part B Charges	$136	$195	$207

NOTES: Prescription claims averages are computed for PACE-enrolled months only. Part A and B charges are deflated to 1990 dollar terms. PACE is Pennsylvania Pharmaceutical Assistance Contract for the Elderly.

SOURCES: Data for 1984-89 are from the Pennsylvania Department of Aging, Bureau of Pharmaceutical Assistance; the Health Care Financing Administration; and the Pennsylvania Department of Health.

sample. As with the decedent cohort, we estimated both unconditional and conditional models on each of the four utilization variables. The conditional models included all of the variables entered in the fixed-effects equations plus the observation month, gender, race, the original PACE enrollee indicator, and PACE exposure and dropout variables. In some regressions, we specified month to death as a linear continuous variable, and in others, as a set of 36 binary dummies (the reference period being months 37 to 72 prior to death). In one set of runs we tested for seasonal effects with calendar dummies (January, February, etc.).

The structure of the ICX data set also permitted us to estimate two-part models that distinguish any use from the level of use by users (Duan et al., 1983). Findings from these latter models are described briefly later in this article but are not formally presented there.[8]

Except where noted, all regressions were estimated with ordinary least squares (OLS) with no data transformations. We conducted tests with alternative estimators and various transformations and found the regression estimates to be quite robust to differences in functional form. For the sake of economy and ease of interpretation, only the OLS findings are presented.

RESULTS

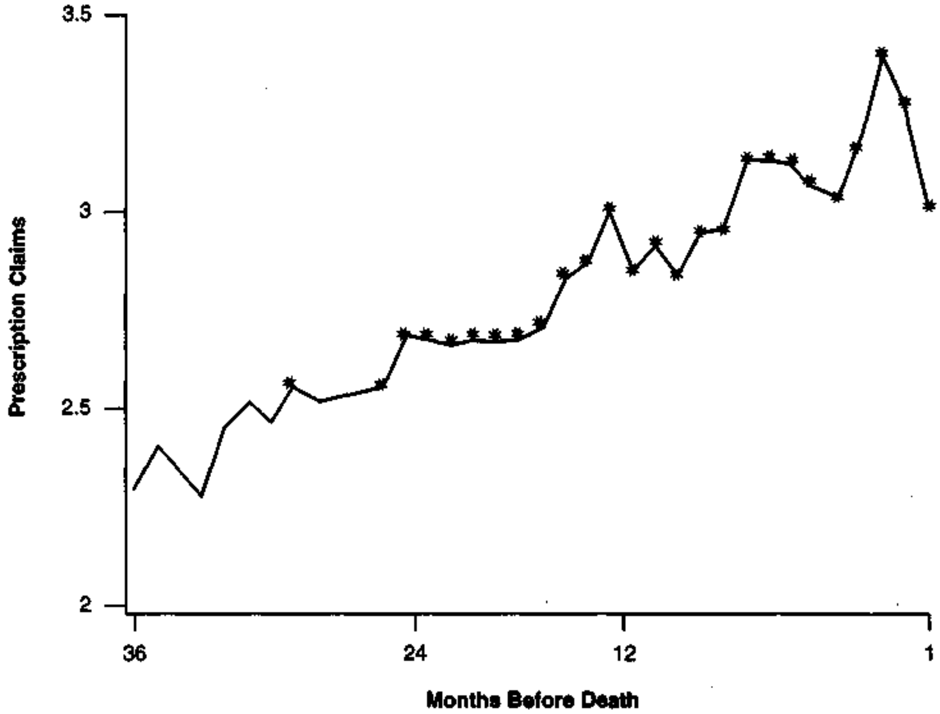
The utilization characteristics of the full sample (survivors and decedents) and the two decedent subsamples are shown in Table 2. The average individual in the host population (column 1) filled two prescription claims per month of PACE enrollment.[9] Prescription use was substantially higher in both decedent samples compared with the full sample: 30 percent greater in the decedent cross-section and 35 percent greater in the decedent cohort. Similar patterns hold for the Medicare variables. The average number of ambulatory physician contacts in the decedent samples was 20 percent above the host population average of one visit every 2 months. Differences in Medicare charges were more pronounced. Average monthly Part B charges in the decedent samples averaged 48 percent above those in the host population. Part A charges averaged 69 percent higher.

Figures 2 through 5 show changes in monthly utilization rates as death approaches for the 758 individuals in the decedent cohort. The 36th month prior to death is the reference point for each

[8]These regression results are available upon request from the authors.

[9]An analysis of prescription utilization based on 100 percent of the PACE claims files produced an identical average of 2 claims per month from July 1984 to December 1988 (Stuart et al., 1992).

Figure 2
Time Path of Outpatient Prescription Claims in the Months Preceding Death: Cohort Sample ($n = 758$)



* Significant difference from the 36th month at $p < .05$.
SOURCE: Regression analysis performed by Stuart, B., and Coulson, N.E., Pennsylvania State University, 1993.

utilization measure. The mean values for that month were added to the regression coefficients in the unconditional OLS regression models to produce the trend lines shown in these figures. An asterisk above a coefficient indicates that it is significantly different from reference month at $p < .05$.

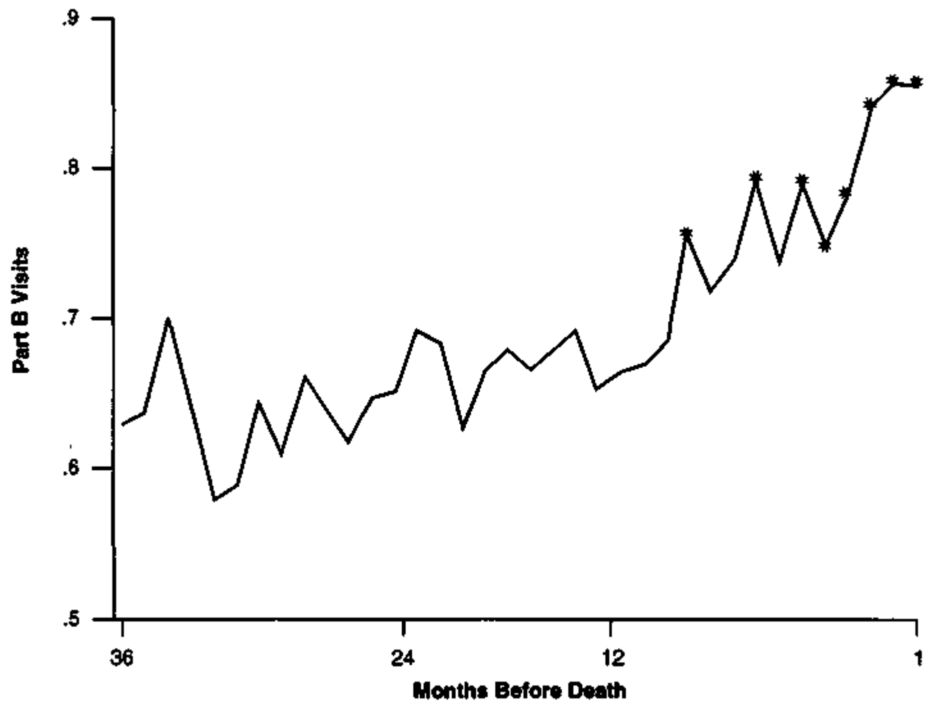
The time path for prescription drug use shows a consistent upward progression during the 3 years prior to death, with short reversals along the way that appear to be randomly distributed. The reversal in the final 2 months of life was anticipated, in part because the death month contains fewer days than prior months and in part because PACE does not pay for inpatient prescriptions. Both our own work and that of others (Long et al., 1984) have shown that the probability of hospitalization increases dramatically in the final 2 months of life.[10] Discounting these final 2 months, the members of the cohort sample increased their prescription volume by more than one prescription per month, a 55-percent increase over the 3-year period.

The time path for ambulatory physician visits is shown in Figure 3. In this case, the anomaly appears at the beginning of the time path, with a sharp rise and an even

[10] For the decedent cohort members, the probability that a month contains 1 or more hospital inpatient days fluctuates between .03 and .09 over the 36th through 3rd month prior to death. It is .12 in the second-to-last month and jumps to .19 in the final month of life. However, as discussed later, this dramatic increase does not appear to explain the drop in prescription use during the final 2 months of life.

Figure 3

Time Path of Ambulatory Physician Visits in the Months Preceding Death: Cohort Sample (n = 758)



* Significant difference from the 36th month at $p < .05$.

SOURCE: Regression analysis performed by Stuart, B., and Coulson, N.E., Pennsylvania State University, 1993.

sharper fall from the 36th to the 32nd month to death. Thereafter, the trend line follows a similar track to prescription drug use, rising erratically to a peak 1 month before death.[11] From the 32nd month to the death month, physician visits increase by 45 percent.

Figures 4 and 5 for Medicare Part A and Part B deflated charges show a flat, albeit somewhat erratic, expenditure profile in the third and second year before death. These coefficients are insignificantly different from zero, indicating no discernable increase in spending during this period. At about the 9th or 10th month, the coefficients show a statistically significant but slow upward trend that carries on until 6 months before death. In the fifth month, there are very steep increases in both Part A and Part B charges, a trend that continues to demise. During this short period, Part A charges increase from a monthly average of about $400 to $1,900 (measured in 1990 dollar values), while Part B charges increase from about $150 to slightly more than $550 per month (also measured in 1990 dollar values).

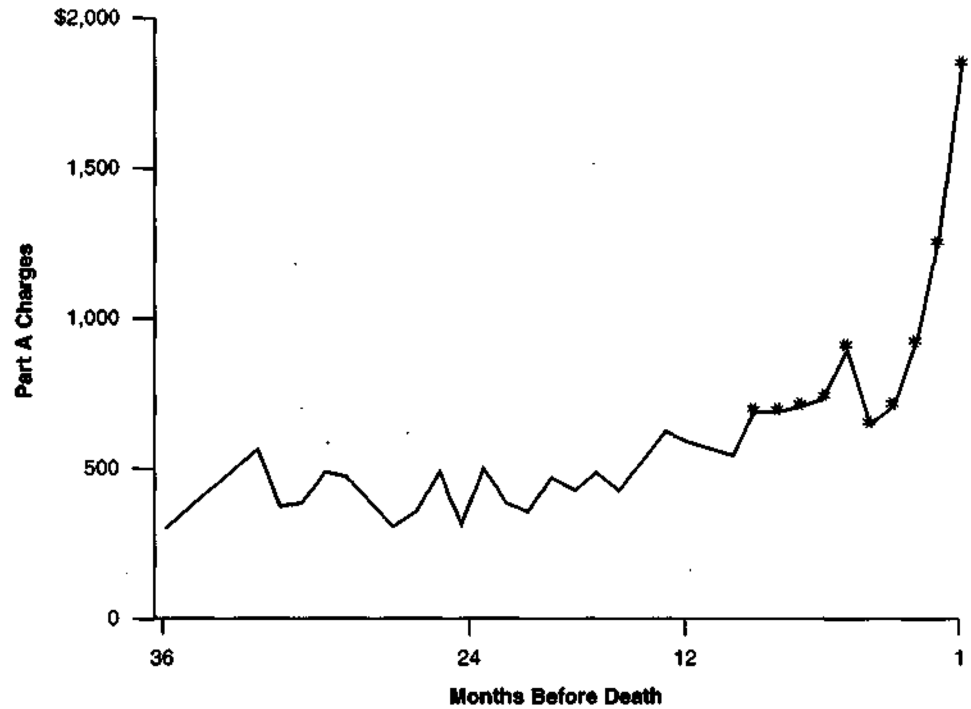
Adding covariates for age, residence, marital status, and income increases slightly the explanatory power of the cohort models, but it has virtually no impact on the month-to-death findings just reviewed.[12] The failure to find ameliorating effects may be a measurement artifact. Observations for these variables are derived from the

[11] The lack of statistically significant coefficients prior to the 11th month to death is an artifact of the choice of 36th month as the excluded reference period.

[12] The R^2s for the unconstrained cohort regressions clustered at .01. Covariates added less than a percentage point to the explained variance in every case.

Figure 4

Time Path of Medicare Part A Charges (Deflated) in the Months Preceding Death: Cohort Sample ($n = 758$)



* Significant difference from the 36th month at $p < .05$.

SOURCE: Regression analysis performed by Stuart, B., and Coulson, N.E., Pennsylvania State University, 1993.

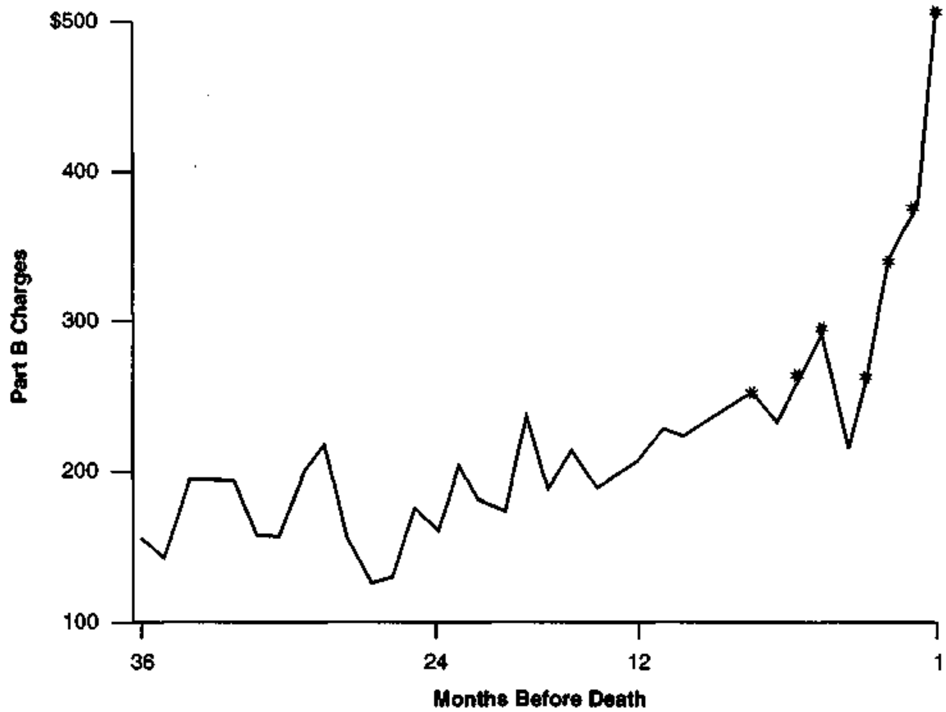
annual PACE application. With only 3 independent observations per individual during the 36 months prior to death, there may be insufficient variation to affect the monthly time series of utilization variables. The conditional-model approach did permit testing the hypothesis that declines in drug use in the final 2 months of life are due to inpatient hospitalization. Including a Medicare inpatient hospital variable in the drug-utilization equation showed—as expected—that persons hospitalized in a given month fill significantly fewer outpatient prescriptions during that month. However, the effect on the month-to-death time path—including the final 2 months—was imperceptible.

The decedent cross-section sample contains more sample variation among individuals (5,261 versus 758) and across time (up to 6 years versus 3 years) and for this reason is better suited for analysis of potential confounding effects in the utilization time paths of impending decedents. On the other hand, because each individual is represented by a single observation, individual-specific heterogeneity also leads to less precise month-to-death utilization coefficients compared with the cohort findings just described.

The unconditional utilization time paths computed from the decedent cross-section sample have the same basic characteristics as the cohort sample plots shown in Figures 2 to 5 (albeit with greater month-to-month variation) and hence are not presented. Instead, we go directly to multivariate results from a series of single and two-part regression models with month to death coded both as continuous and binary dummy variables.[13]

Figure 5

Time Path of Medicare Part B Charges (Deflated) in the Months Preceding Death: Cohort Sample (n = 758)



* Significant difference from the 36th month at $p < .05$.

SOURCE: Regression analysis performed by Stuart, B., and Coulson, N.E., Pennsylvania State University, 1993.

Table 3 presents findings from four linear regressions with month to death represented by a continuous variable ranging from 1 (the death month) to 72 (an individual with an ICX month of January 1984 who died in December 1989). Note that the prescription drug regression is estimated for 3,091 individuals representing PACE-enrolled observation months, whereas the other three regressions contain observations for the entire decedent sample.

The four regressions explain very little of the individual variation in monthly health service use (R^2s range from .01 to .04). Month to death is the only consistently significant variable across the four equations ($p < .01$ in every case). The month-to-death coefficients are low in relation to their respective means, no doubt because the fitted relationship is linear whereas the actual relationship is not. The observation-month variable captures effects associated with the passage of time from 1984 to 1988. The coefficients on this time variable are positive and highly significant in the Medicare Part A and B charge regressions. Given that the two charge series have been deflated, these coefficients can be interpreted as representing real increases in Medicare utilization rates during the study timeframe. PACE exposure is an important determinant of prescription use in this population, with an effect size twice that of month to death. Although health services use is not strongly

[13] Early test runs indicated no seasonality in the time series, and for this reason the calendar month variables were excluded from all further runs.

Table 3

Linear Regression Results on Monthly Utilization Rates for Selected Health Services in the 6 Years Preceding Death

Explanatory Variables	Monthly Utilization Measures: Prescription Claims (n = 3,091)	Ambulatory Physician Visits (n = 5,261)	Medicare Part A Charges (n = 5,261)	Medicare Part B Charges (n = 5,261)
Intercept	**4.62	0.50	**-1,572.88	-296.73
	(3.13)	(1.59)	(-3.03)	(-1.66)
Observation Month	-0.01	0.00	**9.36	**2.25
	(-1.16)	(0.40)	(4.31)	(3.02)
Age:				
70–74 Years	-0.12	0.03	-112.71	-31.30
	(-0.65)	(0.66)	(-1.43)	(-1.16)
75–79 Years	-0.30	-0.003	-94.85	-32.80
	(-1.65)	(-0.07)	(-1.24)	(-1.25)
80–84 Years	-0.35	-0.03	-5.82	-19.95
	(-1.92)	(-0.63)	(-0.07)	(-0.74)
84 Years or Over	**-0.81	*-0.09	-148.18	-47.12
	(-4.47)	(-1.95)	(-1.88)	(-1.74)
Male	**-0.30	0.00	-14.22	21.08
	(-2.69)	(0.00)	(-0.28)	(1.20)
Race Other Than White	-0.24	0.07	**314.97	54.20
	(-1.10)	(1.16)	(3.35)	(1.68)
Marital Status Single or Divorced	-0.19	0.03	60.44	-5.45
	(-1.14)	(0.62)	(0.80)	(-0.21)
Widowed	0.23	0.05	35.20	3.86
	(1.64)	(1.18)	(0.54)	(0.17)
Annual Income:				
$3,000–$5,999	-0.07	0.07	-22.00	-2.36
	(-0.32)	(1.23)	(-0.22)	(-0.07)
$6,000–$8,999	0.01	0.05	-44.44	-8.40
	(0.06)	(0.89)	(-0.45)	(-0.25)
$9,000–$11,999	0.39	0.09	-57.34	-19.10
	(1.68)	(1.47)	(-0.55)	(-0.53)
$12,000–$15,000	**0.78	0.10	-104.79	-24.29
	(2.78)	(1.34)	(-0.85)	(-0.57)
Nursing Home Resident	0.26	**0.15	-140.08	-17.93
	(1.26)	(3.10)	(-1.75)	(-0.65)
Months to Death	*-0.01	**-0.004	**-6.17	**-2.35
	(-2.07)	(-3.91)	(-3.82)	(-4.24)
Original PACE Enrollee	**-0.64	**-0.20	33.53	-12.34
	(-3.17)	(-3.75)	(0.39)	(-0.42)
PACE Exposure	**0.02	0.001	-0.46	0.81
	(2.95)	(0.70)	(-0.19)	(0.96)
PACE Dropout	**-0.66	0.007	-0.15	0.20
	(-4.68)	(-0.11)	(-0.43)	(1.72)
ADJ R^2	0.04	0.01	0.02	0.01
F	7.10	3.62	6.03	5.15

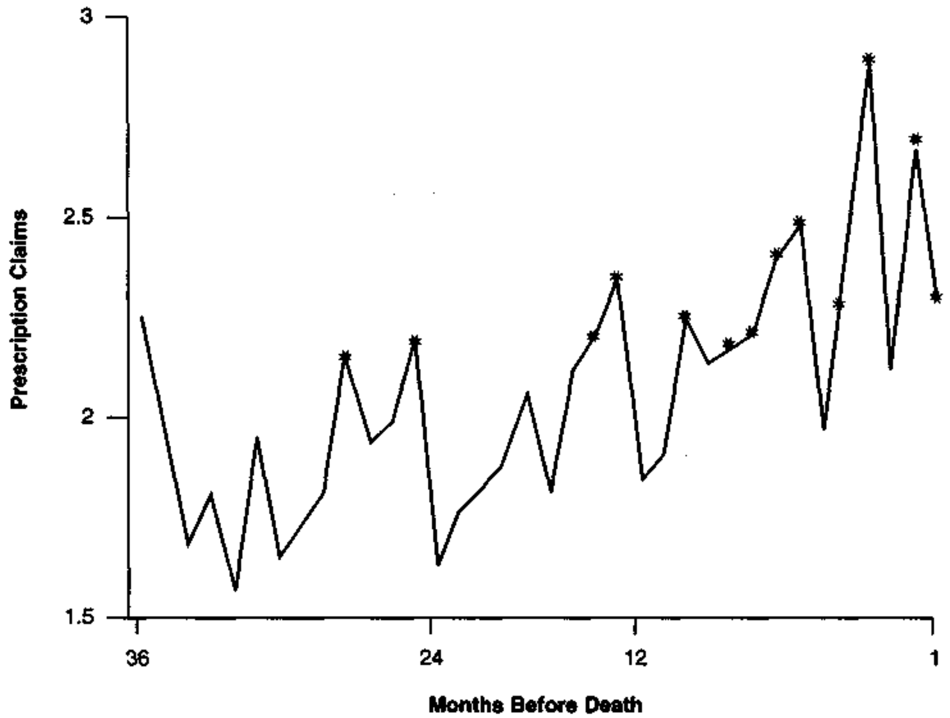
* Coefficient is significant at $p < .05$.
**Coefficient is significant at $p < .01$.

NOTES: Reference categories for the categorical variables are: age 65–69, female, white, married, community dwelling, annual income under $3,000, enrolled in PACE after September 1984, and did not subsequently drop out of the PACE program. t-statistics are in parentheses. PACE is Pennsylvania Pharmaceutical Assistance Contract for the Elderly.

SOURCES: Data for 1984-89 are for the Pennsylvania Department of Aging, Bureau of Pharmaceutical Assistance; the Health Care Financing Administration; and the Pennsylvania Department of Health.

Figure 6

Conditional Time Path of Outpatient Prescription Claims in the Months Preceding Death: Interval Cross-Sectional Sample ($n = 3{,}091$)



* Significant difference from the 36th month at $p < .05$.

SOURCE: Regression analysis performed by Stuart, B., and Coulson, N.E., Pennsylvania State University, 1993.

correlated with age, 15 of 16 age coefficients are negative relative to the excluded category of "young" elderly (65-69 years of age). This finding suggests compression of morbidity in the study population; that is to say, once impending death is controlled for, utilization rates for the surviving elderly tend to fall with advancing age.

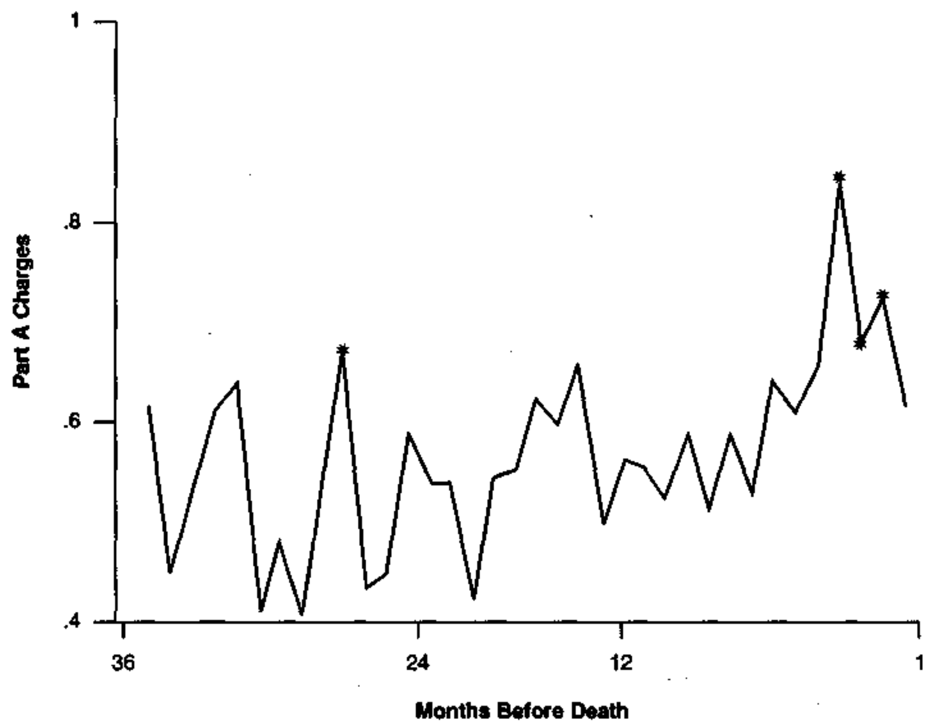
The observation month, exposure, and age variables also serve to purge the month-to-death coefficients of spurious time-related influences. The controls do make a difference. Figures 6 through 9 chart conditional month-to-death coefficients based on ICX regressions identical to those reported in Table 3 except that the continuous measure of time to death is replaced by 36 dummies and a reference period representing the 72nd through 37th month prior to death.[14] Selecting this reference period permits us to compare the utilization time paths from the cohort and ICX models.

Figure 6 shows the marginal effect of time to death on outpatient prescription drug use controlling for all other variables in the ICX model. The time path reflects erratic individual-specific heterogeneity, but this does not obscure the basic pattern of significant impending death effects beginning about 1½ years prior to demise. There is no evidence here that death effects are manifest before this point. Indeed, only 2 of 19 observations prior to the 17th month to death reached conventional significance

[14]The other variable coefficients in these regressions changed very little from the previous version and are thus not presented.

Figure 7

Conditional Time Path of Ambulatory Physician Visits in the Months Preceding Death: Interval Cross-Sectional Sample (n = 5,261)



* Significant difference from the 36th month at $p < .05$.

SOURCE: Regression analysis performed by Stuart, B., and Coulson, N.E., Pennsylvania State University, 1993.

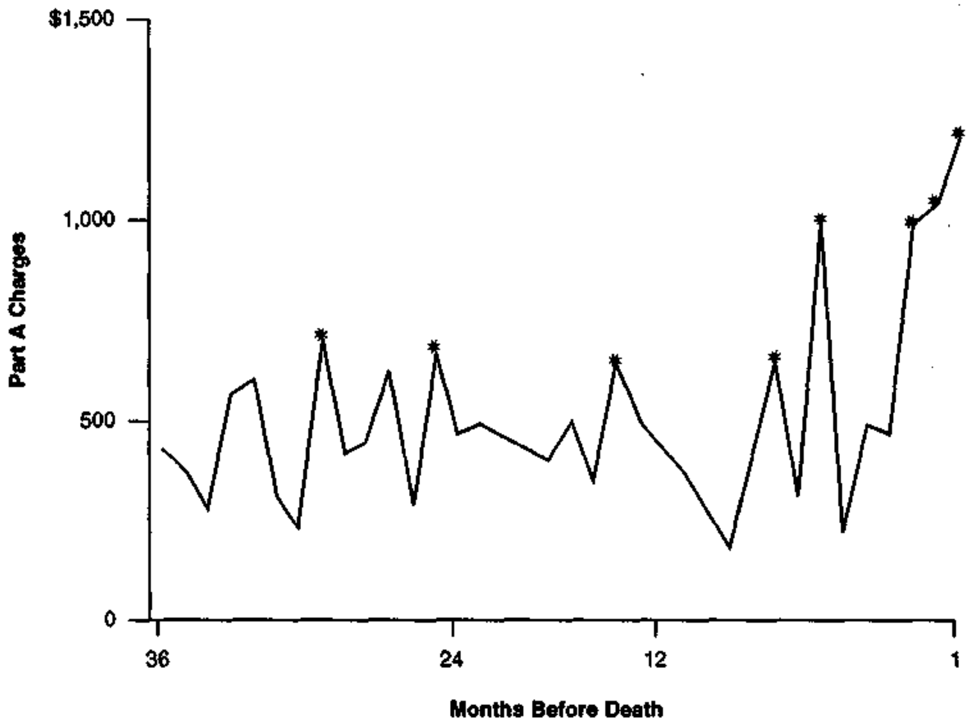
levels. This is a particularly noteworthy finding, given that the reference period extends from 4 to 6 years before death. We conclude from these findings that the upward trend in prescription use in the 36th through 18th month before death observed in the decedent cohort (Figure 2) is, in reality, an artifact of other time-related factors.

The ICX sample time plots for the Medicare variables tell a similar tale. Although there is a clear upward pattern in ambulatory physician visits (Figure 7) in the last year of life, only in 2 of the final 4 months are utilization rates significantly different from the base period of 4 to 6 years prior to death. As in the unconditional time plots, physician visits and prescription claims follow the same general pattern in the second-to-last and last year of life. The plots for Part A and Part B deflated charges (Figures 8 and 9) are also very similar to each other. They confirm the rapid escalation of use during the final 3 to 4 months of life and provide somewhat tenuous evidence of increases up to 6 months earlier.

With one notable exception, findings from the two-part, probability-of-use and level-of-use-by-users equations correspond closely with the single-equation results. Month to death was the only variable to achieve significance in every probability-of-use model estimated ($p < .01$ in each case). Month to death was also highly significant in all but one use-by-users

Figure 8

Conditional Time Path of Medicare Part A Charges (Deflated) in the Months Preceding Death: Interval Cross-Sectional Sample (n = 5,261)



* Significant difference from the 36th month at p < .05.

SOURCE: Regression analysis performed by Stuart, B., and Coulson, N.E., Pennsylvania State University, 1993.

equation.[15] Versions of these models with month-to-death dummies substituting for continuous measures produced plots similar to those in Figures 6 to 9. For the most part, it thus appears that impending death has similar effects on both the probability and level of use.

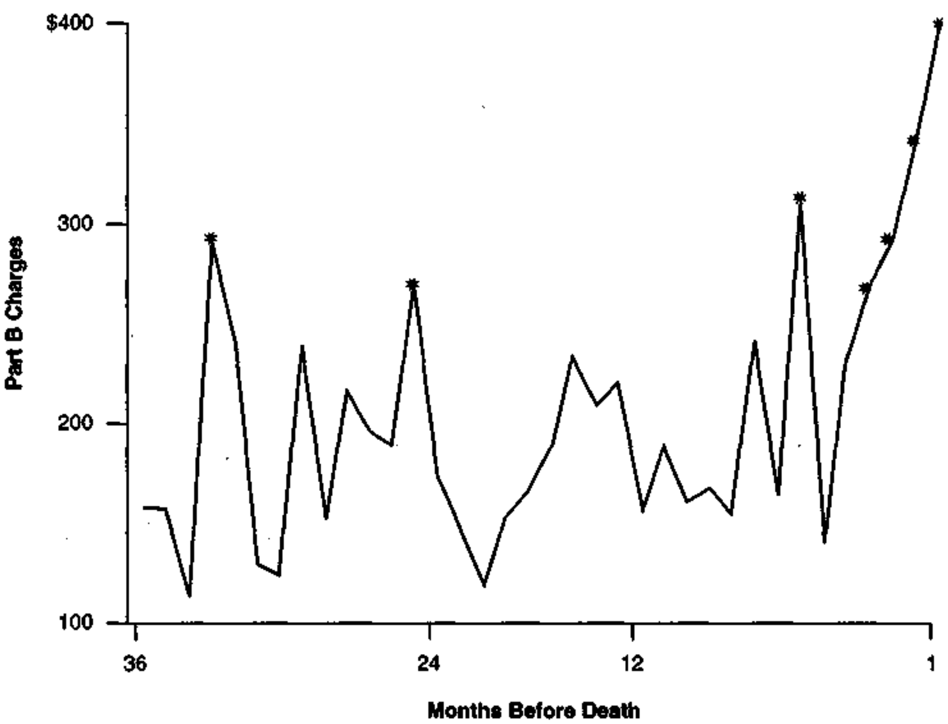
The exception is prescription drugs. We find that the probability of filling any prescription drug in a given month actually declines as death approaches. The negative relationship is highly significant (p < .001) and robust to alternative estimators. Logit and linear probability equations estimated for the same data produced identical results. In dummied-month versions of these models, the probability of use was actually below base-period rates in all but 5 of the final 36 months of life and all but 1 of the final 6 months of life (p < .05). We have no plausible explanation for this finding. It is not because of higher hospitalization rates in the months preceding death. We estimated a version of the probability model that included a Medicare hospital indicator, and the results were unchanged. Nor can it be explained by changes in prescription size or number of refills, because PACE rules limit days supply to 1 month per prescription or prescription refill.

Although the probability of drug use declines in the face of impending death, this is more than compensated for by increases in the level of use by users. We find that in just 7 of the final 36 months is

[15]Month to death is a marginally significant predictor of use of ambulatory physician visits (p = .12).

Figure 9

Conditional Time Path of Medicare Part B Charges (Deflated) in the Months Preceding Death: Interval Cross-Sectional Sample (n = 5,261)



* Significant difference from the 36th month at $p < .05$.

SOURCE: Regression analysis performed by Stuart, B., and Coulson, N.E., Pennsylvania State University, 1993.

use by users below base-period levels. Users' utilization rates are above base levels in each of the final 16 months of life. During the final 12 months of life, users' drug claims climbed to nearly 40 percent above base-period levels.

There is little resemblance between the probability-of-use time plots for prescription drugs and physician visits. Nor is there any direct correspondence between the level-of-use-by-users plots for the two utilization series. However, there is a striking similarity between the probability-of-physician-visit plots and the level-of-prescription-use plots over the final 12 months of life (but not before). This would suggest either that heavy users of prescription drugs are more likely to have outpatient physician contacts as death nears or that persons who visit physicians during this period of life receive relatively more prescriptions per visit than in earlier periods.

SUMMARY AND CONCLUSIONS

This study employed both cohort and cross-sectional techniques to assess patterns in the use of outpatient prescription drugs, ambulatory physician visits, and Medicare Part A and Part B charges in the final months of life for a sample of 5,261 beneficiaries of the Pennsylvania PACE program. Before summarizing our findings, it will be useful to reiterate some of the strengths and limitations to the study design.

The study is the first to examine outpatient prescription drug use and its relationship to Medicare services during this critical period of life. Nonetheless, the sample is restricted to elderly Medicare beneficiaries in a single State who received comprehensive prescription coverage from a means-tested pharmaceutical assistance program. There can be no assurance that the patterns observed here would be replicated in other States or in less comprehensive prescription drug programs.

In one set of analyses, we determined time-to-death effects by computing utilization rates for a cohort of individuals during a 36-consecutive-month period ending in death. In a second set of analyses, we inferred these effects from a single monthly observation drawn from a sample of individuals with known longevity. Both techniques yield more consistent results than the more common practice of computing utilization or reimbursement ratios in which the average utilization levels for individuals at a given point prior to death are divided by the average for the population as a whole during the same calendar period. Utilization and reimbursement ratios are highly sensitive to the demographic composition of the host population (i.e., populations with a high proportion of soon-to-be-decedents in the denominator will have systematically lower ratios at every period prior to death than will populations with younger age and lower mortality-risk profiles). This problem can be handled through stratification if the sample population is large and the analyst has knowledge of the major mortality risks affecting its members. We could satisfy neither condition. However, because we restricted the sample to individuals with known longevity, our results are not sensitive to mortality risk.

This study also differs from most previous work in the selection of the person-month as the unit of observation. Our rationale was to maximize the opportunity to observe short-term changes in use prior to death, and the rapid month-to-month increases in utilization rates immediately preceding death would appear to justify this decision. However, there is also considerably more random fluctuation in monthly utilization rates than in rates for longer periods, and this individual-level variation is reflected in the low R^2s obtained in the estimating equations.

Our findings regarding Medicare services generally conform to those of earlier studies. Part A and Part B service utilization rates rise slowly until the final 4 or 5 months of life, at which point they jump steeply. Ambulatory physician contacts follow the general pattern for Part B services except for the final month of life when use appears to level off or even decline somewhat. Multivariate analysis adds to our understanding of these time trends but does not alter the basic patterns just described. When Medicare utilization is decomposed into separate measures for probability-of-use and level-of-use-by-user, the effect of impending death is evident in rising utilization rates in both time series.

We find that prescription claims also mount as death approaches. However, simple descriptive trends tend to overstate the impact of impending death on drug use, at least for our sample of PACE beneficiaries. Multivariate regressions on a cross-sectional sample of observations from 1 to 72 months prior to death produced no evidence of an impending death effect on drug use before the second-to-last year of life, and only marginal evidence of an effect in the second-to-last year. Separating the utilization measure into probability-of-use and level-of-use components produced unexpected results. The probability of use actually declines significantly during the entire 3 years prior to death, with a particularly steep drop in the last 3 months of life.

Conversely, the level of prescriptions filled by users rises throughout this period. The two trends are self-canceling up to the second-to-last year of life. For the next 20 months or so, rising use by users predominates, leading to an increasing trend over all. In the last 2 to 3 months of life, declining probabilities of drug use drive the overall trend downward.

As expected, there is correspondence between outpatient physician contacts and prescription use as death nears, but the relationship is not a simple one. And even though we can rule out concurrent hospitalization as a cause for declining probability of outpatient drug use, our data base is too limited to investigate other possible reasons. Other researchers have found that time paths for Medicare expenditures prior to death vary substantially between persons suffering from acute and chronic conditions (Riley and Lubitz, 1989). We would expect to see corresponding differences in patterns of prescription use according to individual diagnosis and perhaps by type and class of drug used.

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