Uninsured Info

Comparison of Uninsured and Privately Insured Hospital Patients

Condition on Admission, Resource Use, and Outcome

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

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INCREASING evidence suggests that the amount and content of medical care that an individual receives in the United States is related to whether the individual has health insurance.[1] Previous studies, for example, although uncontrolled for differences between groups in clinical status and need for services on admission, have shown that the uninsured have fewer physician visits per year, are less likely to have a hospital stay, and, if hospitalized, have a shorter length of stay than insured individuals.[2-5] Other studies have shown that, even after controlling for differences in admission diagnosis, the services that uninsured individuals receive are different from those that insured individuals receive.[6-8] In a recent analysis of 1983 hospital discharge data for 65 032 uninsured, Blue Cross, and Medicare patients hospitalized in the Boston area, for example, Weissman and Epstein[9] found, after adjusting for case mix, that insured patients had 7% shorter stays and underwent 7% fewer procedures than Blue Cross patients. More recently, Wenneker et al[7] found, after controlling for clinical and demographic factors, that insurance status was strongly associated with utilization of three expensive cardiac procedures among 28 000 patients admitted to Massachusetts hospitals in 1985 with circulatory disorders or chest pain. Very few data are available regarding differences in health outcomes among insured vs uninsured patients.[4]

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

The research reported herein extends prior work in four important ways. First, we used a large national sample of 592 598 hospital discharges to compare the characteristics of hospital care received by uninsured and privately insured patients. Thus, our findings should be more generalizable to the nation's uninsured than those of studies based on a single, relatively small geographic area, such as Boston or Massachusetts. Second, the only previous study that has examined insured vs uninsured patients' use of high-cost procedures focused exclusively on three cardiac procedures.[8] In contrast, while controlling for diagnosis, we analyzed the frequency with which a broad spectrum of high-cost invasive and noninvasive procedures are provided to insured vs uninsured hospitalized patients. Third, we used multiple measures of condition on admission and resource use, which allow us to explore these factors in greater depth. Finally, we examined differences between the uninsured and privately insured in in-hospital death rates, an important measure of outcome.

DATA, VARIABLES, AND STATISTICAL METHODS

Data

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

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

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

Variables

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

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

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

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

Average Length of Stay.—Length of stay is a common measure of resource use. However, even when controlling for diagnosis, length of stay can reflect the severity of a patient's condition as well as the level of resource use for a patient with a given level of severity. We attempted to control for the effects of severity by focusing the analysis of length of stay on diagnoses that have either a high or a low degree of physician discretion with regard to the need for a hospitalization. Using both prior literature[14-17] and clinical judgment, we initially selected 28 diagnoses thought to differ in the amount of discretion, from the physician's perspective, in the need for a hospital admission. Relatively high-discretion diagnoses may be thought of as being more elective, while relatively low-discretion diagnoses may be considered nonelective.

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

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

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

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

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

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

Condition on Admission
Risk-Adjusted Mortality Index
Probability of a weekend admission

Resource Use
Average length of stay for each of 10 specific diagnoses
Incidence of selected high-cost or high-discretion procedures
- Coronary artery bypass graft surgery
- Total knee replacement
- Total hip replacement
- Stapedectomy
- Surgical correction of strabismus

Incidence of selected procedures, controlling for diagnosis
- Colonoscopy or proctosigmoidoscopy
- Upper endoscopy procedures
- Coronary arteriography
- Contrast myelogram or CT* of spine
- Head CT
- Bronchoscopy with biopsy
- Cholecystectomy with intraoperative cholangiogram

Tissue pathology of "not abnormal" for biopsy specimens obtained during selected endoscopic procedures

Outcome
In-hospital death rate

*CT indicates computed tomography.

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

		Relative Regression-Adjusted Coefficient	
Sex, Age, and Race	Unweighted N	Risk-Adjusted Mortality Index	Probability of a Weekend Admission
Males			
Age 1-17 y			
White	36 527	1.57	1.13†
Black	8813	1.56‡	1.00
Age 18-34 y			
White	74 514	1.82†	1.22†
Black	17 614	2.09†	1.24†
Age 35-49 y			
White	46 429	1.80†	1.17†
Black	11 384	1.59†	1.27†
Age 50-64 y			
White	54 417	1.45†	1.14†
Black	9175	1.44†	1.23†
Females			
Age 1-17 y			
White	29 724	2.00†	1.12†
Black	7098	1.81	1.12‡
Age 18-34 y			
White	72 780	1.80†	1.21†
Black	15 828	1.04†	1.24†
Age 35-49 y			
White	52 371	1.62†	1.16†
Black	11 338	2.24	1.15†
Age 50-64 y			
White	61 611	1.52†	1.11†
Black	10 395	1.88†	1.07

*Null hypothesis is that relative rate or probability equals 1.00.
†$P \leq .01$.
‡$.01 < P \leq .05$.

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

Statistical Methods

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

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

$$Y = \alpha U + \beta PI + \Sigma \mu_i X_i$$

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

The exact list of X_i variables depends on the particular dependent variable being analyzed and the extent of file stratification. In all analyses, the individual discharge was the unit of observation.

The coefficients α and β can be interpreted as mean values of Y for uninsured and privately insured patients, respectively, adjusting for differences in Y due to the X_i variables. The primary null hypothesis throughout the analysis is that there is no difference in Y between uninsured and privately insured patients, controlling for the effects of other factors, which is equivalent to the null hypothesis $\alpha = \beta$. This hypothesis can be evaluated using a t statistic calculated from the variance-covariance matrix of the regression parameters.[18] Note that because the regression model is linear in the parameters, the test of the null hypothesis is independent of the particular values of X_i chosen to standardize the values of α and β.

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

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

Controls for Factors Other Than Insurance Coverage

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

The regression model for average

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

Diagnosis, by Degree of Physician Discretion and Estimated RAMI (*ICD-9* Code)	Unweighted N	Estimated RAMI Value†	Relative Regression-Adjusted Average Length of Stay
High Physician Discretion			
Chronic tonsillitis (474.0)	3815	.001	0.88‡
Noninfectious gastroenteritis, NOS (558.9)	8460	.001	0.82‡
Acute bronchitis (466.0)	2684	.003	0.85§
Unilateral inguinal hernia (550.91)	4144	.004	0.62‡
Uterine leiomyoma, NOS (218.9)	3147	.005	0.77§
Low Physician Discretion			
Gastrointestinal hemorrhage, NOS (578.9)	1600	.028	0.89
Malignant neoplasm, bronchials/lung (162.9)	1548	.036	0.97
Congestive heart failure (428.0, 428.1, 428.9)	2768	.051	0.94
Acute myocardial infarction, inferior wall, NEC (410.4)	2512	.068	0.89‡
Acute myocardial infarction, anterior wall (410.0)	1788	.115	0.95

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

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

‡$P \leq .01$.

§$.01 < P \leq .05$.

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

RESULTS

Condition on Admission

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

Resource Use

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

In Table 4 we present further results pertinent to the question of differences in resource use between the uninsured and insured. Using the entire sample of discharges, we estimated regression models of the probability of undergoing each of five high-cost or high-discretion procedures among the privately insured and the uninsured. The results suggest highly significant and quantitatively large differences between uninsured and privately insured patients in the probability of undergoing these procedures. The uninsured are 29% less likely to have coronary artery bypass graft surgery and almost 75% less likely to have a total knee replacement. Rates for the other three procedures range from between 45% and 56% lower for the uninsured.

In Table 4, we also examine the relative use of other specific procedures, controlling for diagnosis by limiting each analysis to discharges with a specific diagnosis, group of diagnoses, or associated procedure. These analyses thus provide an opportunity to examine whether the clinical management of specific conditions varies with insurance status. In three of the 11 diagnosis-procedure combinations—colonoscopy or proctosigmoidoscopy, upper endoscopy, and coronary arteriography—there is a statistically significant difference ($P<.01$) in the probability of an uninsured person's undergoing the procedure relative to a privately insured person. In these instances, uninsured patients were from 20% to 44% less likely to have the procedure performed. We found no statistically significant difference ($P>.05$) between the uninsured and the privately insured in the use of myelography or computed tomography of the spine among patients with lumbar disk displacement, in the use of computed tomography of the head among patients with any of five diagnoses, in the incidence of biopsy among patients undergoing a bronchoscopy, or in the incidence of an intraoperative cholangiogram among patients undergoing a cholecystectomy.

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

Outcome

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

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

Procedure (Procedure Code)	Unweighted N	Relative Regression-Adjusted Coefficient
High cost or high discretion	592 598	
Coronary artery bypass graft surgery (36.10-36.19)		0.71†
Total knee replacement (81.41)		0.26†
Total hip replacement (81.5, 81.51, 81.59)		0.55†
Stapedectomy (19.1-19.2)		0.50†
Surgical correction of strabismus (15.1-15.9)		0.44†
Colonoscopy (45.23-45.26) or proctosigmoidoscopy‡ (48.23)	35 977	0.56†
Upper endoscopy procedure§ (42.23, 44.13, 45.13)	47 732	0.60†
Coronary arteriography‖ (88.55-88.57)	30 622	0.61†
Contrast myelogram or CT of spine¶ (87.21, 88.38)	5243	0.95
Head CT (87.03) by diagnosis		
Convulsions (DX code 780.3)	11 582	0.91#
Syncope and collapse (DX code 780.2)	4363	0.90
Concussion, NOS (DX code 850.9)	3290	0.88
Concussion without loss of consciousness (DX code 850.0)	2268	0.98
Concussion with brief loss of consciousness (DX code 850.1)	2684	1.02
Bronchoscopy (33.22, 33.23) procedures with a biopsy (33.24, 33.25)	3713	0.98
Cholecystectomies (51.22) with an intraoperative cholangiogram (87.53)	9500	0.94

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

†$P \leq .01$.

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

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

‖All discharges that had any of the following as the primary discharge diagnosis were considered: precordial pain (DX7865.1), chest pain (DX7865.0), angina pectoris (DX413.9), intermediate coronary syndrome (DX411.1), coronary atherosclerosis (DX414.0).

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

#$.05 < P \leq .10$.

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

Procedure (Procedure Code)	Unweighted N	Relative Regression-Adjusted Coefficient for Probability of 'Not Abnormal' Tissue Pathology,† Uninsured to Privately Insured
Esophagoscopy (42.23)	908	0.26‡
Gastroscopy (44.13)	2478	0.38‡
Proctosigmoidoscopy (48.23)	1120	0.87
Colonoscopy (45.23)	3781	0.52
Cystoscopy (57.32)	3320	0.34
Endoscopy of small intestine (45.13)	5408	0.51§
Bronchoscopy (33.22, 33.23)	2919	0.66
All above endoscopic procedures‖	18 934	0.64¶

*Null hypothesis is that relative coefficients equal 1.00.

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

‡$.01 < P \leq .05$.

§$.05 < P \leq .10$.

‖Includes a separate control variable for each endoscopic procedure.

¶$P \leq .01$.

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

Age, Sex, Race	Unweighted N	Relative Regression-Adjusted Coefficient
Males		
Age 1-17 y		
White	38 527	1.45
Black	8813	1.88
Age 18-34 y		
White	75 514	1.34†
Black	17 614	1.92†
Age 35-49 y		
White	46 429	1.45†
Black	11 384	3.20†
Age 50-64 y		
White	64 417	1.23†
Black	9175	1.08‡
Females		
Age 1-17 y		
White	29 724	2.37‡
Black	7098	2.56†
Age 18-34 y		
White	72 780	1.48§
Black	15 828	1.80
Age 35-49 y		
White	52 371	0.92
Black	11 338	1.60‡
Age 50-64 y		
White	61 611	1.15§
Black	10 395	1.35‡

*Null hypothesis is that relative coefficient equals 1.00.

†$P \leq .01$

‡$.01 < P \leq .05$.

§$.05 < P \leq .10$.

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

COMMENT

Our analysis strongly suggests that an individual's condition on admission, use of resources during hospitalization, and likelihood of in-hospital death vary depending on whether the individual has health insurance. In the case of condition on admission, compared with the privately insured, the uninsured were more likely to be admitted for a condition that has a higher expected risk of death (based on the RAMI score) and appeared to be in more urgent need of care. The latter conclusion is suggested by our finding that, when differences in case mix between the uninsured and privately insured are controlled for with the Medicare case-mix index and the RAMI, the uninsured were significantly more likely to be admitted on a weekend.

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

We also found evidence of a difference

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

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

Finally, our analysis demonstrates that the uninsured have a higher relative risk of in-hospital death than the privately insured in 15 of 16 age-sex-race-specific cohorts, even when the Medicare case-mix index, the expected risk of death, and whether a procedure was performed are controlled for. In 10 of these 15 cohorts, the difference is statistically significant ($P \leq .05$). Although it is possible that this observed difference in in-hospital mortality is due to underprovision of needed medical services to hospitalized uninsured patients, the difference also could be due to differences in severity of illness between the uninsured and privately insured that are not reflected fully in the Medicare case-mix index and the RAMI. It is also possible that privately insured patients are more likely than uninsured patients to be discharged to another facility, such as a nursing home or a hospice, where death might occur shortly after discharge from the hospital.

Our results extend the findings of other studies in several ways. First, we have provided evidence that insurance coverage has an effect on resource use for a broad spectrum of clinical problems in addition to lung cancer,[6] cardiovascular disease,[8] and the acquired immunodeficiency syndrome.[11] Second, we have found evidence for differences between the uninsured and the privately insured in resource use for some, but not all, conditions, and of differences in use of some, but not all, types of resources. Although our findings of differences in average length of stay for high- but not for low-discretion hospitalizations and of a consistently higher probability of completely normal tissue pathology in biopsy specimens obtained from the privately insured vs the uninsured do not clearly indicate whether differences in the care provided to the uninsured vs the privately insured reflect "the unmet needs of some of our citizens or the overmet needs of others,"[9] they do provide empirical support of the hypothesis that insurance coverage may have a bigger impact on the use of elective or more highly discretionary services than on medically necessary care.[19]

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

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

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Health Insurance and Mortality

Evidence From a National Cohort

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

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

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

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

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

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

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THE NUMBER of uninsured persons has increased throughout the past decade,[1] reaching 35.4 million in 1991.[2] There are several reasons for this,[3] including a changing economy and increased costs of health care and insurance, making both less affordable both for employers and individuals. As a result, there has been widespread recent interest in ensuring health insurance for all Americans. There is limited systematic evidence that a lack of insurance results in adverse health outcomes, particularly mortality.[4]

It is well established that a lack of health insurance is associated with reduced access to medical care; persons without insurance report fewer physician visits, after adjusting for health status,[6-8] and report delaying or forgoing medical care for serious symptoms. Medicaid coverage has been shown to reduce this differential.[5,6,12,13] A lack of health insurance is also associated with a lower prevalence of recommended preventive services.[14,15] Other studies have found that uninsured persons compared with the insured are more likely to have potentially avoidable hospitalizations,[16,17] may be sicker at the time of hospital admission,[16,18,19] are more likely to experience in-hospital[20-22] and cancer mortality,[23] and are less likely to receive invasive procedures.[18,24] Many of these studies suggesting adverse outcomes do not include uninsured persons who have not entered the health care system, so that the results may simply represent adverse selection due to delayed access to care.

Some studies do point to health hazards associated with lacking health insurance. Braveman et al[25] studied hospital discharge data in California and found that adverse outcomes in newborns were more frequent in the uninsured. The use of hospital discharge data limited the ability of the authors to adjust for potential confounders. Some studies have shown that persons losing health insurance benefits suffer measurable declines in their health.[26-28] In the Rand Health Insurance Experiment, on average, persons randomized to both the health insurance copayment group and the free care group ended the study with similar levels of health.[29] For persons with poor vision and poor persons with high blood pressure, however, free care brought an improvement.[29] Though all persons enrolled in the Rand study had health insurance, these results suggest that financial barriers in the form of copayments resulted in clinically important adverse outcomes. Hubbell et al found lower levels of functional sta-

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tus among 94 poor patients with financial barriers to care compared with 94 control patients without financial barriers. Patrick et al[31] found that poor persons in families with all members uninsured compared with those from insured families had lower self-rated health and greater perceived need of services. Most studies have focused on the poor with the implication that the benefits from health insurance are limited to poor persons. Lack of insurance is correlated with low income, which is itself associated with poor health,[32-35] and most previous studies do not permit isolation of the effects of being poor from that of being uninsured. Many of the studies are cross-sectional in design, limiting confidence in the direction of the causal pathway. Finally, none of these studies have addressed the relationship between insurance and mortality.

Because of the limitations of studies addressing this important policy issue we examined the relationship between health insurance and mortality, using data from the National Health and Nutrition Examination Survey Epidemiologic Follow-up Study (NHEFS), which followed a representative cohort of the US population for up to 16 years. The rationale and analysis used in the present study was informed by the conceptual framework developed in a US Congress Office of Technology Assessment report.[4] The availability of health insurance is viewed as an enabling factor facilitating access to care. As noted above, persons lacking insurance have reduced access to care, evidenced by reduced use of a wide range of health services and self-reported forgone and delayed care.[1,6,7,9-11,14,15] In addition, the process of care for uninsured persons may be inferior. For example, the uninsured have an increased risk of suffering medical injury due to substandard care.[36] Consequent on reduced access and lower quality of care, we hypothesized that persons lacking insurance would experience increased mortality. Our analysis adjusted for the potentially confounding effect of other factors identified in the Office of Technology Assessment report as affecting health. These factors are categorized as predisposing (age, gender, education, employment, and race), need (perceived health and objective evidence of morbidity), other enabling (income), and individual behaviors (smoking status, alcohol consumption, leisure exercise, and obesity status).

METHODS

The first National Health and Nutrition Examination Survey (NHANES I), conducted between 1971 and 1975, collected sociodemographic, health insurance and utilization, medical history, and clinical and laboratory information from several national probability samples of the civilian noninstitutionalized population.[37,38] Not all information was collected on all respondents. Detailed information including insurance status was collected on 6913 adults aged 25 to 74 years. Interviewees were also examined by physicians, who assigned up to 15 diagnoses based on history, physical examination, and the results of laboratory investigations. The NHEFS was designed to trace and reinterview respondents aged 25 to 74 years.[39] Follow-up surveys in NHEFS were conducted in 1982 through 1984, 1986, and 1987. Data collected comprised interview surveys, medical records from health care facilities, and death certificates for all decedents. The age, race, and sex-specific mortality of the NHEFS cohort is similar to that experienced by the US population.[40]

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

Analyses

The NHANES I used multistage stratified probability samples of clusters of persons. In addition, persons living in poverty areas, women of childbearing age, and elderly persons were oversampled. To accommodate the complex survey design the statistical package SUDAAN[43] was used in the analyses reported below. The SUDAAN program uses a Taylor series approximation method to compute variances that allow adjustment for the multistage probability sampling strategy. The revised weights provided on the 1987 NHEFS public use tapes were used to adjust for survey oversampling and nonresponse to yield population estimates of reported baseline descriptors. The univariate relationships between each baseline variable and health insurance status (dichotomized as reporting having no health insurance or having private health insurance) and mortality were examined using χ^2 tests. To examine the relationship between insurance status and subsequent mortality, proportional hazards survival analyses were used to adjust for other baseline variables, including the following: age (years); gender; race (dichotomized as white or black); education (dichotomized as at least 12 years of school or less); family income at baseline (treated as three dummy variables, income less than \$7000 per year, \$7000 to \$9999 per year, and \$10 000 to \$14 999 per year, using income of \$15 000 and higher per year as a reference); employment status (dichotomized as working most of the previous 3 months or not); morbidity (dichotomized as the presence or absence of evidence of morbidity on medical examination and laboratory testing); self-rated health (treated as three dummy variables, reporting health in general to be very good, good, or fair/poor with excellent as the reference value); smoking status (smoker or not); obesity status (dichotomized as body mass index [weight in kilograms divided by height in meters squared] >27 or not); leisure exercise (dichotomized as reporting little or no exercise compared with moderate or much); and alcohol consumption (dichotomized as consuming at least six alcoholic drinks per week or less). Because the family income variables were not adjusted for household size, the survival analyses also included household size as an interval level variable. In addition to a survival analysis including only main effects, a model with the interaction terms between insurance status and the other independent variables was examined. Interaction terms were retained if they were statistically significant ($P<.05$). To avoid the inefficiency of performing weighted multivariate analyses, we followed the recommendations of Korn and Graubard[44] to use unweighted survival analyses that

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

Characteristic	No. (%)	% Uninsured* (SE)	% Died† (SE)
Vital status			
Died	593 (10.7)	22.1 (2.4)	...
Alive	4289 (89.3)	11.7 (0.7)	...
Insurance status			
Uninsured	699 (12.8)	...	18.4 (2.0)
Insured	4183 (87.2)	...	9.6 (0.5)
Age group, y			
≥55	1206 (28.8)	17.4 (3.9)	27.1 (1.6)
45-54	1354 (24.9)	11.2 (2.0)	10.9 (0.8)
35-44	1035 (24.6)	10.1 (2.4)	5.7 (0.8)
25-34	1287 (22.8)	13.0 (3.7)	2.0 (0.5)
Sex			
Male	2136 (46.6)	13.8 (1.0)	13.7 (0.8)
Female	2746 (53.4)	11.7 (0.9)	6.1 (0.5)
Race			
Black	517 (9.1)	23.4 (2.5)	14.7 (1.9)
White	4365 (90.9)	11.8 (0.7)	10.3 (0.5)
Education, y			
<12	1684 (31.5)	22.6 (1.4)	18.7 (1.2)
≥12	3106 (68.5)	8.4 (0.6)	7.0 (0.4)
Income, $			
<7000	950 (18.0)	34.7 (2.2)	19.2 (1.6)
7000-9999	1074 (25.1)	12.9 (1.1)	11.7 (1.0)
10 000-14 999	1283 (27.6)	6.4 (0.7)	7.6 (0.9)
≥15 000	1391 (29.3)	4.6 (0.9)	6.7 (0.7)
Employment status			
Unemployed	1796 (33.6)	20.3 (1.2)	12.0 (1.0)
Employed	3078 (66.4)	9.1 (0.7)	9.6 (0.5)
Self-rated health			
Fair or poor	825 (16.9)	27.6 (1.7)	24.7 (1.6)
Good	1532 (31.4)	13.2 (1.2)	12.4 (0.9)
Very good	1292 (26.5)	12.6 (1.0)	8.7 (0.9)
Excellent	1229 (25.2)	8.4 (0.9)	5.9 (0.8)
Morbidity			
Present	2822 (56.6)	15.7 (1.0)	14.9 (0.9)
Absent	1960 (43.4)	9.1 (1.0)	5.3 (0.6)
Leisure exercise			
Little or none	1885 (39.7)	16.0 (1.1)	12.6 (0.9)
More	2993 (60.3)	11.3 (0.9)	9.4 (0.6)
Smoking status			
Smoker	1968 (42.0)	14.3 (1.0)	13.4 (0.9)
Nonsmoker	2914 (58.0)	11.8 (0.9)	8.8 (0.5)
Alcohol consumption			
≥6 drinks per wk	1269 (27.0)	11.9 (1.3)	13.4 (1.0)
<6	3613 (73.0)	13.2 (0.8)	9.7 (0.6)
Obesity status‡			
BMI, >27	467 (9.3)	16.9 (1.8)	14.4 (1.7)
BMI, ≤27	4415 (90.7)	12.5 (0.7)	10.3 (0.5)

*Percentage of those with the characteristic who were uninsured at baseline.
†Percentage of those with the characteristic who died during the follow-up period. Percentages are weighted to provide population estimates, and SEs are adjusted for cluster sampling strategy.
‡BMI indicates body mass index (weight in kilograms divided by height in meters squared).

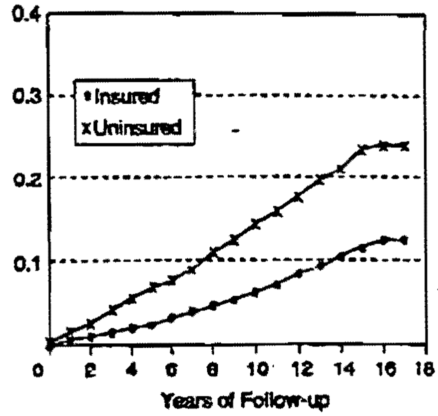

Fig 1.—Cumulative mortality probability by baseline insurance category.

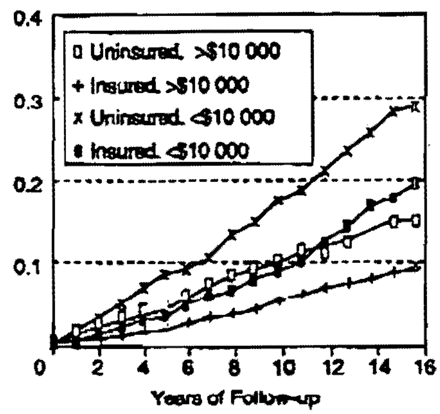

Fig 2.—Cumulative mortality probability by baseline insurance and income categories.

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

RESULTS

Baseline characteristics of the sample are shown in Table 1, which also provides population estimates for each characteristic. Not having health insurance was reported by 699 persons (12.8% of the population). Persons older than 55 years were more likely to be uninsured (χ^2, 26.1; P<.001), as were men (χ^2, 33.8; P<.001), blacks (χ^2, 16.7; P<.001), those with less than 12 years of school (χ^2, 73.5; P<.001), those with lower family incomes (χ^2, 114.3; P<.001), the unemployed (χ^2, 78.8; P<.001), those reporting lower self rated health (χ^2, 59.9; P<.001), those with morbidity found on medical examination (χ^2, 18.4; P<.001), and those reporting little or no leisure exercise (χ^2, 10.1; P=.002). There was no statistically significant (P<.05) relationship between insurance and smoking status, alcohol consumption, or obesity.

By the end of the follow-up period, 593 persons, representing 10.7% of the population, had died (Table 1). Those who died were nearly twice as likely to have been uninsured at baseline compared with those who survived (χ^2, 16.5; P<.001). Increased mortality was also associated with being older (χ^2, 228.8; P<.001), being a male (χ^2, 33.8; P<.001), black (χ^2, 8.9; P<.001), having less than 12 years of school (χ^2, 68.7; P<.001), lower income (χ^2, 49.4; P<.001), unemployment (χ^2, 9.8; P=.003), lower self-rated health (χ^2, 73.2; P<.001), morbidity present on baseline medical examination (χ^2, 66.4; P<.001), reporting little or no leisure exercise (χ^2, 8.9; P=.004), smoking (χ^2, 20.1; P<.001), consuming six or more drinks per week (χ^2, 11.0; P=.001), and obesity (χ^2, 4.9; P=.03). Figures 1 through 3 show the relationship between insurance status at baseline and the probability of mortality over the follow-up period, both unadjusted and stratified by income and morbidity. The adverse association between lacking insurance and mortality was observed in all subgroups.

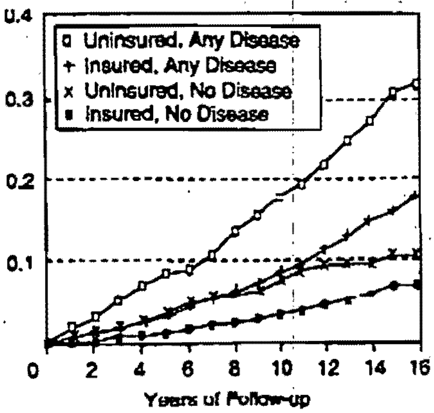

Fig 3.—Cumulative mortality probability by baseline insurance and morbidity categories.

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

COMMENT

This analysis of a nationally representative cohort of the US population suggests that the mortality experience of Americans without health insurance is greater than those with insurance. The effect observed is comparable to that observed for education, income, employment status, and self-rated health. In this study and reported in previous studies, lack of insurance is associated with social and medical factors that increase the risk of poor health.[33-35] However, lacking insurance is associated with subsequent higher mortality independent of other risk factors. The relationship was observed in those with and without baseline morbidity and in those with higher and lower levels of income and education. The absence of statistically significant interactions in the survival analysis between health insurance and the other baseline characteristics suggests that the benefits of health insurance are not confined to particular subgroups. These results are consistent with a previous study of a nationally representative sample suggesting a beneficial association between insurance and health status in persons both below and above 200% of the federal poverty level.[45]

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

Baseline Risk Factor	Hazard Ratio (95% Confidence Interval)
No insurance	1.25 (1.55-1.00)
Age	1.07 (1.09-1.05)
Men	1.90 (2.40-1.50)
Black	1.30 (1.72-1.05)
<12 y school	1.32 (1.61-1.09)
Income	
<$7000	1.26 (1.66-0.96)
$7000-$9999	1.27 (1.67-0.97)
$10 000-$14 999	1.08 (1.48-0.79)
Unemployment	1.21 (1.59-0.99)
Morbidity present	1.43 (1.74-1.18)
Self-rated health	
Fair/poor	1.62 (2.21-1.18)
Good	1.25 (1.61-0.97)
Very good	1.21 (1.59-0.92)
Little or no exercise	1.14 (1.36-0.96)
Present smoker	1.84 (2.15-1.57)
BMI, >27†	1.40 (1.74-1.13)
≥6 drinks per wk	1.46 (1.78-1.20)

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

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

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

Second, the results are consistent with the study hypothesis that a lack of health insurance is causally related to a higher mortality rate, because of decreased access and lower quality of care. This hypothesis is in accordance with the results of previous studies, and the conclusions of the US Congress Office of Technology Assessment report.[4] For the most part, previous studies suggesting the benefits of health insurance have either focused on the poor[27,30,31] or suggested that the benefits of free health care are evident mostly in the poor.[29] Although poorer persons had a higher mortality rate, there was no evidence that the adverse effects of lacking insurance was limited to this group. Our analysis compared any level of insurance with no insurance and included a representative sample of adults not covered by publicly funded insurance programs; thus, the analysis may have been more sensitive to the effect of insurance than the Rand Health Insurance Experiment.[29]

The results obtained probably underestimate the relationship between insurance and mortality. First, many persons in the insured group may be underinsured. Farley found that over a quarter of insured persons under 65 years of age were inadequately protected against the possibility of large medical bills.[47] Second, insurance status was measured only once, at baseline, and no account could be taken of the impact of changes in insurance status over time. Migration of persons with and without insurance to the other group would tend to bias the observed association toward zero. Most persons older than 50 years at the beginning of the study would have been eligible for Medicare by the end of the follow-up period. Significant shifts in insurance status also occur during the course of a year. In the 1987 National Medical Expenditure Survey, 22.4% of persons younger than 65 years were uninsured at some time during the year, 16.1% to 17.2% were uninsured at any one period during the year, and 11.4% were uninsured throughout the year. Third, the variables included in the survival analysis may have resulted in overadjusting of the true relationship between insurance status and mortality. Over time, employment status may be a better measure of insurance than self-reported insurance status, since most persons obtain their insurance through employment.[46] Also, if lacking insurance adversely affects health, then adjusting for baseline health status will tend to reduce the observed relationship between health insurance and mortality. When the baseline employment and health status variables were excluded from the analysis, the relationship between insurance and mortality was increased. Although observational studies cannot exclude the possibility that unmeasured underlying variables explain the relationship between lacking health insurance and adverse health outcomes, it is unlikely that a randomized trial will be conducted to address this important policy issue. Our analysis, consistent with previous studies suggesting an adverse impact of lacking insurance on health, provides evidence that lacking insurance also increases mortality in a nationally representative sample of adults. The existing "safety net" has failed to prevent this increased mortality. The findings support a policy imperative for universal health insurance to reduce both financial barriers to care and the risk of premature mortality.

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Talking Points on Estimates of the Uninsured

Differences in estimates of the number of uninsured reflect differences in methodolgy and interpretation.

- The Administration estimates the number of uninsured to be about 40 million people in 1993. The estmates are based on data from the Current Population Survey (CPS), adjusted to reflect the 1990 census.

- The Urban Institute's estimate of the number of uninsured differs from the Administration estimate primarily for two reasons. First, the Urban Institute estimate is not based to the 1990 census -- leading to a lower count of the number of uninsured. The Urban Institute also adjusts the CPS data to account for a perceived undercounting in the CPS of the number of people on Medicaid. They also project their estimates forward to 1994. They estimate the number of uninsured at about 36 million people in 1994.

- The Employee Benefit Research Institute's estimate differs from the Administration's estimate primarily because EBRI makes a downward adjustment in the number of insured children on the CPS to account for a perceived inconsistency between two questions on the survey. EBRI estimates the number of uninsured to be 41 million people.

NUMBER OF UNINSURED AMERICANS

- There were approximately 40 million Americans without health insurance in 1993. This was about 15% of the U.S. population.

- The number uninsured Americans is growing:

  - The number of uninsured Americans grew from about 30 million people in 1979 to about 40 million people in 1993.
  - The number of uninsured Americans is currently growing by about 1 million people each year.

- The erosion of employer sponsored health insurance is part of the reason for the growth in the number of uninsured Americans.

  - Between 1989 and 1993, the number of Americans with employer-sponsored health insurance fell from 152 million to 148 million.

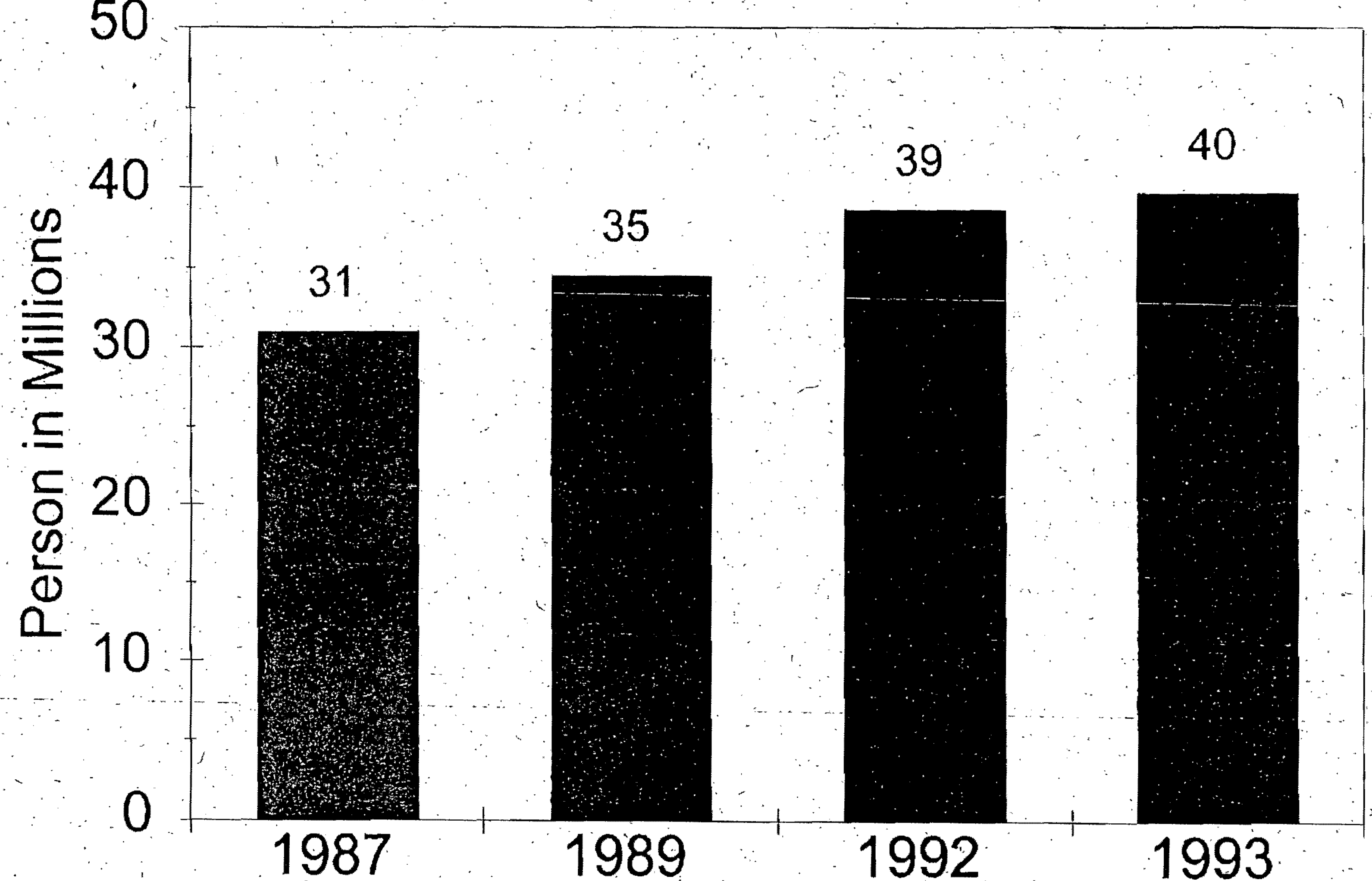
Notes:

Administration estimates based on the Current Population Survey (CPS).

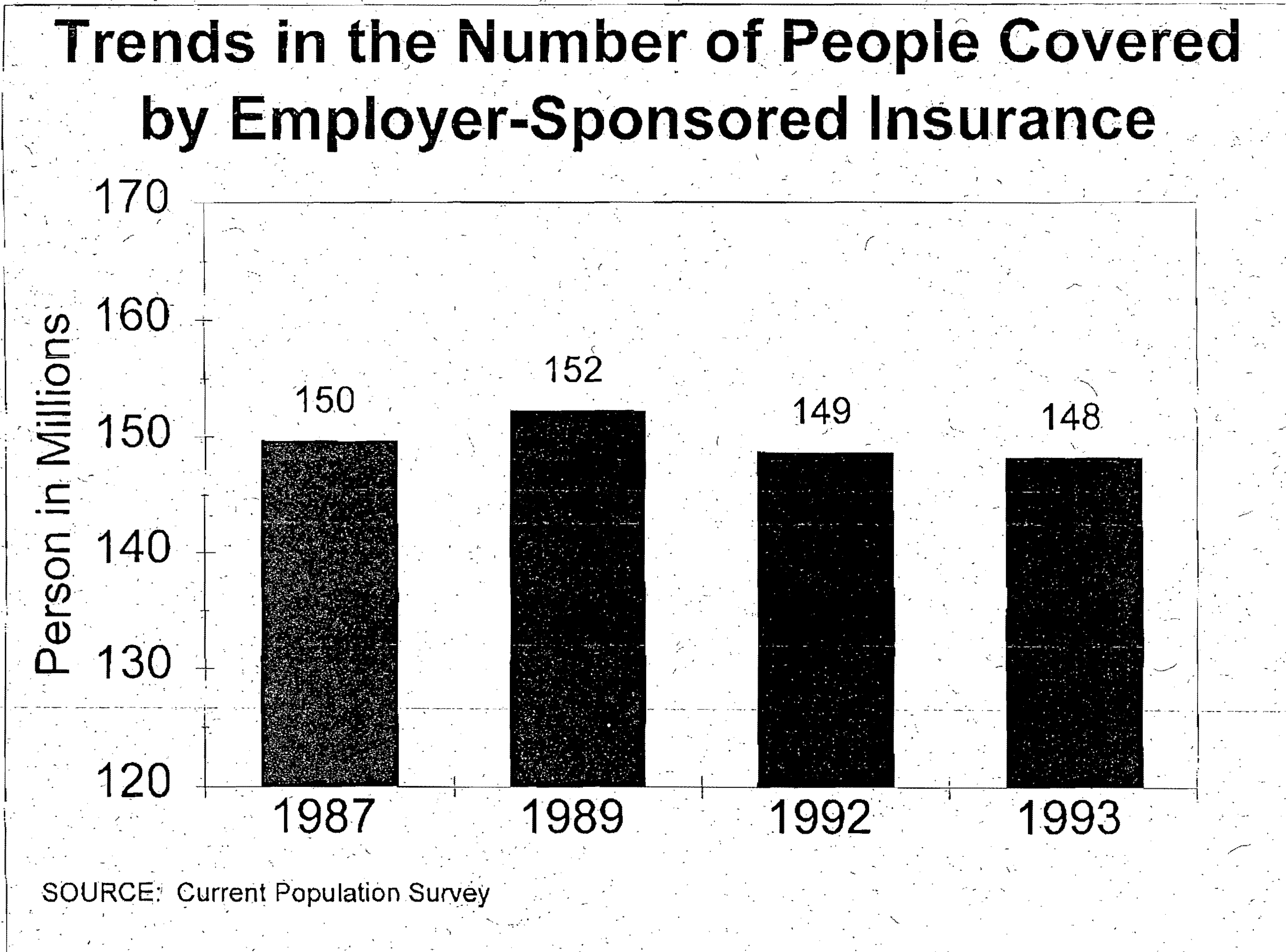
There are differences in estimates of the number of uninsured which reflect differences in methodology and data interpretation. For example, the Urban Institute has a lower estimate (36 million people in 1994) of the number of uninsured primarily for two reasons. The first is that the Urban Institute is not adjusted to the 1990 census, which produces a lower count of the uninsured. Second, the Urban Institute adjusts its estimates to account for a perceived under reporting of the number of people covered by Medicaid. The Employee Benefit Research Institute (EBRI) has a higher estimated of the number of uninsured (41 million in 1993) because they make a downward adjustment in the number of insured children to account for a perceived inconsistency between two questions on the CPS.

Changes in the CPS design in 1988 produce inconsistencies in insurance coverage information before and after 1987. Beginning in 1988, the CPS asked all respondents (rather than just employed people) whether they were insured under employer-sponsored plans. In addition, method of counting the number of insured children was improved.

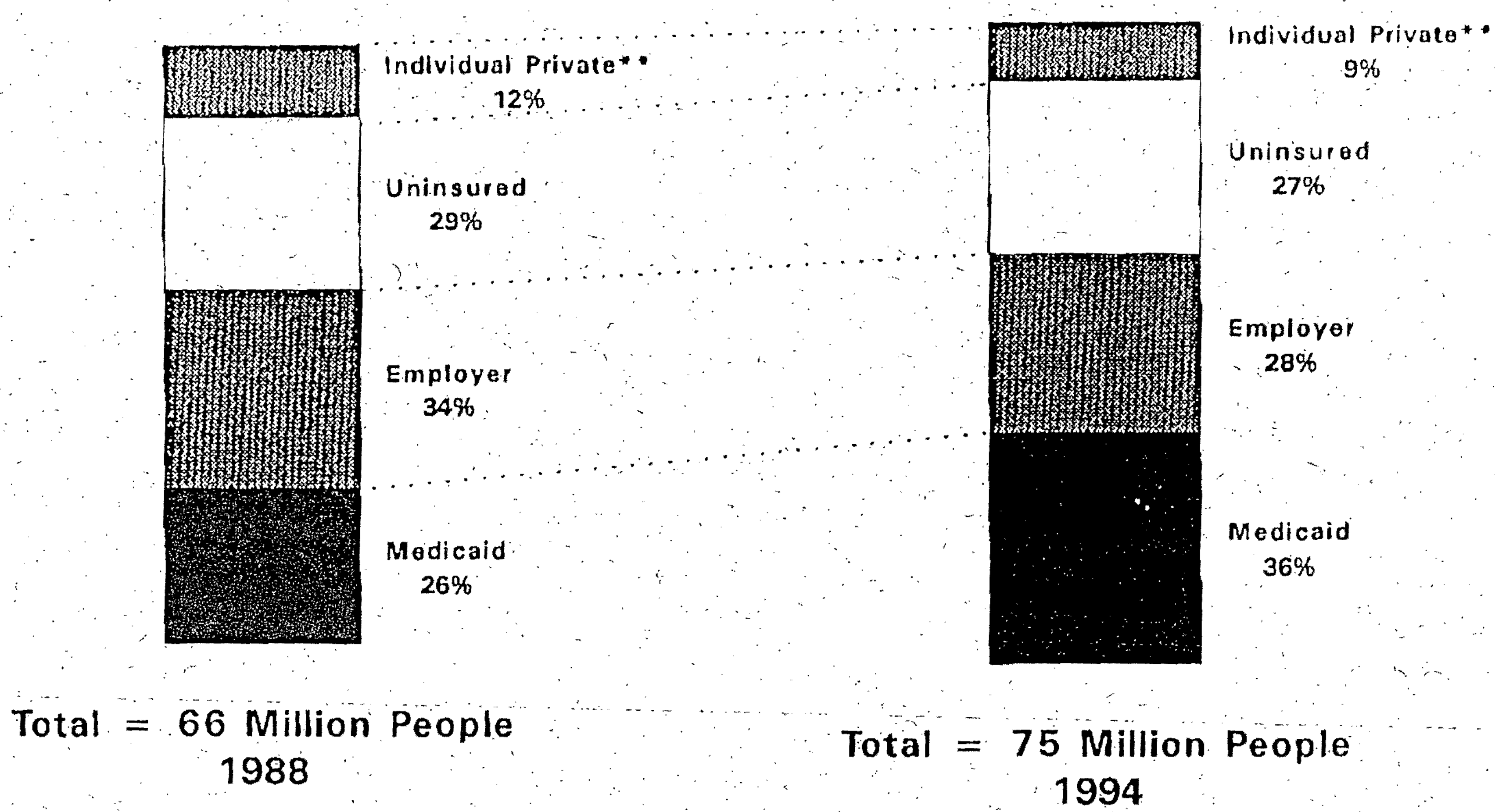
Trends in the Number of Uninsured



SOURCE: Current Population Survey


Trends in the Number of People Covered by Employer-Sponsored Insurance
170
160
150
140
130
120
Person in Millions
150
152
149
148
1987
1989
1992
1993
SOURCE: Current Population Survey

Trends in Health Insurance Coverage for Low-Income Population*, 1988 and 1994



* Below 200 percent of the Federal poverty level.

** Includes coverage for the military and veterans.

Note: The Federal poverty level was $14,800 for a family of four in 1994.

Source: Urban Institute estimates based on 1988 and 1992 Current Population Surveys, 1994.

The Kaiser Commission on

THE FUTURE OF MEDICAID

HEALTH INSURANCE OVER TIME FROM THE MARCH CPS
MAY 17, 1995

PERSON COUNTS IN MILLIONS AS OF MARCH OF THE YEAR

YEAR	UNINSURED	% OF POPULATION	AVERAGE INCREASE	COMPOUND RATE OF GROWTH	EMPLOYER GROUP INSURANCE	% OF POPULATION	AVERAGE INCREASE	COMPOUND RATE OF GROWTH
1994R	39.719	15.3%	1.078	2.8%	148.318	57.1%	-0.478	-0.3%
1993R	38.641	15.0%	1.365	3.8%	148.796	57.9%	-1.181	-0.8%
1990R	34.546	13.9%	-0.502	-1.4%	152.338	61.3%	3.616	2.6%
1985I	37.055	15.8%	1.316	4.0%	134.257	57.4%	-0.180	-0.1%
1980Z	30.474	13.7%			135.157	60.6%		

R: BENCHMARKED ON THE 1990 CENSUS INCLUDING UNDERCOUNT.
I: BENCHMARKED ON THE 1980 CENSUS WITHOUT THE UNDERCOUNT.
Z: BENCHMARKED ON THE 1970 CENSUS WITHOUT THE UNDERCOUNT.

THE 1985-90 CHANGES ARE SPURIOUS BECAUSE OF MAJOR CHANGES IN THE INSURANCE QUESTIONS ASKED ON THE SURVEY.

OPTIONAL FORM 99 (7-90)
FAX TRANSMITTAL
# of pages ▸
To
From
Dept./Agency
Phone #
Fax #
Fax #
NSN 7540-01-317-7368 5099-101 GENERAL SERVICES ADMINISTRATION

DEPARTMENT OF HEALTH AND HUMAN SERVICES
ASSISTANT SECRETARY FOR PLANNING AND EVALUATION
OFFICE OF HEALTH POLICY

PHONE: (202) 690-6870 FAX: (202) 401-7321

Date:

From: Jeanne

To: CHRIS JENNINGS

Phone: (202) 690-
(202) 690-6870
FAX: (202) 401-7321

Phone:

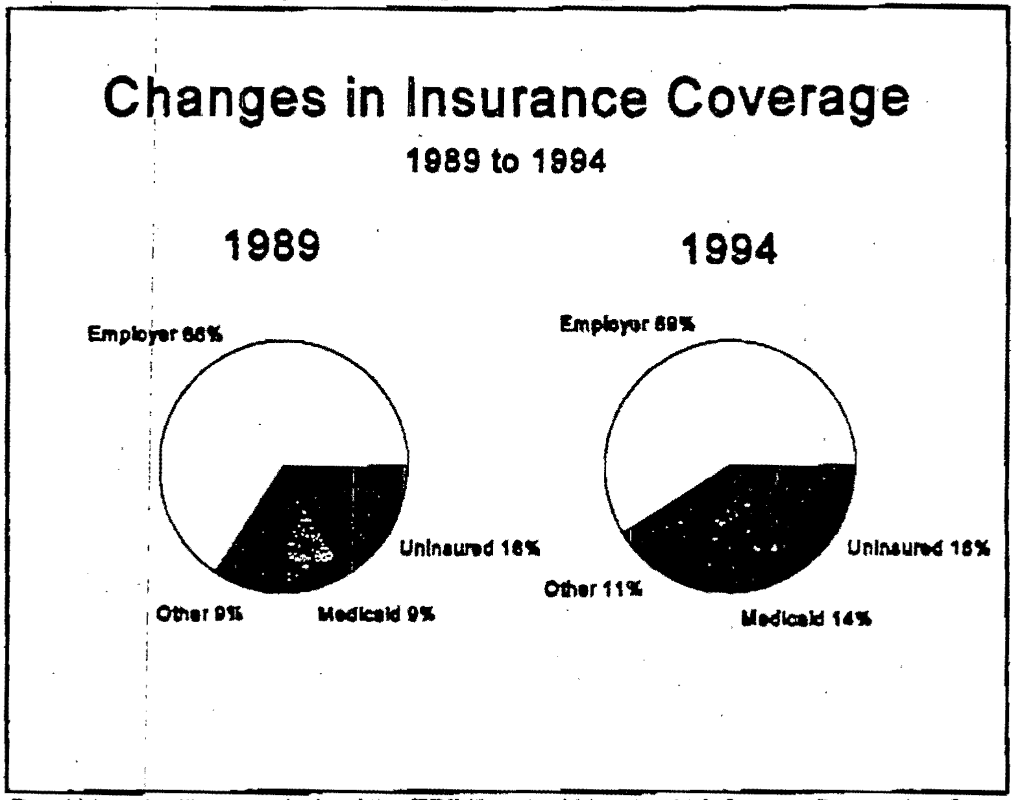
Fax:

Number of Pages (Including Cover):

Comments: ATTACHED ARE A COUPLE TABLES. I DIDN'T WRITE POINTS BECAUSE I REMEMBERED LAST NIGHT THAT LARRY DID THIS A MONTH OR SO AGO. AS SOON AS GARY GETS IN, I'LL FIND THOSE POINTS + FAX THEM.

ALSO: (1) We are working on beautifying the MCR BENEFICIARY + PROVIDER HIT CHARTS - ANY COMMENTS?
(2) SOMEBODY NEEDS TO GIVE THE POST EDITORIAL BOARD A MEDICAID LESSON!

CHANGES IN MEDICAID ENROLLMENT



SOURCE: The Urban Institute analysis of the TRIM2-edited March 1993 Current Population Survey.

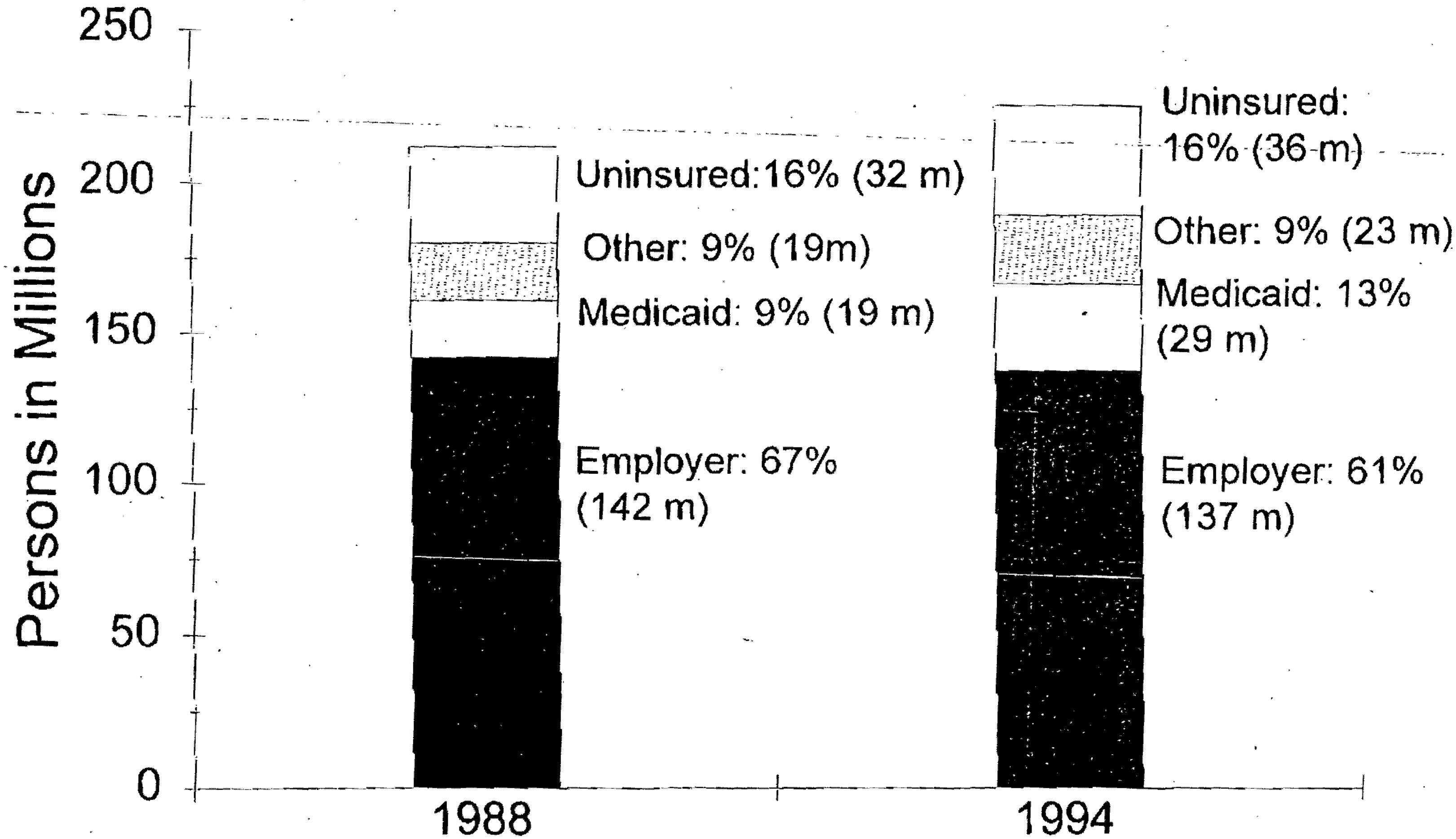
- Medicaid has been a significant and growing source of health insurance for many people.

    - Between 1989 and 1994, the percentage of the population covered by Medicaid grew from 9% to over 14%, while the percentage covered by private health insurance fell from about 66% to about 59%.

    - Without this growth in Medicaid, the number of uninsured would likely have increased significantly.

- Additional Republican proposals to significantly cut federal Medicaid payments through a block grant would likely exacerbate the loss of Medicaid coverage. The magnitude of the suggested cuts would leave states with little choice but to reduce eligibility and benefits.

Change in Insurance Coverage: 1988-94

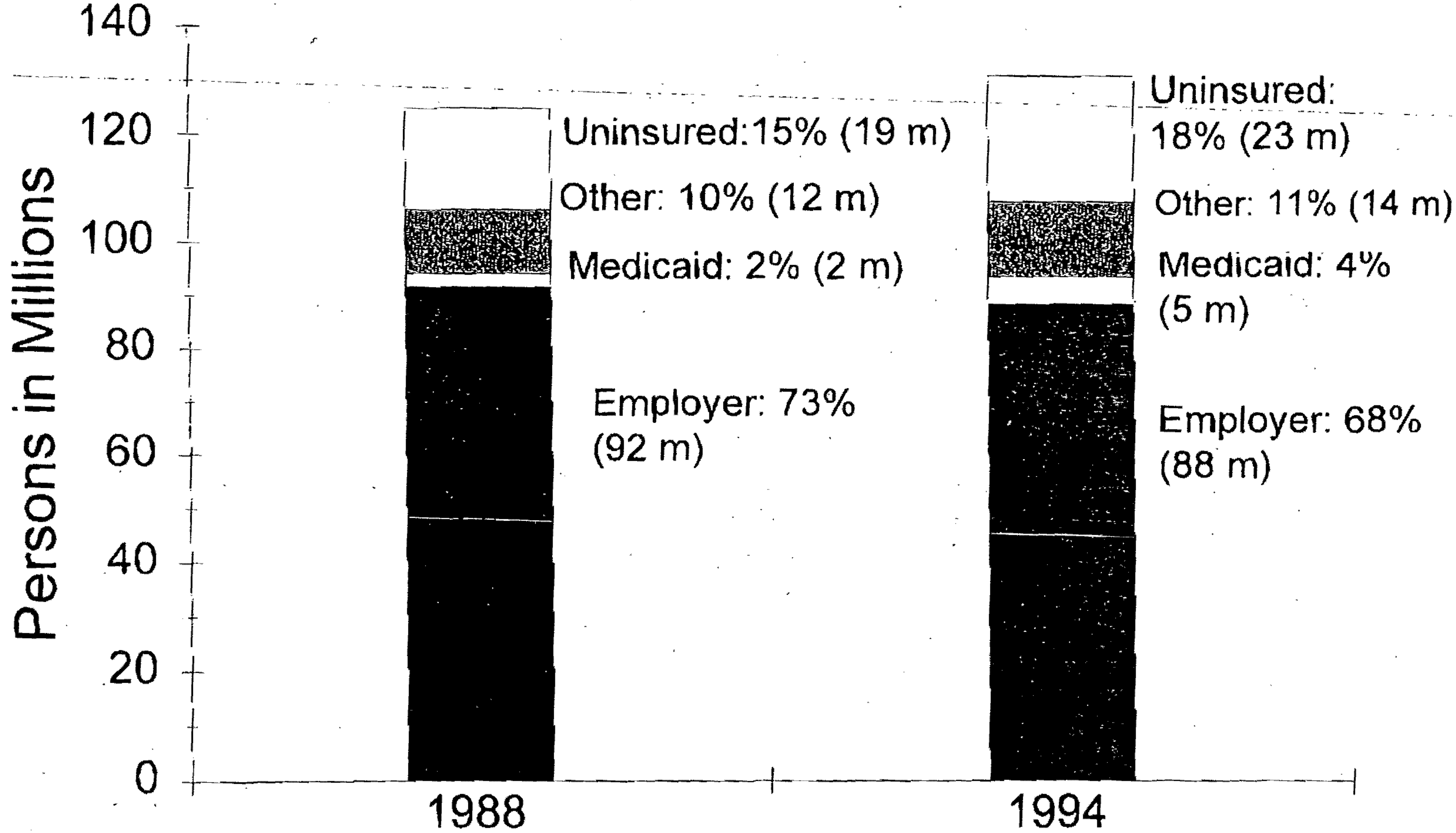


SOURCE: THE URBAN INSTITUTE

Nonelderly only; Number of uninsured does not match total of 40 million on CPS

Change in Insurance Coverage: 1988-94

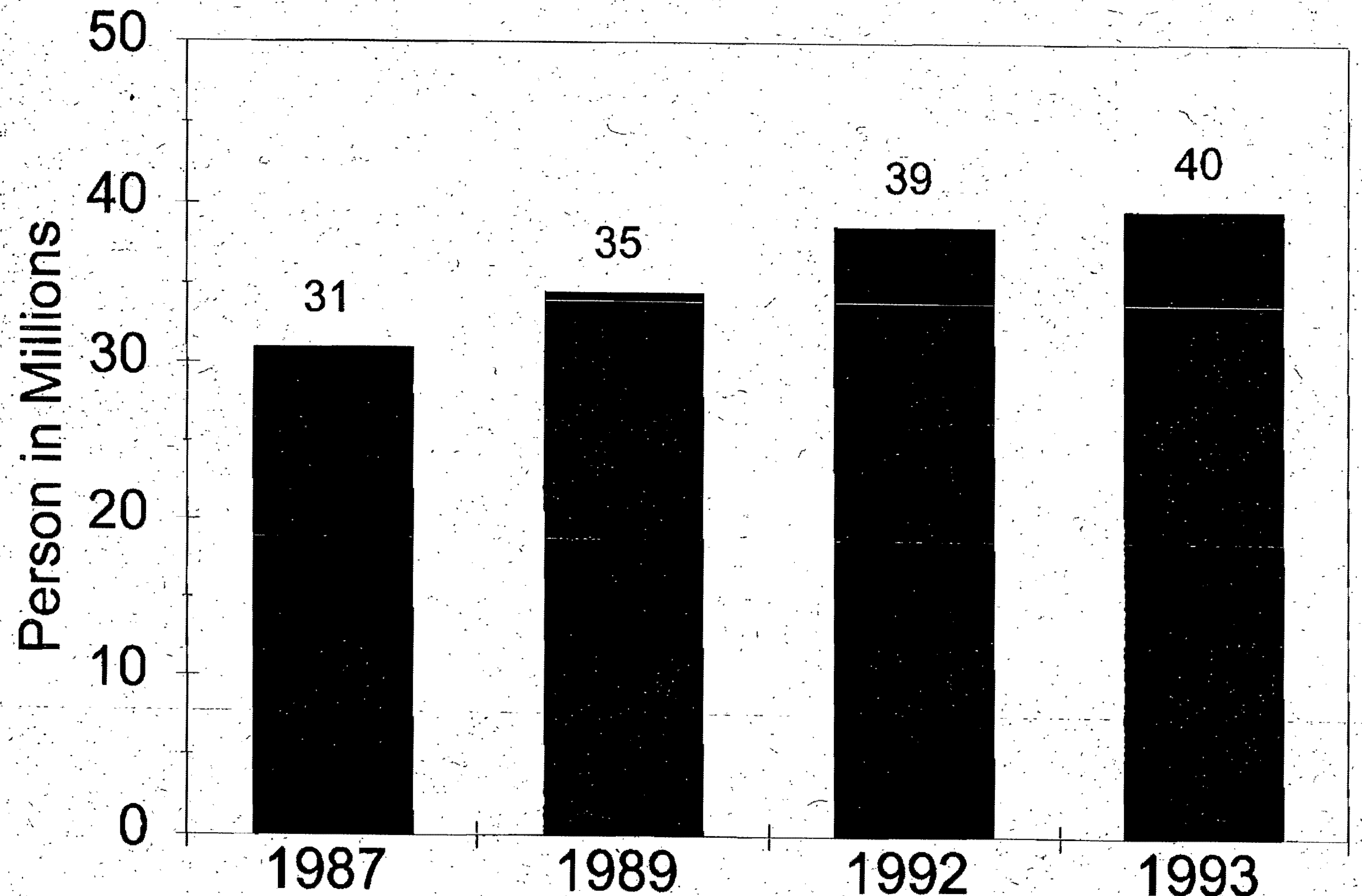
Workers



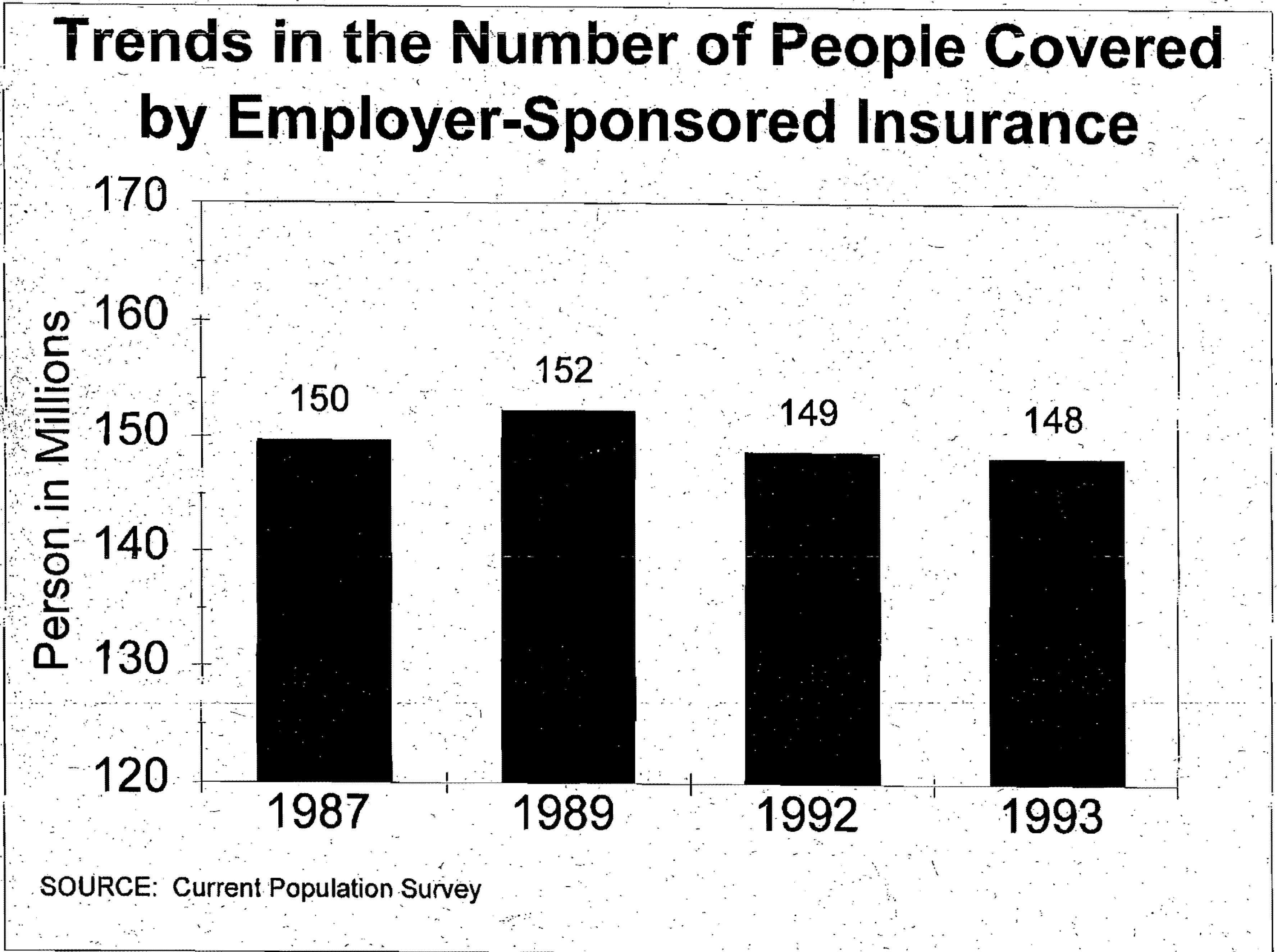
SOURCE: THE URBAN INSTITUTE

Nonelderly only; Number of uninsured does not match total of 40 million on CPS

Trends in the Number of Uninsured



SOURCE: Current Population Survey


Trends in the Number of People Covered by Employer-Sponsored Insurance
Person in Millions
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150
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1993
SOURCE: Current Population Survey

DEPARTMENT OF HEALTH AND HUMAN SERVICES
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OFFICE OF HEALTH POLICY

DEPARTMENT OF HEALTH & HUMAN SERVICES · USA

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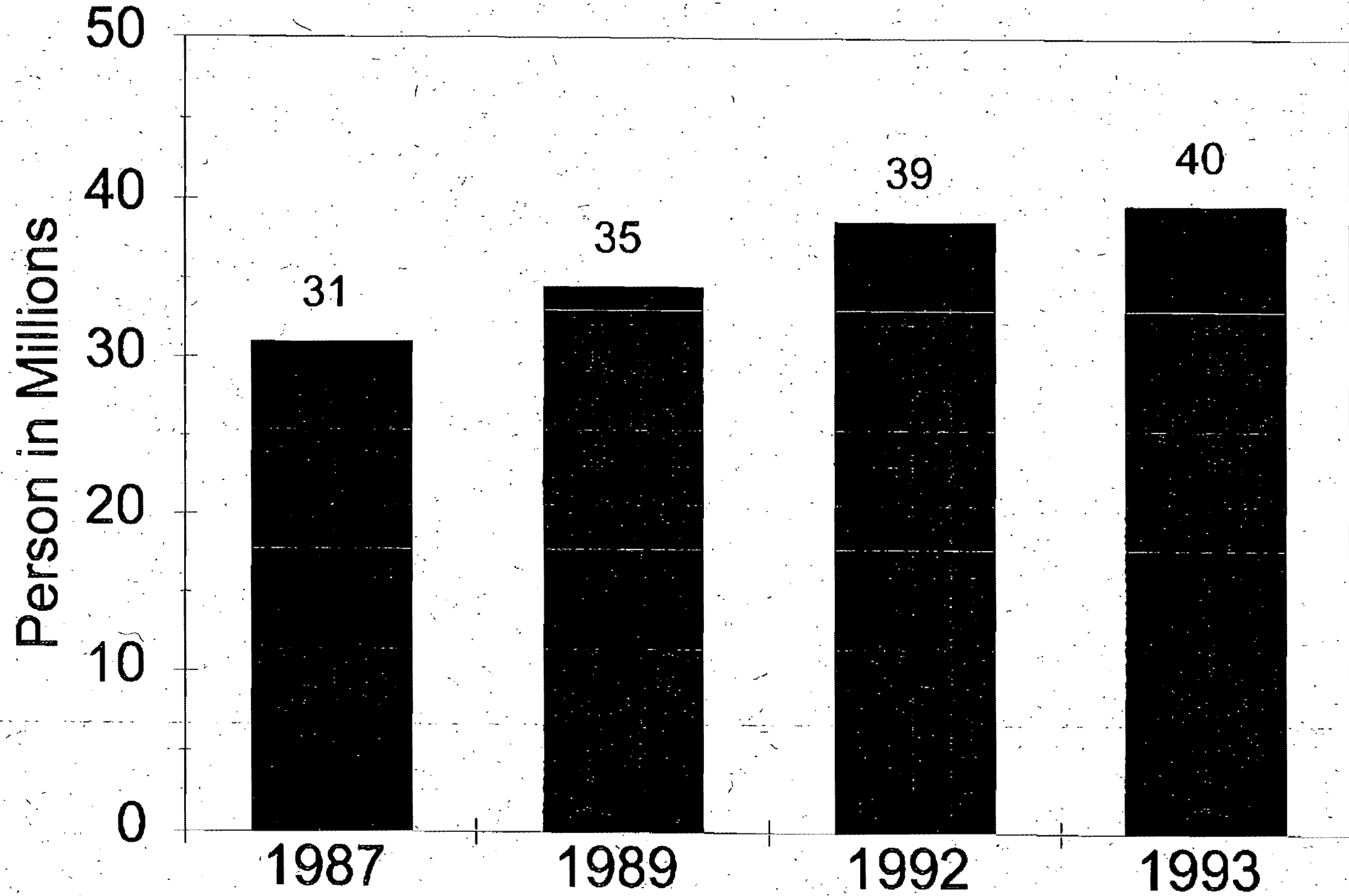
Number of Pages (Including Cover):

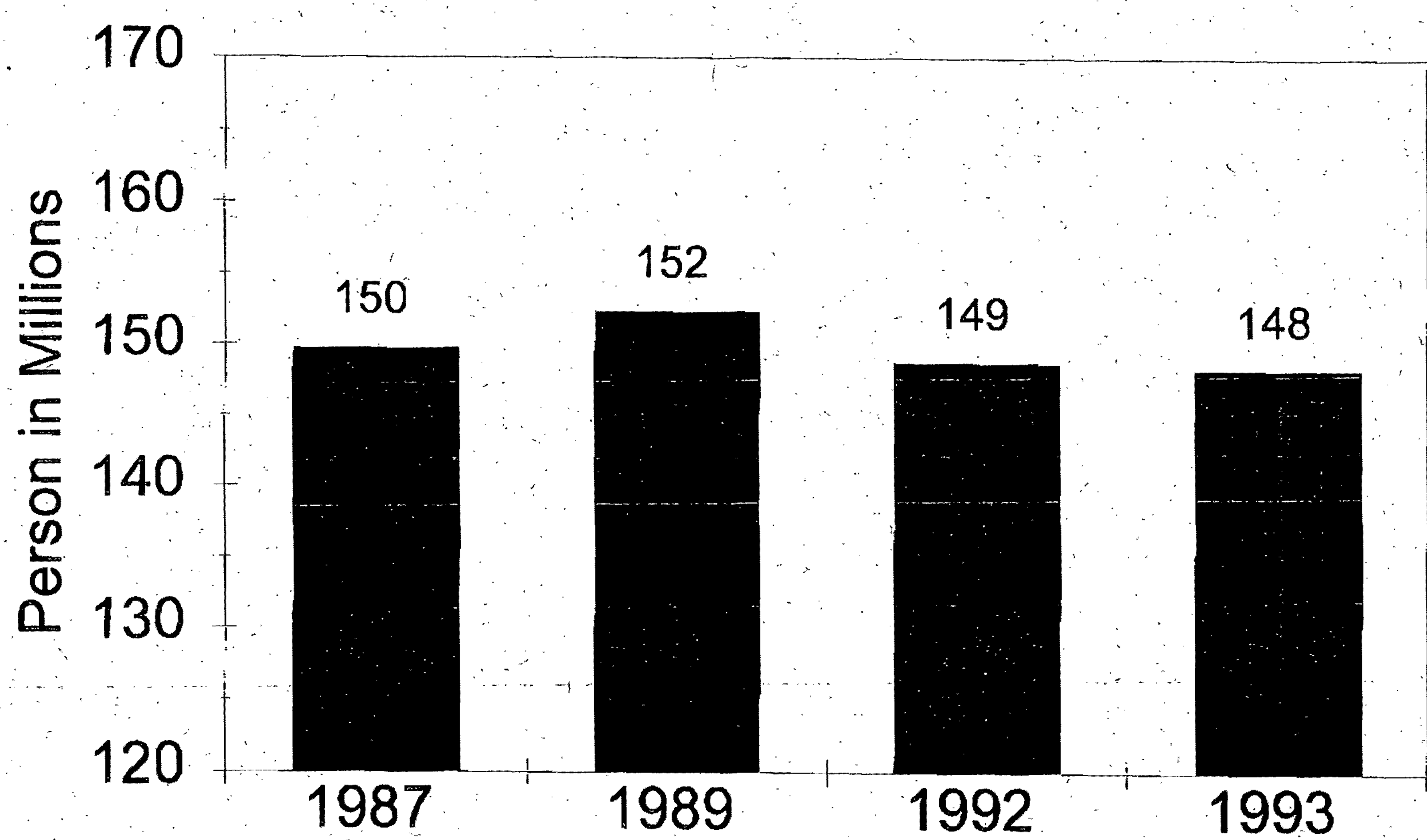
Comments: I've added 87 - it's the Best we can do.

Trends in the Number of Uninsured



SOURCE: Current Population Survey

Trends in the Number of People Covered by Employer-Sponsored Insurance



SOURCE: Current Population Survey

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OFFICE OF HEALTH POLICY

DEPARTMENT OF HEALTH & HUMAN SERVICES · USA

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Number of Pages (Including Cover):

Comments: Uninsured Stuff.

NUMBER OF UNINSURED AMERICANS

- There were approximately 40 million Americans without health insurance in 1993. This was about 15% of the U.S. population.

- The number uninsured Americans is growing:

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    - Between 1989 and 1993, the number of Americans with employer-sponsored health insurance fell from 152 million to 148 million.

Notes:

Administration estimates based on the Current Population Survey (CPS).

There are differences in estimates of the number of uninsured which reflect differences in methodology and data interpretation. For example, the Urban Institute has a lower estimate (36 million people in 1994) of the number of uninsured primarily for two reasons. The first is that the Urban Institute is not adjusted to the 1990 census, which produces a lower count of the uninsured. Second, the Urban Institute adjusts its estimates to account for a perceived under reporting of the number of people covered by Medicaid. The Employee Benefit Research Institute (EBRI) has a higher estimated of the number of uninsured (41 million in 1993) because they make a downward adjustment in the number of insured children to account for a perceived inconsistency between two questions on the CPS.

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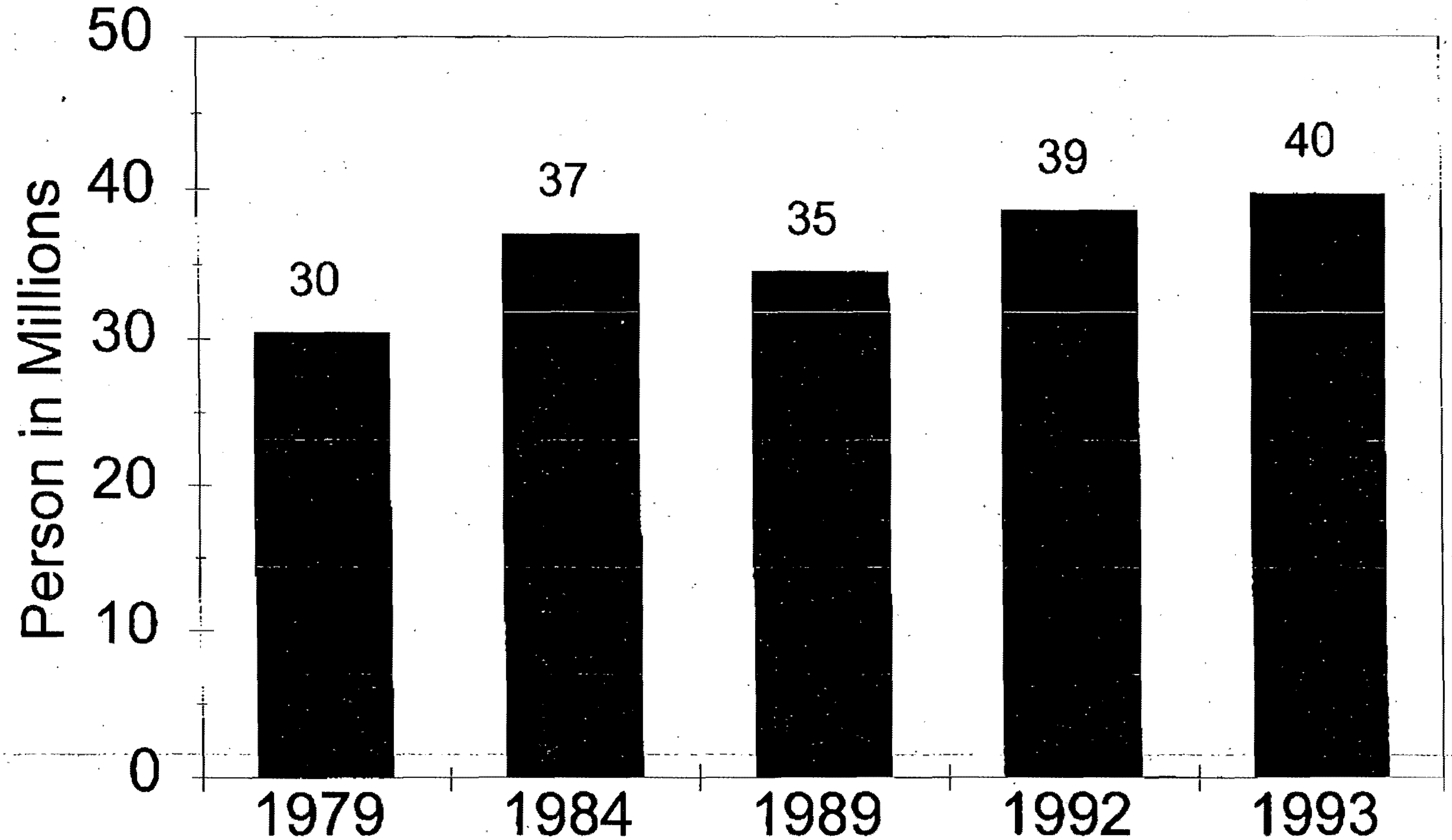
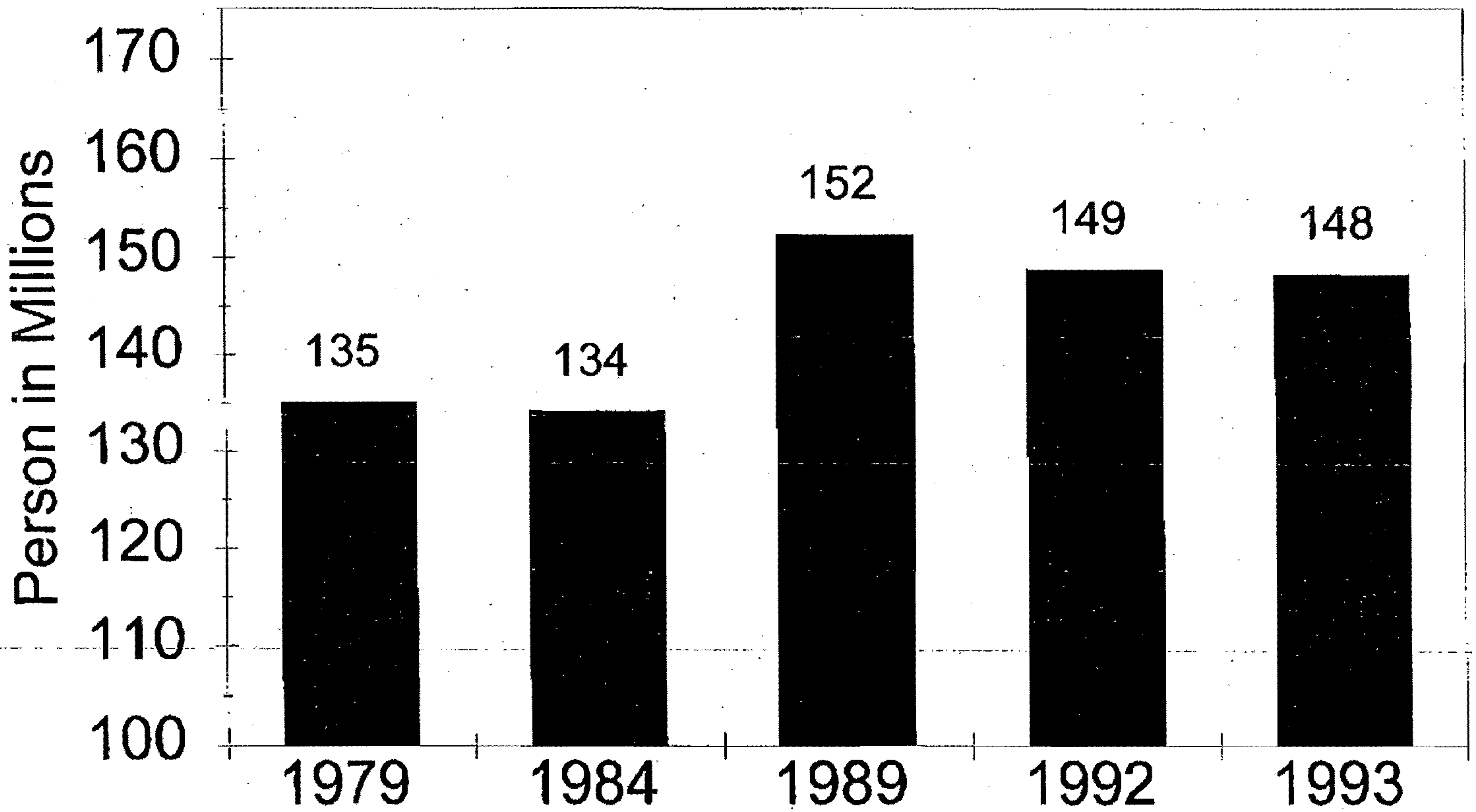
Trends in the Number of Uninsured



SOURCE: Current Population Survey;
Note: changes in the survey in 1986 account for the inconsistent trend between 1984 & 1989

Trends in the Number of People Covered by Employer-Sponsored Insurance



SOURCE: Current Population Survey;
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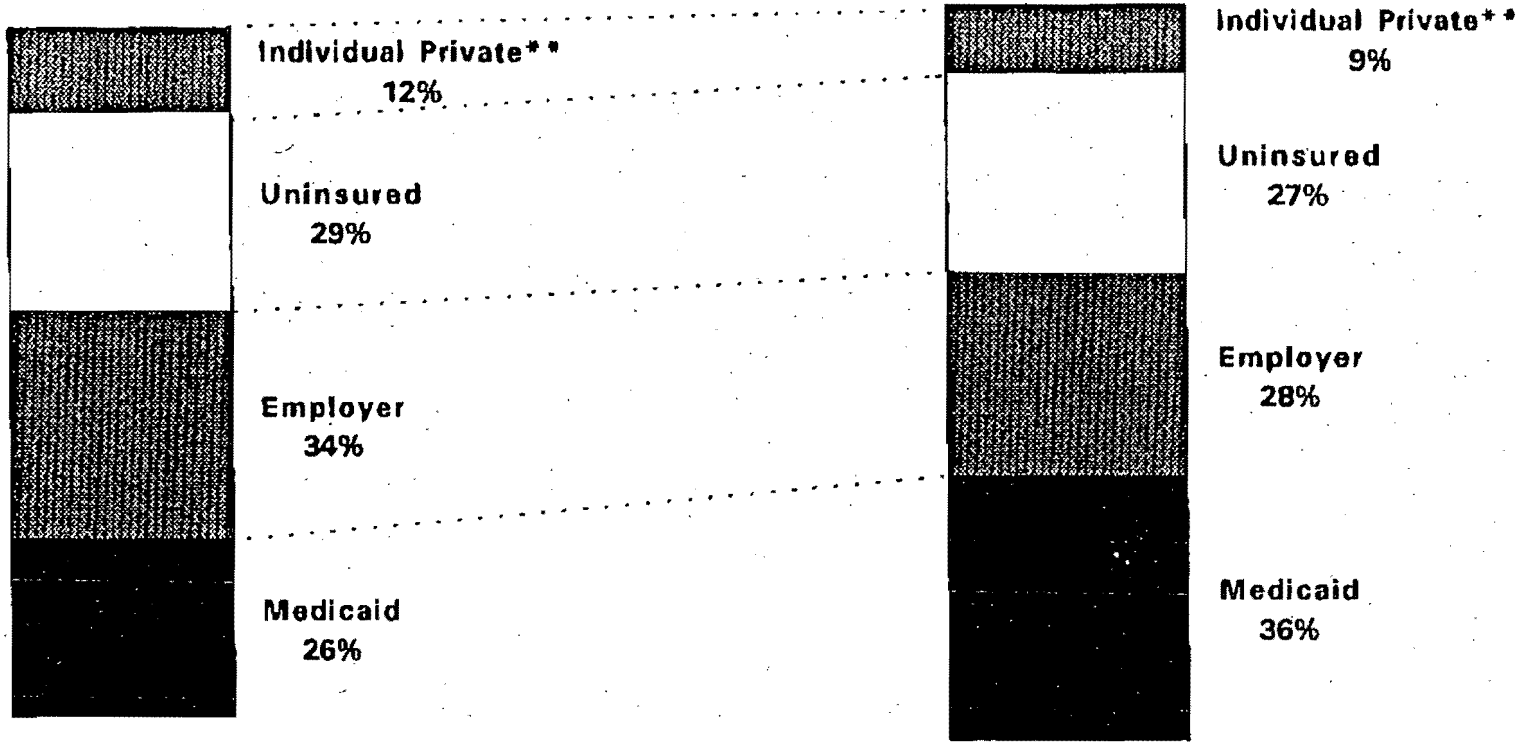
Trends in Health Insurance Coverage for Low-Income Population*, 1988 and 1994



Total = 66 Million People
1988

Total = 75 Million People
1994

* Below 200 percent of the Federal poverty level.

** Includes coverage for the military and veterans.

Note: The Federal poverty level was $14,800 for a family of four in 1994.

Source: Urban Institute estimates based on 1988 and 1992 Current Population Surveys, 1994.

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OFFICE OF HEALTH POLICY

DEPARTMENT OF HEALTH & HUMAN SERVICES · USA

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(202) 690-6870
FAX: (202) 401-7321

Phone:

Fax:

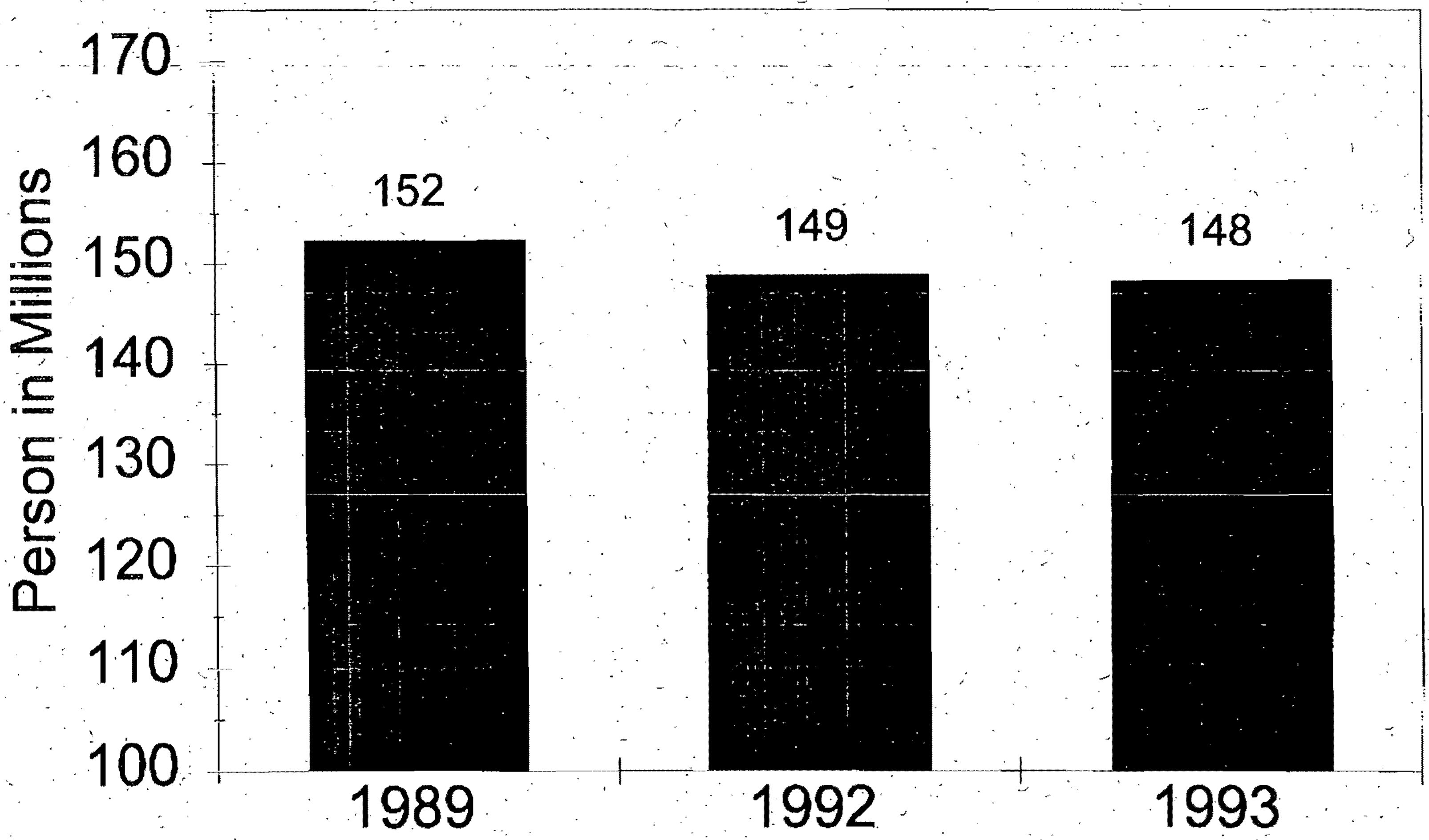
Number of Pages (Including Cover):

Comments: Exciting New Charts.

REVISED —

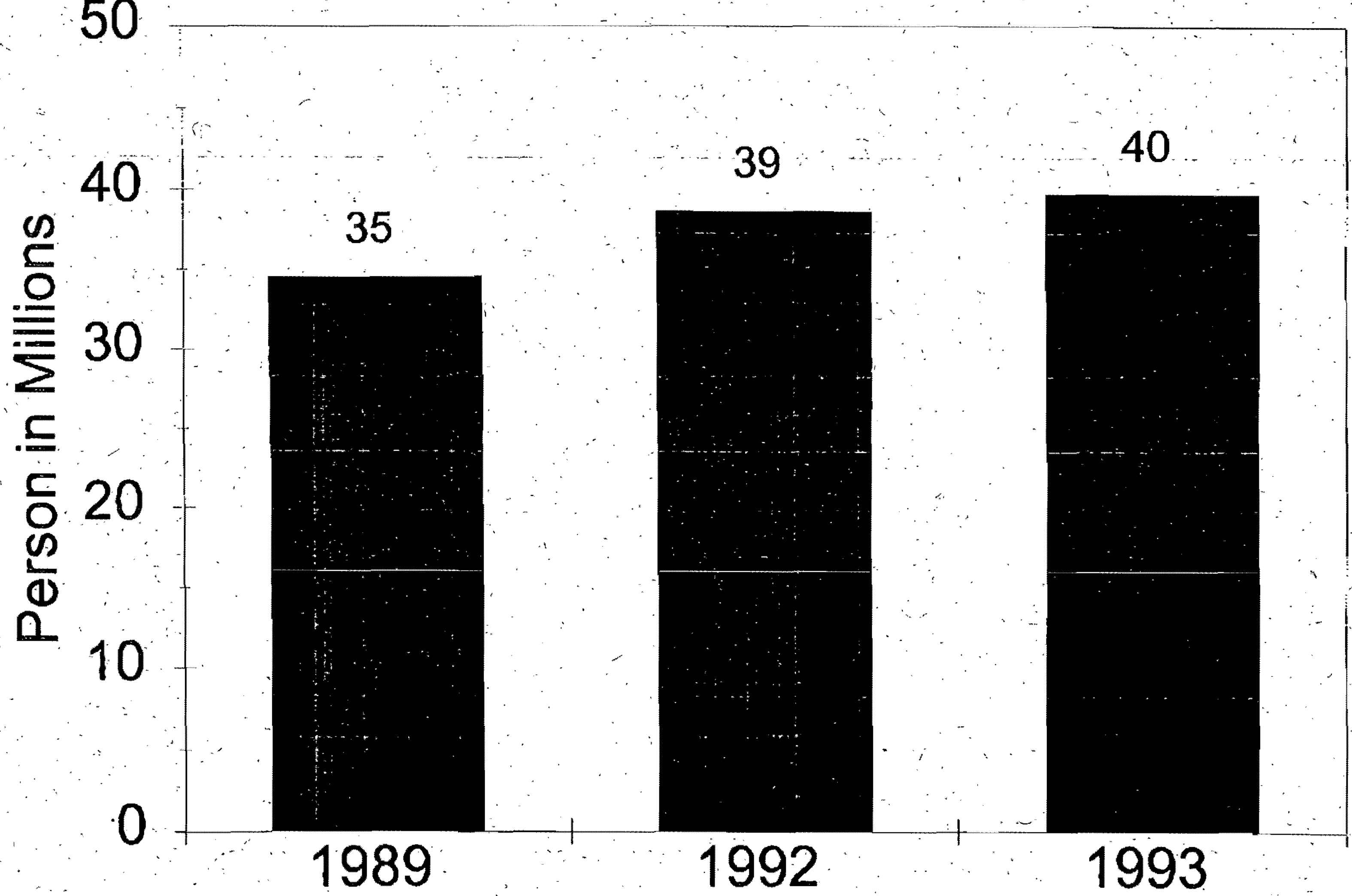
NOTE: I'm trying to get the 20→91 estimates

Trends in the Number of People Covered by Employer-Sponsored Insurance



SOURCE: Current Population Survey

Trends in the Number of Uninsured



SOURCE: Current Population Survey

American
Psychiatric
Association

1400 K Street, N.W.
Washington, D.C. 20005
Telephone: (202) 682-6000

6-25-97

Dear Sarah,

Thanks for following up with me today re the Domenici-Wellstone amendment on mental health parity for children.

Enclosed is a copy of a House dear colleague sent today as well as our actuarial study the we discussed.

We appreciate the support of the President behind this effort and particularly during the conference committe.

Sincerely,

Julie Shroyer

File Mental Health

Congress of the United States
House of Representatives
Washington, DC 20515

June 25, 1997

Dear Colleague:

Yesterday, the Senate approved the Dominici-Wellstone-Reid-Conrad amendment to require parity coverage of mental health care for children as part of the Senate child health reconciliation package. This is another major step toward ending discrimination in insurance coverage of mental illness treatment.

We will urge the House to accept this important amendment in conference with the Senate, and we encourage you to join in these efforts. To that end, we are pleased to bring to your attention the attached report on mental health parity costs for children, prepared by the nationally recognized firm of Milliman & Robertson. The most important finding is that we can afford children's mental health parity now, with projected costs ranging from at most 3.7 percent, to as little as 0.3 percent. Actual costs are likely to fall in the lower end of the range.

The Milliman & Robertson study follows on the heels of an April, 1997 report to Congress by the National Advisory Mental Health Council, which found that actual mental health parity experience in Texas, Maryland and Rhode Island had caused a negligible increase in premiums, or no increase at all, to whit: "the overall increase in premium cost, if any, appears to be less than 1 percent."

Children are our most precious national resource, and study after study has shown that there is a serious need for better access to comprehensive mental health care. The more we learn about mental illness treatment and insurance parity, the clearer it becomes that treatment is both effective and cost-effective. We commend the Senate for its latest effort to curb insurance discrimination against mental health care, and urge you to join us in supporting parity in the House and in conference.

Sincerely,

Marge Roukema
Marge Roukema

Bob Wise
Bob Wise

Peter DeFazio
Peter DeFazio

Coalition for Fairness in Mental Illness Coverage

June 25, 1997

Dear Representative:

The Coalition for Fairness in Mental Illness Coverage, a coalition of organizations representing patients, family groups, and health care systems and providers, is pleased to transmit to you the attached report on the cost of including mental illness parity as an addition to the children's health initiative now being considered by the Congress.

The report, prepared by the nationally respected actuarial firm of Milliman & Robertson, shows clearly that a non-discriminatory (parity) mental health standard is affordable, with projected costs measured against baseline standards ranging from a high of 3.7 percent (loosely managed FEHBP standard plan including substance abuse) to a low of 0.3 percent (tightly managed Washington State plan including substance abuse).

Relevant variables measured include whether substance abuse is covered and whether the plan is loosely, moderately, or tightly managed. Parity for children could be implemented in a tightly managed FEHBP standard plan for as little as 1.9 percent (not including substance abuse which would add a mere 0.1 percent to the cost of parity in this example).

Yesterday, the Senate approved a children's mental health parity amendment to the reconciliation bill sponsored by Senators Pete Domenici, Paul Wellstone, Harry Reid, and Kent Conrad. We believe that children's mental health parity is necessary and we urge you to support this amendment when the House and Senate go to conference. The amendment is urgently needed. The Finance Committee proposal, for example, uses an actuarial average FEHBP plan as a standard, but this would leave large gaps in children's mental health coverage, since none of the plans now offers parity coverage.

On behalf of the millions of children who require mental health care, we urge you to support children's mental health parity amendment of the budget reconciliation package. Parity for children is the right thing to do. It is also clearly affordable.

National Alliance for the Mentally Ill
National Mental Health Association
American Managed Behavioral Health Association
American Medical Association
American Psychiatric Association
American Psychological Association
Federation of American Health Systems
National Association of Psychiatric Health Systems

Mental illness coverage. It's time to be fair by treating it equally in health care.

1400 K Street, NW, Third Floor, Washington, DC 20005
Phone: 202-682-6393 Fax: 202-682-6287

Premium Rate Estimates for Parity Mental Illness Insurance Coverage for Children

by

Stephen P. Melek, FSA, MAAA
Bruce S. Pyenson, FSA, MAAA

Milliman & Robertson, Inc.
June 20, 1997

II. Premium Rate Estimates for Parity Coverage for Children

A behavioral health parity provision would eliminate benefit limitations specific to behavioral health disorders, and it would require that beneficiary cost sharing provisions for such services equal those for non-behavioral care. These plan changes would also increase costs somewhat.

Managed care has resulted in greater cost reductions for behavioral health than for other kinds of health services. In order to fairly present the cost effects of behavioral health parity, we show results for relatively loosely managed utilization, for moderately managed utilization and for aggressively managed utilization.

The following table presents the results of our cost estimates:

Table I	Percent Increase in Premium Rate Estimates of Benefit Plans - Loosely Managed Delivery System		
Benefit Plan	Parity for Mental Health	Parity for Substance Abuse	Parity for both MH and SA
Federal Employee Health Benefits - Standard Plan	3.4%	0.4%	3.7%
Washington State Health Benefits Plan	2.1%	0.2%	2.3%

Table II	Percent Increase in Premium Rate Estimates of Benefit Plans - Moderately Managed Delivery System		
Benefit Plan	Parity for Mental Health	Parity for Substance Abuse	Parity for both MH and SA
Federal Employee Health Benefits - Standard Plan	2.8%	0.3%	3.1%
Washington State Health Benefits Plan	1.4%	0.1%	1.5%

APPENDIX I

Assumptions and Limitations

This section describes key assumptions and sources for our premium rate estimates, including cautions about how the estimates should be interpreted and used.

We estimated costs for children in the currently-insured commercial population. This does not include individuals covered by Medicare or Medicaid. We used standard Milliman & Robertson, Inc. demographic assumptions, intended to represent the age and sex mix of children for a typical employee group with the demographics of the US labor force population.

The plan provisions that we used are summarized in Appendix II.

The starting point for our cost increase estimates is the M&R *Health Cost Guidelines* (*HCGs*) (1997 edition). The *HCGs* are M&R's proprietary information base that shows how the components of per-capita medical claim costs vary with benefit design, demography, location, provider reimbursement arrangements, degree of managed care delivery, and other factors. In most instances, cost assumptions are based on our evaluation of several data sources and are not specifically attainable to a single source. The *HCGs* are used by client insurance companies, HMOs, and other organizations for, primarily, pricing and evaluating insurance products.

We incorporated estimates of the effects of managed care delivery in our cost increase estimates. We utilized the M&R *Healthcare Management Guidelines* (*HMGs*) in developing our assumptions of the impact on utilization and service intensity in managed care scenarios. The *HMGs* describe loosely and aggressively managed care scenarios as follows:

- Loosely managed delivery: Typical of most fee-for-service, indemnity plans which use some pre-admission certification, concurrent review, and hospital audit.

- Moderately managed delivery: Typical of plans which use some inpatient care management protocols sand standards with moderate conformity to those standards.

- Aggressively managed delivery: Typical of very aggressive HMOs which employ very close conformity to inpatient care and ambulatory management standards such as those contained in the HMGs.

For each plan, we developed results assuming a loosely managed and a well managed delivery system.

We assume that the health plan could negotiate physician reimbursement consistent with 100% of RBRVS. Hospital per diems were assumed to be $1,100, except for mental health and substance abuse, for which we assume a $500 per diem. We also assumed a 35% discount from national average charges for hospital outpatient services.

APPENDIX II

Summary of Plan Provisions of "Typical" Benefit Plans

A. The 1997 Federal Employee Health Benefits Plan - Blue Cross Blue Shield, Standard Option, assuming PPO/Preferred Physicians

Service Category	Benefit Provision
Hospital Inpatient	
Medical	100% paid by plan
Mental Health	\$150 copay per day, up to 100 days per year
Substance Abuse	\$150 copay per day, up to 100 days per year
Outpatient	
Medical	\$10 copay per visit
Mental Health	40% copay per visit, up to 25 visits per year
Substance Abuse	40% copay per visit, up to 25 visits per year
Prescription Drug	20% copay per script

NAMI

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Andrew Sperling
Director of Public Policy

Mental Health

July 30, 1997

President Bill Clinton
The White House
Washington, D.C. 20500

Dear Mr. President:

On behalf of the 168,000 members of the National Alliance for the Mentally Ill (NAMI) and its 1,140 affiliates and chapters, I am writing to thank you for your efforts to ensure that a meaningful standard for inclusion of mental illness benefits was made a part of the new childrens' health program in the balanced budget bill.

While NAMI had pushed hard for the original childrens' mental illness parity amendment that was added to the bill last month by Senators Paul Wellstone and Pete Domenici, we understand that adoption of any federal standard for the program's benefit package was very controversial and difficult to achieve. The spirited (and in NAMI's opinion shortsighted) advocacy by some of the nation's governors and a few congressional leaders against any benefit package requirement was a significant obstacle. Overcoming this opposition and ensuring that there will be accountability in how states spend these federal funds is a tremendous accomplishment for you and advocates for equitable mental illness coverage in Congress.

We are deeply grateful for your leadership and perseverance on this important issue for NAMI. As the nation's largest consumer and family organization, we appreciate what you have done to ensure that uninsured children with severe mental illness will get access to health care treatment that truly meets their needs. While the result is not full parity for mental illness treatment, inclusion in the benefits package does help bring us another step closer to acceptance of severe mental illness into the mainstream of our nation's health care system.

Again, thank you for your efforts on behalf of all children with severe mental illnesses and their families.

Sincerely,

Laurie Flynn

Laurie Flynn
Executive Director

cc: Tipper Gore, Office of the Vice President
Chris Jennings, Special Assistant to the President for Health Policy