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The Impact of Nearly Universal Insurance Coverage on Health Care Utilization and Health: Evidence from Medicare

The Impact of Nearly Universal Insurance Coverage on Health Care Utilization: Evidence from Medicare

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One-fifth of nonelderly adults in the United States lacked health insurance coverage in 2005. Most of these were from lower-income families, and nearly one-half were African American or Hispanic (Carmen DeNavas-Walt, Bernadette Proctor, and Cheryl Hill Lee 2005). Many analysts have argued that unequal insurance coverage contributes to disparities in health care utilization and health outcomes across socioeconomic groups. Even among the insured there are differences in copayments, deductibles, and other features that affect service use. Nevertheless, credible evidence that better insurance causes better health outcomes is limited (M. E. Brown, A. B. Bineman, and N. Lurie 1998; Helen Levy and David Meltzer 2001). Both the supply and demand for insurance depend on health status, confounding observational comparisons between people with different insurance characteristics.

In contrast to the heterogeneity among the nonelderly, fewer than 1 percent of the elderly population are uninsured, and most have fee-for-service Medicare coverage. The transition occurs abruptly at age 65, the threshold for Medicare eligibility. Building on this fact, in this paper we use a regression-discontinuity framework to compare health-related outcomes among people just before and just after the age of 65. Our analysis extends existing research on the effects of the age 65 threshold (Frank R. Lichtenberg, 2002; William H. Dow 2003; Sandra Decker and Carol Rapaport 2002a; Decker 2002; Decker and Rapaport 2000b; J. Michael McWilliams et al. 2003) in two main ways. First, we examine a wider range of outcomes. We use survey data from the National Health Interview Survey (NHIS) to analyze changes in self-reported access to care, and in the number of recent doctor visits and hospital stays. We supplement these data with hospital discharge records from California, Florida, and New York, which allow us to measure changes in hospital admissions for specific conditions and procedures, and by hospital type. Second, we focus on the differential effects of Medicare eligibility on different subgroups, and use the pattern of intergroup differences to assess whether these impacts arise through changes in insurance coverage, insurance generosity, or other channels. We also quantify the extent to which the onset of Medicare eligibility reduces or increases disparities in use of different types of services.

Our main finding is that Medicare eligibility causes a sharp increase in the use of health care services, with a pattern of gains across groups that varies by the type of service. For relatively low-cost services, such as routine doctor visits, the onset of Medicare eligibility leads to increases in utilization that are concentrated among groups with the lowest rates of insurance coverage

for persons under the age of 65. For relatively high-cost procedures—including hospitalization for procedures like bypass surgery and hip and knee replacement—the gains are concentrated among groups that are more likely to have supplementary insurance coverage after 65. These patterns, coupled with evidence of a redistribution of patients across hospital ownership categories once Medicare is available, suggest that the distribution of gains in use of health services is driven by an *interaction* between supply-side incentives and shifts in insurance characteristics for different socioeconomic groups.

I. Measuring the Causal Effect of Health Insurance

We work with a simple reduced-form model of the causal effects of health insurance status:

$$(1) \qquad y_{ija} = X_{ija}\alpha + f_j(a; \beta) + \sum_k C^k_{ija}\delta^k + u_{ija},$$

where y_{ija} is a measure of health care use for individual i in socioeconomic group j at age a, u_{ija} is an unobserved error component, X_{ija} is a set of covariates (e.g., gender and region), $f_j(a; \beta)$ is a smooth function representing the age profile of outcome y for group j, and C^k_{ija} ($k = 1, 2, \ldots, K$) are characteristics of the insurance coverage held by the individual. These can include a simple coverage indicator as well as variables summarizing other features such as copayment rates or the presence of gatekeeper restrictions.

A fundamental problem for the estimation of equation (1) is that insurance coverage is endogenous. The age threshold for Medicare eligibility at 65 provides a credible source of exogenous variation in insurance status. To illustrate this claim, Figure 1 shows the age profiles of health insurance coverage estimated with data from the pooled 1999–2003 NHIS, where age is measured in quarters (the sample is described below and in more detail in the online Appendix, available at http://www.aeaweb.org/articles.php?doi=10.1257/aer.98.5.2242). Overall coverage rates (plotted with open diamonds) rise from 85 to 96 percent at age 65. Even more striking is the impact of Medicare eligibility on differences across socioeconomic groups. Prior to age 65, less educated minorities (blacks, Asians, and Hispanics with under 12 years of education) have 25 percentage points lower coverage rates than highly educated whites. After 65 the gap falls to 10 points or less.

Figure 1 also shows the fractions of individuals with two or more insurance policies. Before age 65, multiple coverage is relatively rare. The incidence rises at 65, with a bigger gain for highly educated whites, reflecting a greater likelihood of enrollment in supplemental "Medigap" policies (see Section II below). Thus, Medicare eligibility is associated with a narrowing of disparities in the probability of any coverage, but a widening of disparities in at least one indicator of the generosity of coverage.

To proceed, suppose that a person's health insurance coverage can be summarized by two indicator variables: C^1_{ija} indicating any coverage, and C^2_{ija} indicating a relatively generous insurance package (i.e., low copayments and few gatekeeper restrictions). Consider linear probability models for the events of any coverage and generous coverage of the form

$$(2a) \qquad C^1_{ija} = X_{ija}\beta^1_j + g^1_j(a) + D_a\pi^1_j + v^1_{ija},$$

$$(2b) \qquad C^2_{ija} = X_{ija}\beta^2_j + g^2_j(a) + D_a\pi^2_j + v^2_{ija},$$

where β^1_j and β^2_j are group-specific coefficients, $g^1_j(a)$ and $g^2_j(a)$ are smooth age profiles for group j, and D_a denotes an indicator for being age 65 or older. Combining equations (2a) and (2b) with equation (1), the reduced-form model for outcome y is

$$(3) \qquad y_{ija} = X_{ija}(\alpha_j + \beta^1_j\delta^1_j + \beta^2_j\delta^2_j) + h_j(a) + D_a\pi^y_j + v^y_{ija},$$
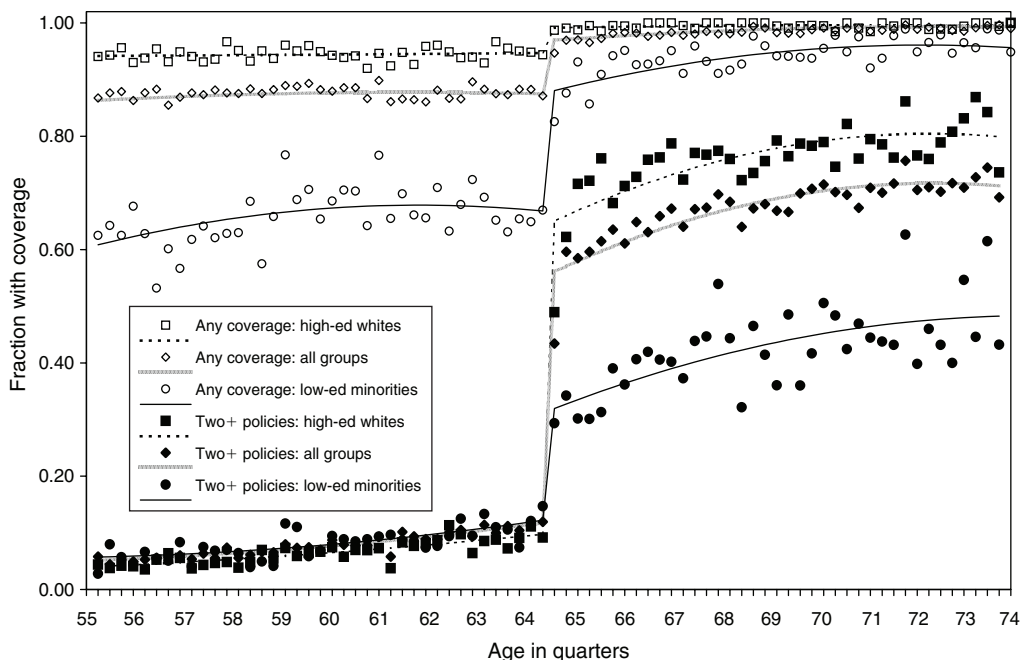
FIGURE 1. COVERAGE BY ANY INSURANCE AND BY TWO OR MORE POLICIES, BY AGE AND DEMOGRAPHIC GROUP

where $h_j(a) = f_j(a) + \delta^1 g_j^1(a) + \delta^2 g_j^2(a)$ represents the reduced-form age profile for group j, $\pi_j^y = \pi_j^1 \delta^1 + \pi_j^2 \delta^2$, and $v_{ija}^y = u_{ija} + v_{ija}^1 \delta^1 + v_{ija}^2 \delta^2$ is an error term. Assuming that the profiles $f_j(a)$, $g_j^1(a)$, and $g_j^2(a)$ are all continuous at age 65, any discontinuity in y can be attributed to discontinuities in insurance. The magnitude depends on the size of the insurance changes at 65 (π_j^1 and π_j^2), and on the associated causal effects (δ^1 and δ^2).

For some basic health care services—for example, routine doctor visits—it is arguable that only the presence of insurance matters. In this case, the implied discontinuity in y at age 65 for group j will be proportional to the change in insurance coverage experienced by the group. For more expensive or elective services, the generosity of coverage may also matter, if patients are unwilling to cover the required copayment or if managed care programs will not cover the service. This creates a potential identification problem in interpreting the discontinuity in y for any one group. Since π_j^y is a linear combination of the discontinuities in coverage and generosity, δ^1 and δ^2 can be estimated by a regression across groups:

$$(4) \qquad \pi_j^y = \delta^0 + \delta^1 \pi_j^1 + \delta^2 \pi_j^2 + e_j,$$

where e_j is an error term reflecting a combination of the sampling errors in π_j^y, π_j^1, and π_j^2.

This framework can be extended to include additional dimensions of insurance coverage. In practice, however, a key limitation is the lack of information on the insurance packages held by different individuals. In the absence of more complete data, we use the presence of at least two forms of coverage as an indicator of "generous" coverage. We also explore a simple measure of gatekeeper limitations, based on whether an individual's primary insurer is a managed care provider.

In our empirical analysis, we fit regression discontinuity (RD) models like (2a), (2b), and (3) by demographic subgroup to individual data using OLS estimators. We then combine the estimates

across groups in the final section of the paper to estimate models like (4). For our main results we follow John E. DiNardo and Lee (2004) and assume the age profiles in equations (1), (2a), and (2b) are continuous polynomials with potential discontinuities in the derivatives at age 65. We have also fit many of the models using local linear regression (as suggested by Jinyong Hahn, Petra Todd, and Wilber van der Klaauw 2001). As discussed below, the estimated discontinuities are generally quite robust, reflecting the relative smoothness of the age profiles in insurance features and health outcomes.

We use data from two main sources: the 1992–2003 NHIS and 1992–2003 hospital discharge records for California, Florida, and New York. The NHIS reports respondents' birth year and birth month, and the calendar quarter of the interview. We use these data to construct an estimate of age in quarters. We adopt the convention that a person who reaches his sixty-fifth birthday in the interview quarter is age 65 and 0 quarters. Assuming a uniform distribution of interview dates, about one-half of these people will be 0–6 weeks younger than 65, and one-half will be 0–6 weeks older. We limit our analysis to people who are over 55 and under 75: the final sample size is 160,821, although some outcomes (e.g., detailed insurance characteristics) are available only in later years. Sample counts and descriptive statistics by age are reported in the online Appendix. The discharge files represent a complete census of discharges from all hospitals in the three states (except federally regulated institutions). The data files include information on age in months at the time of admission. For our analysis we drop records for people admitted as transfers from other institutions, and limit attention to people between 60 and 70 years of age at admission. The sample sizes are 4,017,325 for California; 2,793,547 for Florida; and 3,121,721 for New York.

II. Changes in Insurance Coverage at Age 65

Medicare is available to people who are at least 65 and have worked 40 quarters or more in covered employment (or have a spouse who did).[1] Coverage is also available to younger people with severe kidney disease and recipients of Social Security Disability Insurance (DI). Eligible individuals can obtain Medicare hospital insurance (Part A) free of charge, and medical insurance (Part B) for a modest monthly premium. Individuals receive notice of their impending eligibility for Medicare shortly before their sixty-fifth birthday, and are informed that they have to enroll in the program and choose whether to accept Part B coverage. Coverage begins on the first day of the month in which they turn 65.

Table 1 shows the effects of reaching age 65 on five insurance-related variables: the probability of Medicare coverage, the probability of any health insurance coverage, the probability of private coverage, the probability of two or more forms of coverage, and the probability that an individual's primary health insurance is a managed care program. As in Figure 1, the data are drawn from the 1999–2003 NHIS. For each characteristic we show the incidence rate at ages 63–64, and the change at age 65, based on a version of equations (2a/2b) that includes a quadratic in age, fully interacted with a post-65 dummy.[2] Alternative specifications—including a parametric model fit to a narrower age window (ages 63–67) and a local linear regression specification using a rule-of-thumb bandwidth selection procedure—are shown in the online Appendix (see Appendix Table 1a), and yield very similar estimates of the change at age 65.

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[1] Individuals who do not qualify may still enroll in Medicare at age 65 by paying monthly premiums for both Part A and Part B coverage. This option is limited to US citizens and legal aliens with at least five years of residency in the United States.

[2] The models also include controls for gender, education, race/ethnicity, region, and sample year.

TABLE 1—INSURANCE CHARACTERISTICS JUST BEFORE AGE 65 AND ESTIMATED DISCONTINUITIES AT AGE 65

	On Medicare		Any insurance		Private coverage		2+ Forms coverage		Managed care	
	Age 63–4 (1)	RD at 65 (2)	Age 63–4 (3)	RD at 65 (4)	Age 63–4 (5)	RD at 65 (6)	Age 63–4 (7)	RD at 65 (8)	Age 63–4 (9)	RD at 65 (10)
Overall sample	12.3	59.7 (4.1)	87.9	9.5 (0.6)	71.8	−2.9 (1.1)	10.8	44.1 (2.8)	59.4	−28.4 (2.1)
Classified by ethnicity and education:										
White non-Hispanic:										
High school dropout	21.1	58.5 (4.6)	84.1	13.0 (2.7)	63.5	−6.2 (3.3)	15.0	44.5 (4.0)	48.1	−25.0 (4.5)
High school graduate	11.4	64.7 (5.0)	92.0	7.6 (0.7)	80.5	−1.9 (1.6)	10.1	51.8 (3.8)	58.9	−30.3 (2.6)
At least some college	6.1	68.4 (4.7)	94.6	4.4 (0.5)	85.6	−2.3 (1.8)	8.8	55.1 (4.0)	69.1	−40.1 (2.6)
Minority:										
High school dropout	19.5	44.5 (3.1)	66.8	21.5 (2.1)	33.2	−1.2 (2.5)	11.4	19.4 (1.9)	39.1	−8.3 (3.1)
High school graduate	16.7	44.6 (4.7)	85.2	8.9 (2.8)	60.9	−5.8 (5.1)	13.6	23.4 (4.8)	54.2	−15.4 (3.5)
At least some college	10.3	52.1 (4.9)	89.1	5.8 (2.0)	73.3	−5.4 (4.3)	11.1	38.4 (3.8)	66.2	−22.3 (7.2)
Classified by ethnicity only:										
White non-Hispanic (all)	10.8	65.2 (4.6)	91.8	7.3 (0.5)	79.7	−2.8 (1.4)	10.4	51.9 (3.5)	61.9	−33.6 (2.3)
Black non-Hispanic (all)	17.9	48.5 (3.6)	84.6	11.9 (2.0)	57.1	−4.2 (2.8)	13.4	27.8 (3.7)	48.2	−13.5 (3.7)
Hispanic (all)	16.0	44.4 (3.7)	70.0	17.3 (3.0)	42.5	−2.0 (1.7)	10.8	21.7 (2.1)	52.9	−12.1 (3.7)

Note: Entries in odd-numbered columns are percentages of age 63-64-year-olds in group with insurance characteristic shown in column heading. Entries in even-numbered columns are estimated regression discontinuities at age 65, from models that include quadratic control for age, fully interacted with dummy for age 65 or older. Other controls include indicators for gender, race/ethnicity, education, region, and sample year. Estimates are based on linear probability models fit to pooled samples of 1999–2003 NHIS.

Medicare coverage rises by 60 percentage points at age 65, from a base level of 12 percent among 63-64-year-olds. Consistent with DI enrollment patterns (David H. Autor and Mark G. Duggan 2003), Medicare enrollment prior to 65 is higher for minorities and people with below-average schooling, and these groups experience relatively smaller gains at age 65 (see rows 2–7). The pattern is reversed for the probability of any insurance coverage (columns 3 and 4): groups with lower insurance coverage rates prior to 65 experience larger gains at age 65. There is still some disparity in insurance coverage after 65, but the 28-point gap between more educated whites and less educated minorities narrows to about 10 points. Similarly, as shown in rows 8–10, the 21-point gap in coverage between whites and Hispanics before age 65 closes to only 12 points after. Thus, the onset of Medicare eligibility dramatically reduces disparities in insurance coverage.

Columns 5 and 6 present information on the prevalence of private insurance coverage (i.e., employer-provided or purchased coverage). Prior to age 65 private coverage rates range from 33 percent for less educated minorities to 86 percent for better educated whites. The RD estimates in column 6 show that these differences are hardly affected by the onset of Medicare eligibility. This stability reflects the fact that most people who hold private coverage before 65 transition to a combination of Medicare and supplemental coverage, either through an employer-provided plan or an individually purchased Medigap policy.[3] Columns 7 and 8 of Table 1 analyze the age patterns of multiple coverage (i.e., reporting two or more policies). Prior to age 65, the rate

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[3] Across the six groups in rows 2–7 of Table 1, for example, the correlation between the private coverage rate at ages 63–64 shown in column 5 and the fraction of 65-66-year-olds with private supplemental Medicare coverage is 0.97.

of multiple coverage is quite low (around 11 percent) and similar across groups. As shown in column 8, the overall rate of dual coverage rises by about 44 percentage points at age 65, with gains close to 60 percentage points for better educated whites, but only on the order of 20 percentage points for less educated minorities. Thus, significant disparities in dual coverage arise after age 65 (see also Figure 1).

Another important dimension of coverage is managed care versus indemnity coverage. The entries in column 9 of Table 1 show that among 63- 64-year-olds with insurance coverage, nearly 60 percent have managed care, with higher rates for whites and better educated groups. At age 65, the overall fraction of people with managed care for their primary insurance drops sharply, with larger declines for groups with higher rates prior to 65, leading to rough convergence in managed care rates across groups. This decline reflects the relatively low enrollment in Medicare managed care: about 85 percent of 65- 66-year-old Medicare recipients are enrolled in fee-for-service Medicare, which offers patients and providers substantial leeway over the use of services.

Overall, there are major changes in health insurance at age 65. Many of those who lacked insurance prior to 65 obtain coverage, equalizing coverage rates across groups. There is also a sharp rise in multiple coverage, particularly among whites and the better educated. Coupled with the shift from managed care to fee-for-service coverage, it appears that the relative "generosity" of insurance coverage among more advantaged groups actually increases with the onset of Medicare eligibility.

A. *Other Changes at Age 65*

Formal identification of an RD model that relates an outcome y (e.g., insurance coverage) to a treatment (Medicare age-eligibility) that depends on age (a) relies on an assumption about the expectation of y conditional on age and treatment status. Let $y(0)$ and $y(1)$ denote the potential outcomes for a given person if he or she was or was not "treated." Note that $y(0)$ is observed only for people under 65, while $y(1)$ is observed only for people 65 or over. The key assumption is that $E[y(0)|a]$ and $E[y(1)|a]$ are both continuous at $a = 65$ (Guido W. Imbens and Thomas Lemieux 2008, Assumption 2.1). In this case the average treatment effect at age 65 is identified as $\lim_{a\downarrow 65} E[y(1)|a] - \lim_{a\uparrow 65} E[y(0)|a]$. Continuity requires that all other factors that might affect the outcome of interest trend smoothly at age 65.

An obvious concern in our context is employment, since 65 is a traditional age of retirement, and any abrupt change in the fraction of people working at 65 could lead to differences in health care utilization if nonworkers have more time to visit doctors.[4] As noted by Lee (2008), a simple test for the potential impact of discontinuities in confounding variables like employment is fitting a model like (3) for the confounding variable and testing for jumps at age 65.

Table 2 presents estimation results for a set of models that test for discontinuities in the age profiles of employment, using data from the 1992–2003 NHIS (with age measured in quarters) and the 1996–2004 March Current Population Surveys (with age measured in years).[5]

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[4] Robin L. Lumsdaine, James H. Stock, and David A. Wise (1995) report evidence of a spike in the retirement hazard rate at age 65. More recent data from the Health and Retirement Study, however, show little or no spike at 65 (Till von Wachter 2002, fig. 3). Moreover, a spike in the retirement *hazard* implies a discontinuity in the second derivative of the employment survivor function, rather than a discontinuity in the employment *rate*.

[5] For the NHIS, we use labor market information for the survey week: employment status is available from 1992 to 2003 but full time status is available only for 1997–2003. For the 1996–2004 March CPS, we use employment, hours of work, and self-reported retirement status as of the survey week, as well as total annual earnings last year. See the online Appendix for information on the CPS sample.

TABLE 2—ESTIMATED DISCONTINUITIES IN EMPLOYMENT MEASURES AT AGE 65

	Data from NHIS		Data from March CPS			
	Employed (1)	Full time (2)	Employed (3)	Hours/Wk (4)	Retired (5)	Earnings (1000s) (6)
Overall sample	0.3 (0.8)	0.8 (1.2)	1.8 (1.1)	1.3 (0.7)	−1.6 (1.7)	−0.2 (0.4)
Classified by ethnicity and education:						
White non-Hispanic:						
High school dropout	1.0 (1.4)	1.1 (1.9)	1.9 (1.1)	1.8 (0.5)	−3.7 (2.4)	0.3 (0.8)
High school graduate	1.7 (1.0)	2.4 (1.8)	2.3 (1.5)	1.4 (0.7)	−2.5 (2.3)	−0.2 (0.5)
At least some college	−1.6 (1.5)	−0.4 (2.0)	0.9 (1.4)	0.9 (1.0)	−0.6 (1.7)	−1.5 (0.8)
Minority:						
High school dropout	2.6 (1.6)	1.5 (2.0)	1.2 (1.3)	0.7 (0.5)	−0.3 (2.4)	0.1 (0.4)
High school graduate	0.0 (3.2)	0.2 (2.6)	6.1 (2.3)	2.5 (1.1)	−5.3 (1.5)	0.7 (0.6)
At least some college	−4.6 (2.6)	−2.6 (3.0)	−1.2 (1.5)	0.1 (0.6)	2.4 (1.8)	0.7 (1.4)
Classified by ethnicity only:						
White non-Hispanic	0.2 (0.9)	0.9 (1.3)	1.6 (1.2)	1.3 (0.7)	−1.7 (2.0)	−0.5 (0.5)
Black non-Hispanic (all)	1.3 (2.6)	0.1 (2.0)	5.7 (1.4)	2.2 (0.5)	−4.6 (0.9)	1.3 (0.4)
Hispanic (all)	−0.5 (2.0)	0.7 (2.0)	−0.4 (1.7)	0.8 (0.9)	−0.7 (1.7)	−1.1 (1.2)
Classified by gender:						
Men	1.7 (1.3)	3.0 (1.6)	2.8 (1.5)	1.7 (0.9)	−1.6 (1.6)	−0.5 (0.6)
Women	−1.0 (1.1)	−1.2 (1.2)	0.9 (1.0)	0.9 (0.6)	−1.4 (2.0)	0.1 (0.3)

Note: Entries are estimated regression discontinuties at age 65, from models that include quadratic controls for age fully interacted with a dummy for age 65 or older. Other controls include indicators for gender, race/ethnicity, education, region, and sample year. Estimates in column 1 are from models fit to pooled samples of 1992–2003 NHIS. Estimates in column 2 are fit on a pooled sample from 1997–2003 NHIS. Estimates in columns 3–6 are from models fit to pooled samples of 1996–2004 March CPS. NHIS samples include people over 54 and under 75. CPS samples include people age 55–75 (inclusive). Standard errors in parentheses are clustered by quarter (columns 1–2) or year (columns 3–6) of age.

Using either data source, the estimated jumps in employment-related outcomes are small in magnitude and statistically insignificant. Figure 3 displays the actual and fitted age profiles of employment for the overall NHIS sample and for highly educated whites and less educated minorities. These profiles all trend relatively smoothly through age 65, though there is some evidence of a discrete drop at age 62, reflecting the large fraction of people who retire as soon as Social Security benefits are available (e.g., John Rust and Christopher Phelan 1997). When we estimate the discontinuity at 65 using a local linear regression procedure or a narrow sample window (see Appendix Table 1a in the online Appendix), the estimated changes are slightly larger (−2.3 percent for the overall sample) and statistically significant, but still small in comparison to the changes in insurance.

One concern is that the smoothness in overall employment trends at 65 may be masking differences between men and women. The bottom two rows of Table 2 present results by
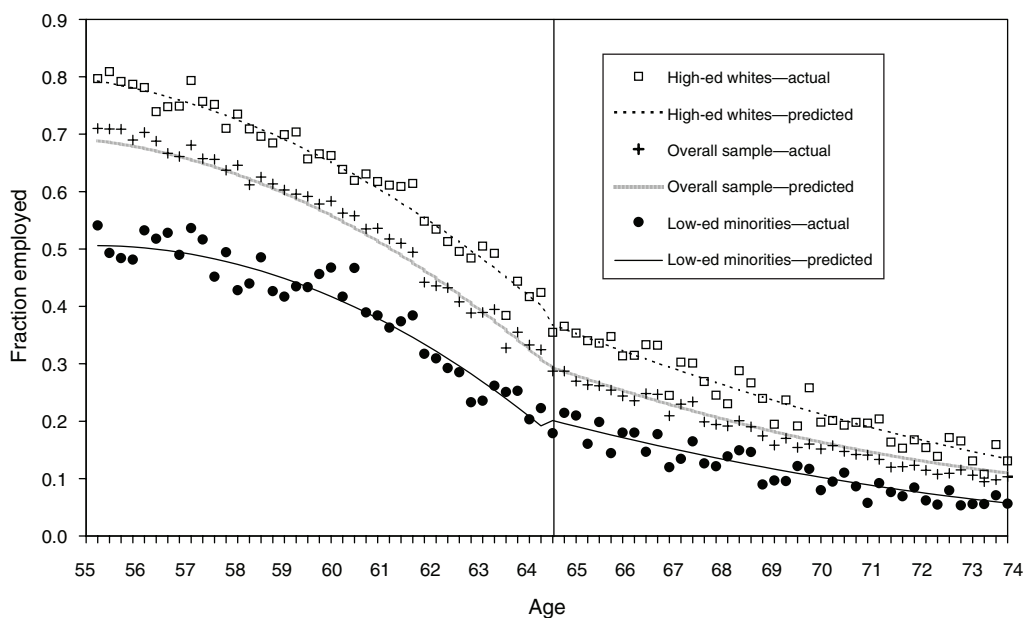
FIGURE 2. EMPLOYMENT RATES BY AGE AND DEMOGRAPHIC GROUP (1992–2003 NHIS)

gender, and show no large discontinuities for either men or women.[6] As an additional check, we used longitudinal data from the 2001 Survey of Income and Program Participation to estimate month-to-month changes in individual employment (see the online Appendix). Consistent with the results here, we found no evidence of a discontinuity in employment at age 65. We also investigated the age profiles of marriage, being classified as poor, and receiving food stamps in the NHIS, as well as residential mobility, marital status, and the incidence of low income in the CPS. As summarized in the online Appendix to this paper, none of these outcomes shows significant discontinuities at age 65 for the overall sample or the subgroups used in Tables 1 and 2. We conclude that employment, family structure, family income, and location, taken as a whole, all trend relatively smoothly at age 65, and are unlikely to confound our analysis of the impact of Medicare eligibility.

III. Changes in Health Care Access and Utilization at Age 65

We now turn to an analysis of the effects of reaching age 65 on access to care and utilization of health care services. Since 1997 the NHIS has asked two questions: (1) "During the past 12 months has medical care been delayed for this person because of worry about the cost?" and (2) "During the past 12 months was there any time when this person needed medical care but did not get it because (this person) could not afford it?" Columns 1 and 3 of Table 2 show the fractions of people ages 63–64 in the pooled 1997–2003 NHIS who responded positively to these two questions. Overall, about 7 percent of the near-elderly reported delaying care, and 5 percent reported not getting care, with relatively higher rates for less educated minorities and for Hispanics. Our RD estimates in columns 2 and 4 imply significant declines at age 65 in both measures of access

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[6] Graphs similar to Figure 2 by gender are available in our online Appendix.

TABLE 3—MEASURES OF ACCESS TO CARE JUST BEFORE 65 AND ESTIMATED DISCONTINUITIES AT 65

| | 1997–2003 NHIS | | | | 1992–2003 NHIS | | | |
| | Delayed care last year | | Did not get care last year | | Saw doctor last year | | Hospital stay last year | |
	Age 63–64 (1)	RD at 65 (2)	Age 63–64 (3)	RD at 65 (4)	Age 63–64 (5)	RD at 65 (6)	Age 63–64 (7)	RD at 65 (8)
Overall sample	7.2	−1.8 (0.4)	4.9	−1.3 (0.3)	84.8	1.3 (0.7)	11.8	1.2 (0.4)
Classified by ethnicity and education:								
White non-Hispanic:								
High school dropout	11.6	−1.5 (1.1)	7.9	−0.2 (1.0)	81.7	3.1 (1.3)	14.4	1.6 (1.3)
High school graduate	7.1	0.3 (2.8)	5.5	−1.3 (2.8)	85.1	−0.4 (1.5)	12.0	0.3 (0.7)
At least some college	6.0	−1.5 (0.4)	3.7	−1.4 (0.3)	87.6	0.0 (1.3)	9.8	2.1 (0.7)
Minority:								
High school dropout	13.6	−5.3 (1.0)	11.7	−4.2 (0.9)	80.2	5.0 (2.2)	14.5	0.0 (1.4)
High school graduate	4.3	−3.8 (3.2)	1.2	1.5 (3.7)	84.8	1.9 (2.7)	11.4	1.8 (1.4)
At least some college	5.4	−0.6 (1.1)	4.8	−0.2 (0.8)	85.0	3.7 (3.9)	9.5	0.7 (2.0)
Classified by ethnicity only:								
White non-Hispanic	6.9	−1.6 (0.4)	4.4	−1.2 (0.3)	85.3	0.6 (0.8)	11.6	1.3 (0.5)
Black non-Hispanic (all)	7.3	−1.9 (1.1)	6.4	−0.3 (1.1)	84.2	3.6 (1.9)	14.4	0.5 (1.1)
Hispanic (all)	11.1	−4.9 (0.8)	9.3	−3.8 (0.7)	79.4	8.2 (0.8)	11.8	1.0 (1.6)

Note: Entries in odd numbered columns are mean of variable in column heading among people ages 63–64. Entries in even numbered columns are estimated regression discontinuities at age 65, from models that include linear control for age interacted with dummy for age 65 or older (columns 2 and 4) or quadratic control for age, interacted with dummy for age 65 and older (columns 6 and 8). Other controls in models include indicators for gender, race/ethnicity, education, region, and sample year. Sample in columns 1–4 is pooled 1997–2003 NHIS. Sample in columns 5–8 is pooled 1992–2003 NHIS. Samples for regression models include people ages 55–75 only. Standard errors (in parentheses) are clustered by quarter of age.

problems, especially for less educated minorities and Hispanics. The onset of Medicare eligibility leads to a fall in cost-related access problems and a narrowing of intergroup disparities in access.[7]

The right-hand columns of Table 3 present results for two key measures of health care utilization: (1) "Did the individual have at least one doctor visit in the past year?" and (2) "Did the individual have one or more overnight hospital stays in the past year?" based on pooled samples of the 1992–2003 NHIS. Inspection of the utilization rates among 63- 64-year-olds shows a well-known fact: less educated and minority groups are *less* likely to have a routine doctor visit than better educated and nonminority groups, but *more* likely to have had a hospital spell. The RD estimates in column 6 suggest that the age 65 threshold is associated with a (modest) increase in routine doctor visits, with relatively larger gains for the groups with lower rates before 65.[8] For example, among the near-elderly there is a 7.4 percentage point gap in the probability of a routine

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[7] Because the questions refer to the previous year, our estimates of the effect of reaching 65 on access problems may be attenuated. Specifically, people who recently turned 65 could have had problems in the past year, but before their birthday. Such attenuation may be reduced if people tend to recall only their most recent experiences.

[8] Lichtenberg (2002) also found a discontinuous rise in physician visits in the National Ambulatory Medical Care Surveys, but did not disaggregate visits by race/ethnic group.

doctor visit in the previous year for better educated whites (87.6 percent) versus less educated minorities (80.2 percent). The relatively large 5.0 percentage point gain for the latter group at 65 closes 5.0/7.4 = 68 percent of the pre-65 disparity.

The RD estimates in column 8 are harder to interpret. Overall, there is a rather large rise in hospitalization rates at 65 (on the order of 10 percent), but the gains are larger for better educated whites than other groups. Indeed, the 2.1 percentage point RD at 65 represents a 20 percent increase in hospitalization for this group. The gains for other groups are smaller, and for blacks in particular are quite small, though somewhat imprecise. In the next section we use 100 percent samples of hospital discharge records from California, Florida, and New York to refine these estimates. These data have the advantages of very large sample sizes, and the ability to compare reasons for hospitalization, which turn out to be helpful in understanding the changes at age 65.

IV. Changes in Hospitalization—Evidence from Discharge Data

In this section we use hospital discharge records from 1992–2002 for people between the ages of 60 and 70 in California, Florida, and New York to examine changes in the number and characteristics of hospital admissions at 65. For some of our analyses we convert the numbers of hospital admissions into rates, using population estimates derived from Census Bureau data as denominators. The advantage of hospitalization *rates* is that they can be compared across groups to evaluate disparities in the pre-65 population. The disadvantage is that the denominators must be interpolated from Census Bureau population estimates, introducing some noise in the age profiles of the hospitalization rates. For our RD models we therefore estimate discontinuities in the log of the number of admissions at age 65. Under the assumption that the underlying population counts trend smoothly, the estimated discontinuities in log admission counts can be interpreted as estimates of the percentage discontinuities in admission rates (see Card, Dobkin, and Maestas (2004) for a formal justification of this approach). [9]

Figure 3 shows the actual and fitted age profiles of hospital admission rates based on our pooled data. The markers in the figure represent actual averages (by month of age) of the number of admissions divided by the estimated population of that age. The lines represent fitted regressions from models that assume a quadratic age profile with a full set of post-65 interactions. Overall admission rates rise steadily prior to age 65, then jump sharply at age 65. The increase appears to be "permanent," with no tendency after age 65 to revert to the previous level, as might occur if the jump in admissions represented only catch-up for deferred care.

Table 4 shows estimated hospital admission rates among 60- 64-year-olds in the three states, and the percentage changes in the numbers of admissions at age 65. The entry in row 1 shows that the jump in overall admissions in Figure 3 corresponds to a 7.6 percent increase, comparable to the national estimate of about 10 percent in Table 3.[10] As shown in online Appendix Table 1b, the corresponding estimate from a local linear regression procedure is 8.0 percent (standard error 0.3 percent). Entries in the other columns show the average admission rates at ages 60–64 (per 10,000 person-years) by race and ethnicity, and the increases in these rates at age 65. Hispanics have the lowest admission rates prior to age 65 but experience a slightly larger gain than whites (9.5 percent). Blacks, in contrast, have much higher admission rates prior to 65 but experience a much smaller gain (4.5 percent). On net, these estimates suggest that both the black-white and the

[9] Note that we are using 11 years of discharge records. Thus, the people in a given age group in our samples are actually drawn from 11 different age cohorts, smoothing any differences in cohort size.

[10] Lichtenberg (2002) found a similar jump in admissions in the 1972–1992 National Hospital Discharge Survey; however, unlike Lichtenberg, our more recent data (1992–2002) for California, New York, and Florida do not suggest that the jump at 65 is the result of postponement of hospitalizations in the prior two years.
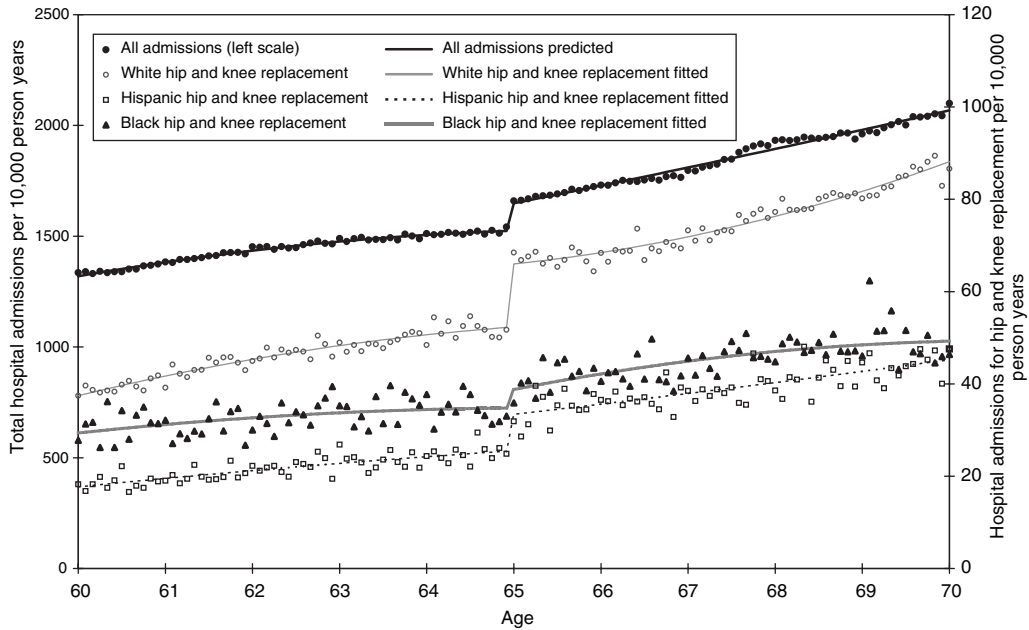
FIGURE 3. HOSPITAL ADMISSION RATES BY RACE/ETHNICITY

Hispanic-white differences in admission rates narrow at age 65, as whites gain relative to blacks and Hispanics gain relative to whites.

For reference, the bottom row of Table 4 shows insurance coverage rates among 60- 64-year-olds in the three states, along with estimated jumps in insurance coverage at 65.[11] Coverage rates in the three states are below the national average prior to 65, but rise by more (15 percent versus a national average of about 10 percent). Consistent with the national data in Table 1, the gains in insurance coverage in the three states are largest for Hispanics (20.3 versus 17.3 percent nationally), a little smaller for blacks (17.6 versus 11.9 percent nationally) and smallest for whites (12.7 versus 7.3 percent nationally).

A key advantage of our hospital data is that we can break down admissions by route into the hospital, and by admission diagnosis and primary procedure. A comparison of rows 2 and 3 in Table 4 shows that most of the jump in admissions at age 65 is driven by non–emergency room admissions, although for each race/ethnic group there is also some increase in ER admissions.[12] Further insights can be gleaned from the admissions patterns across diagnoses. The most common admission diagnosis for near-elderly patients is chronic ischemic heart disease (IHD), which is often treated by coronary artery bypass surgery. There are substantial disparities in IHD

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[11] These data are drawn from the 1996–2004 CPS data for California, New York, and Florida. Given the small sample sizes and the coarseness of the age measure in the CPS, we estimated the insurance RDs assuming a linear age profile but allowing a different slope before and after 65.

[12] ER admissions include extremely urgent cases (which one might expect to be unresponsive to insurance status) as well as patients who have presented at the ER without being referred by a physician. Some analysts have argued that provision of health insurance would reduce ER use and shift patients to outpatient care. Nevertheless, our results are consistent with the RAND Health Insurance Experiment, which found ER use as responsive to copayment rates as use of outpatient care (Joseph P. Newhouse 1993).

TABLE 4—HOSPITAL ADMISSIONS AND INSURANCE COVERAGE AT AGE 65: CALIFORNIA, FLORIDA, AND NEW YORK

	All		Whites		Hispanics		Blacks	
	Rate age 60–64	RD at 65	Rate age 60–64	RD at 65	Rate age 60–64	RD at 65	Rate age 60–64	RD at 65
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Hospital admissions								
All admissions	1,443	7.57	1,407	7.74	1,262	9.47	2,008	4.39
		(0.29)		(0.33)		(0.55)		(0.71)
By route into hospital								
ER admission	761	3.30	688	3.70	774	2.63	1,313	1.93
		(0.39)		(0.40)		(0.92)		(0.95)
Non–ER admission	682	12.16	718	11.51	488	19.89	695	8.92
		(0.46)		(0.49)		(1.05)		(1.04)
By admission diagnosis								
Chronic ischemic heart disease	83	11.58	90	11.05	59	18.45	66	8.29
		(0.96)		(1.16)		(2.45)		(2.78)
AMI	48	4.41	50	5.31	38	3.90	45	−3.43
		(1.43)		(1.65)		(3.33)		(4.78)
Heart failure	56	0.44	45	2.33	62	−4.85	130	−1.47
		(1.11)		(1.24)		(2.63)		(2.43)
Chronic bronchitis	34	7.50	36	6.50	19	9.76	38	13.05
		(1.51)		(1.52)		(5.58)		(4.43)
Osteoarthrosis	34	26.97	38	27.16	18	29.27	27	22.08
		(1.39)		(1.64)		(5.05)		(4.01)
Pneumonia	34	2.44	32	2.05	30	3.39	51	3.81
		(1.42)		(1.74)		(4.34)		(3.21)
By primary procedure								
None	419	5.70	400	5.73	388	7.23	614	3.86
		(0.33)		(0.40)		(1.23)		(1.25)
Diagnostic procedures on heart	51	9.18	53	8.17	40	16.78	58	8.76
		(1.03)		(1.20)		(3.21)		(3.32)
Removal of coronary artery obstruction	38	10.67	43	10.49	23	18.77	22	0.49
		(1.46)		(1.60)		(3.94)		(5.33)
Bypass anastomosis of heart	26	15.91	29	16.17	17	18.97	13	5.15
		(1.39)		(1.44)		(5.62)		(6.09)
Joint replacement lower extremity	41	22.69	46	23.16	22	26.40	33	12.14
		(1.47)		(1.60)		(4.69)		(4.20)
Diagnostic procedure on small intestine	35	7.35	31	6.60	37	13.07	58	4.09
		(1.27)		(1.47)		(3.27)		(3.13)
Cholecystectomy (gall bladder removal)	26	17.93	26	16.00	29	29.25	18	12.27
		(2.10)		(1.84)		(5.11)		(7.50)
Insurance coverage								
Probability of coverage (March CPS data)	82.7	15.0	86.7	12.7	69.1	20.3	79.0	17.6
		(0.8)		(0.8)		(2.7)		(2.7)

Notes: Insurance estimates are based on pooled March CPS 1996–2004 data for California, Florida, and New York. Entries in top row columns 1, 3, 5, 7 are fractions of 60- and 64-year-olds with insurance coverage. Entries in top row columns 2, 4, 6, 8 represent regression discontinuity estimates (×100) of the increase in coverage at age 65 from a model with a quadratic in age, fully interacted with a post-65 dummy. Entries in lower rows columns 1, 3, 5, 7 are hospital admission rates (per 10,000 person years for 60- to 64-year-olds) for California, Florida, and New York 1992–2002. Entries in lower rows columns 2, 4, 6, 8 are regression discontinuity estimates (×100) of the increase in the log of the number of admissions at age 65, from models with a quadratic in age, fully interacted with a post-65 dummy. Standard errors are in parentheses.

admission rates prior to age 65, with a black-white relative admission rate of only 72 percent (see row 4). Overall admissions for this diagnosis rise by 12 percent at age 65, with the largest proportional rise for Hispanics (who have the lowest base rate) and the smallest gain for blacks. The onset of Medicare eligibility therefore leads to a surprising *widening* of the black-white disparity in admission rates for IHD, but a slight narrowing of the Hispanic-white disparity.

The second most common admission diagnosis is acute myocardial infarction (AMI or heart attack). AMI admission rates among 60- 64-year-olds are fairly similar for whites and blacks,

but lower for Hispanics. There is a small rise in total AMI admissions at age 65 (4.4 percent), although estimates for the different race/ethnic groups are not precise enough to rule out equality across the groups.

The third most common admission diagnosis is heart failure (HF), a progressive disorder caused by deterioration of the functional capacity of the heart. Unlike IHD, there is no surgical treatment for HF: the main treatment is medication. HF is particularly prevalent among African Americans, and consistent with this fact, the data show that 60- 64-year-old blacks have a 280 percent higher admission rate for HF than whites. Interestingly, there is no evidence of a jump at age 65 for HF admissions for any group.

We can also classify admissions by the primary procedure (rows 11–16). The leading primary procedure is diagnostic procedures on the heart (such as cardiac catheterization, often performed for people admitted with AMI). Admissions for this group of procedures rise by about 9 percent at age 65, with the largest rise for Hispanics (16.8 percent), leading to a narrowing in Hispanic-white disparities. Blacks have somewhat higher admission rates for this procedure than whites prior to 65, but show a similar proportional increase at 65.

Admission rates for the second most common group of procedures—removal of a coronary artery obstruction (including angioplasty and related procedures)—rise by around 11 percent. Again, the rise is larger for Hispanics than for whites, leading to a narrowing of the Hispanic-white disparity. The gains for blacks, by comparison, are small and statistically insignificant. Black admission rates for these interventions prior to 65 are notably below those for whites, so the changes associated with Medicare eligibility further increase an already large disparity. A similar pattern emerges for admissions for bypass surgeries, which rise by 19.0 percent for Hispanics, 16.2 percent for whites, but by only a statistically insignificant 5 percent for blacks, again leading to a widening of the black-white disparity in admission rates.

The fourth most common group of admission diagnoses is osteoarthrosis (degenerative joint disease), which has substantial overlap with admissions for hip and knee replacement surgery. The impact of Medicare on the age profile of hip/knee surgery admissions is revealing because on the one hand these procedures are readily deferred, and on the other, they are relatively expensive but "routine" interventions that are thought to have a beneficial effect on quality of life. Prior to 65, whites have a much higher admission rate for hip and knee replacements than blacks or Hispanics.[13] At 65, whites experience a large (23 percent) increase in admissions for these procedures. The proportional gain for Hispanics is about the same, but given their much lower base rate, the Hispanic-white disparity actually increases at age 65. Blacks show a much smaller gain than whites or Hispanics, and coupled with their lower rates pre-65, the net effect is a substantial widening in the black-white disparity in hip/knee replacement surgeries.

These conclusions are visually confirmed in Figure 3. Close inspection of the age profile for whites in Figure 3 also reveals a drop-off in procedure rates just prior to 65, coupled with a temporary surge shortly after 65. This pattern suggest that some people who are close to 65 delay knee and hip replacements until they become eligible for Medicare. A recent panel convened by the National Institutes of Health concluded that hip and knee replacement surgeries are under-performed in the United States. If true, this implies that the rise in admissions for these procedures at age 65 may be due to stringent limitations in the insurance coverage available prior to age 65, rather than to excessive generosity of Medicare.

Looking across the patterns in Table 4 by diagnosis and procedure, an interesting contrast emerges between conditions that typically are treated with medication or bed rest (HF, bronchitis, and pneumonia), and those that are treated with specific procedures (IHD, osteoarthrosis,

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[13] Racial disparities in the rates of hip and knee replacement surgery among Medicare recipients have been well documented in the medical literature (e.g., see Mary E. Charleson and John P. Allegrante 2000).
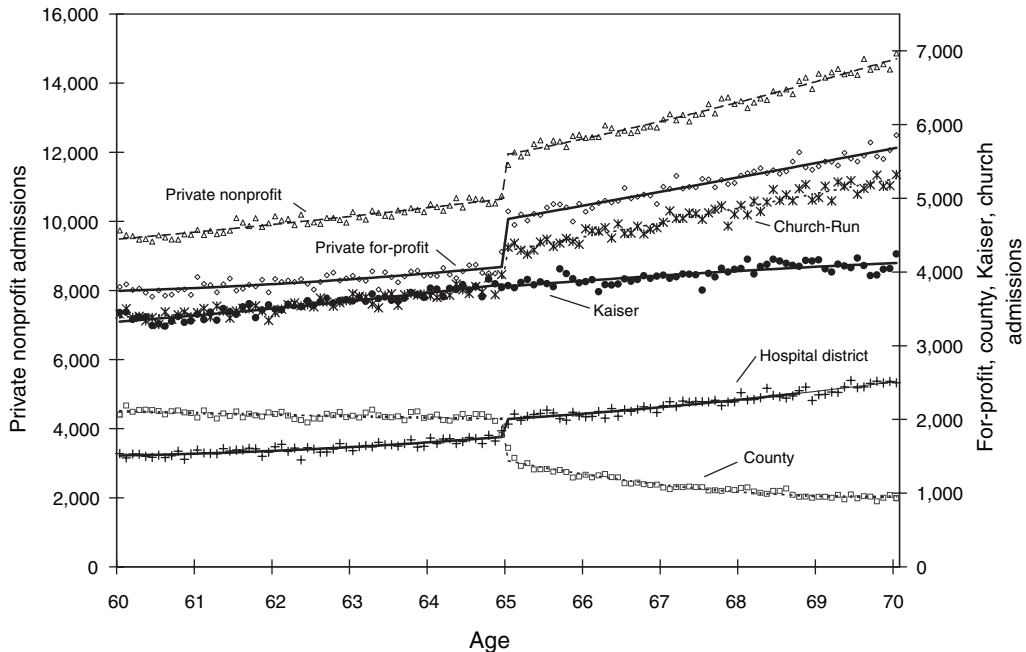
FIGURE 4. HOSPITAL ADMISSION IN CALIFORNIA BY OWNERSHIP TYPE (1992–2002)

gall bladder removal). The first group of admissions tends to increase only slightly at age 65, with similar gains across groups. The second group rises by more, with a pattern that often widens existing racial/ethnic disparities—especially between blacks and whites. This contrast suggests there is an interaction between the availability and generosity of insurance, on one hand, and the existence of specific surgical procedures, on the other, that lead to differences in rises in hospital admissions once Medicare becomes available. In this light, it is interesting to note that the increase in admissions for people who receive no procedures is below the overall growth in admissions (5.7 percent versus an overall average of 7.6 percent).

Further evidence that supply-side reactions to the changes in insurance status at age 65 play a role is presented in Figure 4, which plots the age profiles of hospital admissions by hospital ownership type in California.[14] Private nonprofit hospitals (the largest category) and private for-profits (the second largest category) experience relatively large increases in admissions at age 65. By comparison, hospitals owned by Kaiser Permanente (a large and long-established HMO) show little change in admissions at age 65, and county hospitals experience a sharp decline. These patterns point to two important conclusions. First, the jump at age 65 in overall hospital admissions masks a significant redistribution of caseloads across hospitals. Once Medicare is available, some patients no longer have to use county hospitals and choose a private alternative. Second, the lack of any discontinuity at the Kaiser hospitals suggests that changes in managed care status may have some role in explaining the rise in admissions in the hospital system as a whole. In particular, Kaiser patients remain under a similar managed care regime before and after 65, and these patients show no rise in hospitalization at 65, whereas other patients appear to

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[14] Hospital ownership data are available in our California files only.

be entering the hospital more frequently after 65 and at the same time switching between hospitals. Moreover, physicians at Kaiser are paid on a salary basis and face no particular incentives to seek out Medicare patients for high-cost procedures.

V. Summary of Patterns across Groups

To summarize our findings, we use the framework of equation (4) to relate changes in insurance characteristics at 65 to the changes in health related outcomes. The entries in column 1 of Table 5 represent estimates of the coefficients δ^1 or δ^2, obtained by regressing the estimated discontinuities in the outcome variable indicated in the row heading on the discontinuities in insurance coverage (panel A) or insurance "generosity" (measured by the incidence of multiple coverage, panel B) across six ethnicity-education groups (rows 1–4) or nine state-ethnicity groups (rows 5–8 and 9–12).[15] Column 2 reports the corresponding *R*-squared coefficients. These should be close to one if the change in the outcome variable at age 65 is driven by the measured change in insurance status. Columns 3–5 summarize the pre-65 disparities, while columns 6–8 show the predicted changes in the disparities at 65, computed by multiplying the estimate of δ^1 or δ^2 by the difference in the jumps in coverage or generosity for the disparate groups (from column 4 of Table 1).

Looking first at the measures of access to care in rows 1–3, it appears that the changes across groups at age 65 are closely related to the corresponding changes in insurance coverage. Differential increases in coverage at 65 are estimated to close 25–40 percent of the intergroup disparity in delaying or not getting care, and 74 percent of the gap in the likelihood of a regular doctor visit. This contrasts with the estimated changes in the probability of a hospital stay (row 4), which are slightly negatively related to the increases in insurance coverage. We also fit versions of equation (4) that related these outcomes to the discontinuities in multiple insurance coverage and managed care, but these had less explanatory power.

In contrast to access to care, the entries in rows 5–8 show no evidence of a link between insurance coverage and hospital admissions. If anything, there appears to be a negative relationship between increases in insurance coverage and the size of the RD in admissions for bypasses and hip/knee replacements. The results in panel B yield a similar conclusion for the link between multiple coverage and overall admissions, or admissions for diagnostic heart procedures. For bypass surgery and hip/knee replacement surgery, however, there is more consistent evidence of a link. The *R*-squared coefficient is particularly high for hip and knee surgery, suggesting that the widening disparities in admissions for hip and knee replacements are attributable to the fact that whites are more likely to obtain supplemental coverage after 65 than blacks or Hispanics.

VI. Conclusions

In this paper we use the discrete changes generated by the rules of the Medicare program to identify the impact of health insurance on access to care and utilization. The Medicare eligibility threshold at age 65 is associated with an increase in overall insurance coverage and a narrowing of coverage disparities across different subgroups. There is also an increase in the incidence of multiple coverage and a reduction in managed care, concentrated among higher educated and

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[15] We use weighted least squares, weighting each observation by the inverse sampling variance of the estimated discontinuity in the outcome. Since the estimated discontinuities are independent, this procedure is efficient. The models based on hospital data include state dummies to control for unobserved state-wide factors that affect the responsiveness of health care providers to the onset of Medicare.

TABLE 5—SUMMARY OF EFFECTS OF INSURANCE COVERAGE ON SOCIOECONOMIC DISPARITIES

Outcome	Coefficient on coverage RD (1)	R^2 (2)	Disparities at ages 63–64			Percent change in disparity due to change in coverage at 65		
			Low-ed minority-Hi-ed Whites (3)	Black-White (4)	Hispanic-White (5)	Low-ed minority-Hi-ed Whites (6)	Black-White (7)	Hispanic-White (8)
Panel A: Based on change in insurance coverage at 65								
Delay in care last year	−0.19 (0.06)	0.72	7.6	—	—	−42.8	—	—
No care last year	−0.12 (0.06)	0.39	8.0	—	—	−25.6	—	—
Regular doctor visit last year	0.32 (0.09)	0.77	−7.4	—	—	−73.9	—	—
Hospital stay last year	−0.09 (0.08)	0.26	4.7	—	—	−32.7	—	—
Total hospital admissions	0.06 (0.18)	0.74	—	724	−193	—	1.8	−0.5
Diagnostic procedures of the heart	0.59 (0.31)	0.62	—	9	−25	—	35.6	4.5
Bypass anastomosis of heart	−0.54 (0.93)	0.54	—	−18	−14	—	−4.9	−4.8
Joint replacement of lower extremity	−0.19 (0.64)	0.89	—	−7	−28	—	2.7	−2.3
Panel B: Based on change in incidence of multiple coverage at 65								
Total hospital admissions	0.03 (0.08)	0.74	—	724	−193	—	−0.1	5.8
Diagnostic procedures of the heart	−0.21 (0.14)	0.55	—	9	−25	—	22.3	−18.7
Bypass anastomosis of heart	0.46 (0.34)	0.64	—	−18	−14	—	29.4	37.1
Joint replacement lower of extremity	0.42 (0.21)	0.94	—	−7	−28	—	59.4	25.7

Notes: Each entry in panel A, column 1, is estimated coefficient from regression of RDs in listed health outcome on RDs in insurance coverage over six ethnicity/education groups (rows 1–4) or nine state-ethnicity groups (rows 5–8). All regressions weighted by the inverse sampling variance of the estimated discontinuity in each outcome, and regressions in rows 5–8 include state dummies. Entries in column 2 are corresponding R-squared coefficients from each regression. Entries in columns 3, 4, and 5 are the observed disparities in each health outcome at ages 63–64, and entries in columns 6, 7, and 8 are the percent change in the disparity attributable to the change in insurance coverage based on the coefficient in column 1. Health disparities measured in the NHIS are characterized in terms of low-ed minorities versus hi-ed whites, whereas health disparities measured in the hospital discharge data are characterized in terms of black-white or hispanic-white differences. Panel B is similar to panel A except that the RDs in each health outcome are regressed on the RDs in the incidence of multiple coverage at 65. Panel B regressions are based on data for New York and Florida only (i.e., six state-ethnicity groups).

nonminority groups, as people with insurance prior to 65 enroll in fee-for-service Medicare and supplementary coverage plans.

We find that the onset of Medicare eligibility leads to increases in the use of medical care services, with a pattern of gains across groups that varies with the type of service. Routine doctor visits and access to care increase more for groups that previously lacked coverage, and experience the largest gains in coverage at age 65. Overall hospitalizations increase sharply, but the patterns of gains across groups differ by type of admission. For certain elective hospital admissions, including hip and knee replacements and bypass surgery, the increases are larger for groups that are more likely to have Medicare combined with supplemental coverage after 65. For other conditions like heart failure that are mainly treated by drugs, all groups show very small increases in hospitalization rates at 65. Coupled with evidence of a redistribution of the caseload across hospitals of different ownership types, these patterns suggest that the rise in hospitalization at 65 is driven by an interaction between the increase in insurance "generosity" at 65—specifically for groups who move to fee-for-service Medicare with supplemental coverage—and the existence of profitable treatments (like bypass surgery) for certain diagnoses.

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