Employment, Health Insurance, and Health Care for Vulnerable Populations: Early Retirees, Low-Income Adults, and Racial/Ethnic Minorities

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Abstract

In the first paper, I examine the potential consequences of the recent decline in employer-sponsored retiree health insurance (RHI) offer for the near-elderly population. I find that an RHI offer increases the probability of early retirement by 35 percent. While the results suggest that an RHI offer has little, if any, effect on health in the short term, there is strong evidence that it provides significant protection from high out-of-pocket medical costs. Estimates of the value of retiree health insurance suggest that increasing opportunities for the near-elderly to purchase coverage through the individual market or public programs could significantly reduce the projected increase in uninsurance.

In the second paper, I examine the impact of the introduction of the Medicaid program on labor force participation among single women. Using variation in the timing of Medicaid implementation across states and in eligibility across demographic groups, I find no evidence that women who were eligible for Medicaid decreased their labor supply relative to women who were not. These results add to an emerging consensus in the literature suggesting that public health insurance programs for low-income parents and children may be able to achieve health benefits and improve access to care without substantial indirect costs from labor supply distortions.

Racial/ethnic concordance between patients and physicians may affect health care disparities by reducing discrimination. In the third paper, I investigate the role of concordance on rates of preventive screening and the length of outpatient, primary care visits. I find little evidence that concordance plays an important role in these outcomes. Physician race tends to be a much more important predictor of these outcomes than patient race or concordance, but the direction of the effect varies. The results highlight the importance of measuring the role of concordance separately from patient and physician race. They also suggest that policies aimed at increasing the number of minority physicians need to be combined with other methods to improve the quality of primary care.

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Introduction

Characteristics of the health system in the United States include health insurance that is often tied to employment, unequal access to health care, and variance in the quality of care received by different groups. In this dissertation, I examine the implications of these factors for employment, health care use and health outcomes among vulnerable populations including early retirees, low-income adults, and racial/ethnic minorities. The consequences for policies to improve access and the quality of health care are also discussed.

The proportion of large employers offering retiree health insurance has declined by half in the past 20 years. The first paper examines the potential implications of this change by estimating the effects of retiree health insurance (RHI) offer on a comprehensive set of outcomes in the near-elderly population (ages 47-64) over a 10-year period. I find that an RHI offer increases the probability of early retirement by 35 percent for both men and women. While the results suggest that a retiree health insurance offer has little, if any, effect on health in the short term, there is strong evidence that RHI provides significant protection from high out-of-pocket medical costs. In the top 40 percent of the out-of-pocket spending distribution, those with an offer of retiree coverage spend 21 percent less on average. An RHI offer also moderately increases outpatient medical care utilization. Estimates of the value of retiree health insurance of $3,400 per year for men and $3,100 per year for women suggest that increasing opportunities for the near-elderly to purchase coverage at actuarially-fair prices through the individual market or public programs could significantly reduce the projected increase in uninsurance for this age group.

A thorough assessment of social insurance programs must account for direct costs as well as indirect costs incurred by distortions to individuals' behavior. The second paper examines the impact of the introduction of the Medicaid program on labor supply decisions among single women in the late 1960's and early 1970's. I use a difference-in-differences-in-differences methodology to estimate the effect of Medicaid on eligible women's labor force participation, identifying the effect using variation in the timing of Medicaid implementation across states and in eligibility across demographic groups. I find no evidence that women who were eligible for Medicaid decreased their labor supply relative to women who were not, in contrast to clear theoretical predictions of a negative supply response. The point estimates are positive, suggesting that positive health impacts from health insurance coverage may have contributed to relative increases in labor supply. I find a larger, though still statistically insignificant, labor supply effect among previously married women, compared to their never married counterparts. This is consistent with the hypothesis that previously married women are less likely to be long-term welfare recipients, and therefore were less affected by the negative labor supply incentives associated with Medicaid at implementation. These results add to an emerging consensus in the literature suggesting that public health insurance programs for low-income parents and children may be able to achieve health benefits and improve access to care without substantial indirect costs from labor supply distortions.

Racial/ethnic concordance between patients and physicians may affect health care disparities by reducing discrimination, which operates through the mechanisms of favorable prejudice, modification of negative stereotypes, and increased clinical certainty, trust and compliance. In the third paper, I investigate the role of concordance on rates of preventive screening and the length of time spent with the physician in the outpatient primary care setting. After restricting my sample to physicians who see both white and minority patients and controlling for patient characteristics, physician demographics, and geographic location, I find little evidence that concordance plays an important role in these primary care outcomes. Two exceptions are that racial/ethnic concordance increases cholesterol screening rates by 2-3 times among black and Hispanic men and appears to reduce rates of tobacco cessation counseling among black and Hispanics. Generally speaking, physician race is a much more important predictor of preventive screening and duration of visit than patient race or concordance, but the direction of the effect varies by outcome. The results highlight the methodological importance of measuring the role of concordance separately from patient and physician race. They also suggest that policies aimed at increasing the number of minority physicians need to be combined with physician training and other methods to improve the quality of primary care.

Employer-Sponsored Health Insurance for Early Retirees: Impacts on Retirement, Health, and Health Care*

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Many employers have stopped offering health insurance coverage to retirees in apparent response to rapidly increasing health care costs, the aging of the American population, and recent policy changes such as modified accounting standards governing employer-provided health benefits. Offer rates by large employers have declined from 66 percent to 33 percent between 1988 and 2004.[1] These trends have been well documented, but the implications of this decline in coverage for older Americans, especially for future cohorts of early retirees, are not understood.

Health insurance coverage is of particular importance to older individuals given their relatively poor health, high rates of chronic and acute illnesses, and high levels and variance of medical spending. Group health insurance coverage is valuable to the near-elderly (under age 65) since policies purchased on the individual market tend to be very expensive or even unavailable due to pre-existing health conditions. In addition to affecting retirement, the decline in employer-sponsored retiree coverage may have implications for older Americans' access to health care, financial protection from high medical costs, and their health. This paper explores how retiree health insurance (RHI) affects outcomes on all of these margins and what consequences we can expect from the decline in employer offer rates.

I use individual-level panel data from the Health and Retirement Survey (HRS) from 1992-2002 and examine the effects of a retiree health insurance offer on near-elderly individuals aged 45-64. This large and growing age group included nearly 62

[1] Large employers have 200+ workers. The Kaiser/HRET Employer Health Benefits Annual Survey, 1988-2005.

million individuals and made up 22 percent of the US population in 2000.[2] I compare outcomes over time for individuals who do and do not report an RHI offer to estimate the effects of an RHI offer. The identifying assumption is that retiree health insurance offer is exogenous, conditional on having employer-sponsored health insurance at baseline and controlling for other important covariates. I address the potential endogeneity of RHI offer with several robustness checks, and in general find that these results match the baseline estimates.

In addition to examining the effect of retiree health insurance on early retirement, my analysis focuses on the effects on health outcomes, health insurance coverage, health care utilization, and out-of-pocket medical spending for the near-elderly population. To my knowledge, the effect of RHI on these latter four outcomes has not been studied. The results show large and strongly significant positive effects of retiree health insurance offer on early retirement. I find no significant effects on health, either for the whole sample or among retirees. RHI offer significantly decreases the probability of being uninsured, while increasing the probability of employer-sponsored health insurance coverage and decreasing the probability of public coverage. There is suggestive evidence of an increase in outpatient and prescription drug utilization, but these results are less conclusive.

I find that retiree health insurance reduces out-of-pocket spending on health care among retirees by 21 percent in the top 40 percent of the spending distribution.[3] This highlights the importance of health insurance as protection from high medical

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[2] In 1992, this age group included nearly 50 million individuals and made up 20 percent of the population. Author's calculations using the 1993 and 2001 Current Population Survey.

[3] The 60th percentile of the out-of-pocket spending distribution is $1,260 per year.

expenditures and that the effects of health insurance on both health and financial risk protection need to be examined to understand the full scope of effects. The results from an expected utility analysis suggest that older individuals are willing to pay a significant amount for the risk protection that RHI coverage provides ($3,400/year for men and $3,100/year for women).

Taken together, these results imply that the large impact of RHI offer on retirement is unlikely to be driven by actual health consequences. That is, while those facing early retirement without an RHI offer may not expect an immediate negative impact on health, the results suggest that they experience significant risk aversion with respect to health shocks and high out-of-pocket expenditures. Though the health and medical expenditure risks are relatively small, they are important enough for this age group to result in a relatively large willingness-to-pay for retiree health insurance.

The results imply that the marked and continuing decline in RHI offer rates may result in much higher labor force participation rates among future cohorts of near-elderly individuals. This may substantially reverse the decline in labor force participation for older men seen since 1960 and could affect the efficiency of retirement patterns depending on the productivity of older workers and how employers respond on other margins.

The sizeable estimates of the value of retiree health insurance suggest that the decline in employer-based private group coverage may increase demand for other forms of insurance, both via public programs and private purchase. Without changes to make these sources of health insurance more affordable and available, the decline in RHI offer also may increase rates of uninsurance and financial risk among the near-elderly. While it does not appear that the decline in RHI offer will have a significant negative impact on health in the near term, continued erosion of the employment-based health insurance system in the US may have longer-term ramifications.

Section 1: Background

This section discusses recent trends in retiree health insurance and the importance and availability of other types of coverage to the near-elderly in the US. Employment-based retiree health benefits are an important source of coverage for early retirees who do not yet qualify for Medicare. In 2000, 57 percent of retirees aged 55 to 64 had employer-provided health insurance coverage, 15 percent had public coverage, 12 percent purchased private individual coverage, and 17 percent were uninsured (United States General Accounting Office 2001b). One third of those who reported employer-sponsored coverage were covered through a spouse who was either working or retired. Other sources of coverage (COBRA, individual private purchase) are often very expensive for older individuals, offer more limited coverage than group policies, and may be unavailable to those with pre-existing health conditions.

Rates of employer offer of health insurance coverage for retirees have declined significantly over the past two decades. Surveys of employers by William M. Mercer, Inc. and the Kaiser Family Foundation and Health Research and Educational Trust find that between 1988 and 2005 the percentage of large firms (200 workers or more) that offered health benefits to active workers that also offered health benefits to retirees declined from 66 percent to 33 percent. (Kaiser Family Foundation and Health Research and Educational Trust 2005) Some employers have stopped offering retiree health coverage entirely, especially for future retirees, while many others continue to offer coverage but have made changes to restrict eligibility and benefits (Fronstin, 2005).

Employers tend to make changes to retiree benefits by eliminating coverage for new workers or future retirees (Schieber 2002, Fronstin 2005). New firms may decide not to offer retiree coverage at all. Though individuals are not particularly likely to lose coverage over time, different cohorts of workers often have different retiree benefits. As a result of this cohort effect, the shares of *current* retirees reporting employer-sponsored retiree coverage have remained relatively constant since 1994 (Fronstin 2001, United States General Accounting Office 2001a). However, the decline in employer offer rates for *future* cohorts of retirees portends much lower coverage rates in the future.

A number of recent private and public-sector policy changes likely influenced the decline in employer offer rates. The Financial Accounting Standards Board's Financial Accounting Statement (FAS) 106 required publicly-held, private-sector firms to report the costs of retiree health insurance benefits using accrual accounting, which includes the costs of promised future benefits for current and future retirees. This prompted many employers to reexamine the costs of their sponsorship of retiree health benefits.[4]

The Medicare Prescription Drug, Improvement, and Modernization Act of 2003 may also impact retiree coverage offered by employers, particularly coverage for prescription drugs. Despite financial incentives that subsidize costs for employers that continue to offer coverage for prescription drugs, there is concern that they may drop

[4] Previously, employers used pay-as-you-go reporting. Similar accounting standards will be implemented for state and local government employers starting in December 2006, which will likely increase pressure on public sector employers to reduce or eliminate retiree health insurance benefits. The potential decline among one of the last remaining sectors to offer these benefits provides further motivation for understanding the implications.

coverage or reduce it to the level provided under Medicare Part D, leaving many retirees with less generous prescription drug coverage (Fronstin 2005).

Recent decisions by the US Court of Appeals for the Third Circuit and judicial action on an Equal Employment Opportunity Commission rule have limited employers' ability to offer different benefits to Medicare-eligible and early retirees. Representatives of large employers and AARP have expressed concern that these requirements and the resulting costs may induce employers to drop or reduce coverage for all retirees (Freudenheim 2005, Pear 2005).

Health insurance is particularly valuable to older individuals due to their higher likelihood of health problems and their higher health care spending. Table 1.1 shows that self-report of being in fair or poor health increases with age, as do total medical spending, out-of-pocket medical spending, and both types of spending as a percent of individual income. Furthermore, as new medical treatments become available and health care costs continue to rise rapidly, health insurance will become more valuable to older individuals.

Table 1.1
Health Status and Health Care Expenditures by Age, 2003

	Age			
	35-44	**45-54**	**55-64**	**65-74**
Self-reported fair/poor health	11%	15%	19%	23%
	0.60	0.68	0.97	1.06
Total health care expenditures in top quintile (~$3,000+)	19%	28%	41%	50%
	0.66	0.77	1.10	1.35
Total health care expenditures as a percent of individual income	7%	9%	16%	30%
	0.01	0.00	0.01	0.02
Total out-of-pocket health care expenditures in top quintile (~$750+)	17%	28%	41%	49%
	0.69	0.86	1.10	1.37
Total out-of-pocket health care expenditures as a percent of individual income	1%	2%	3%	5%
	0.00	0.00	0.00	0.00

Standard errors are presented below the percentages.
Agency for Healthcare Research and Quality. Medical Expenditure Panel Survey Household Component Data. Generated using MEPSnet/HC. http://www.meps.ahrq.gov/mepsnet/HC/MEPSnetHC.asp (May 26, 2006)

While other forms of coverage are available to protect older individuals against the financial costs associated with the health risks described above, they tend to be quite expensive and difficult to get. Collins et al. (2006) report that 55 percent of adults ages 50 to 64 with health insurance purchased on the individual market spend $3,600 or more annually on premiums. In addition to higher premiums, nearly half of older adults with individual coverage have per-person annual deductibles of $1,000 or higher and 37 percent spend $1,000 on out-of-pocket health care costs (Collins et al. 2006). Insurance costs may be even higher, or the policies unavailable entirely, for those individuals with pre-existing medical conditions (Pollitz, Sorian, and Thomas 2001). Public sources of coverage (Medicaid for low-income individuals and Medicare for those covered under Disability Insurance) provide coverage options for these subgroups, but tend to be less generous than employer-sponsored group coverage.[5]

There has been a fair amount of attention paid to the phenomenon of older workers using Disability Insurance (DI) as a way to exit the labor force (Bound and Waidmann 1992, Stapleton and Burkhauser 2003, Autor and Duggan 2006). My analysis, however, is based on individuals who report employer-sponsored coverage at baseline. Only 12 percent of individuals with employer-sponsored coverage apply for DI, compared to 39 percent of those who do not. Receipt rates are 5 and 12 percent, respectively.[6] So while labor force exit via the DI program is an important phenomenon

[5] Health care providers are less likely to accept Medicaid. Medicare often covers fewer services and charges higher copayments/coinsurance than group coverage.

[6] Among the 1,923 individuals who have not applied for DI in 1992 but apply between 1994 and 2002, 58 percent have no employer-sponsored insurance, 14 percent have ESI but no RHI offer, and 29 percent have both ESI and an RHI offer.

more generally, it does not appear to be as prominent among the subgroup that has employer-sponsored coverage.

This paper focuses on older adults between the ages of 47 and 64.[7] At age 65 nearly all individuals become eligible for Medicare, the federally-administered health insurance program for the elderly. At this point, retiree health insurance coverage fundamentally changes from a benefit that closely resembles the comprehensive coverage offered to current workers to "wrap-around" coverage that fills in Medicare's coverage gaps. So this analysis of the near-elderly focuses on a population where differences in insurance coverage are greater and where retiree health insurance coverage is most extensive.

Section 2: Related Literature

2.1 Health Insurance and Retirement

Early studies of how health insurance affects retirement used reduced-form models and found that retiree health insurance increases the probability of retirement before age 65 by 50 percent or more (Madrian 1994, Karoly and Rogowski 1994).[8] Later papers provide suggestive evidence of large effects of employer-sponsored health insurance on retirement expectations and of spouse's eligibility for Medicare at age 65 on own retirement (Hurd and McGarry 1999, Madrian and Beaulieu 1998). All of these studies lack good controls for firm-level and individual characteristics correlated with both health insurance and retirement (i.e., pension incentives, individual preferences for

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[7] About 14 percent of this group is retired full-time in 1992.

[8] These studies use variation in self-reported RHI offer to identify the effect.

leisure) which potentially bias the results upward. Studies which have used more convincing sources of exogenous variation in health insurance (Gruber and Madrian 1995, 1996) tend to confirm the basic results from the earlier studies, albeit with a slightly smaller magnitude.[9]

Structural models of the determinants of retirement can better account for pension incentives and tend to estimate smaller effects of RHI (Gustman and Steinmeier 1994, Lumsdaine, Stock, and Wise 1994, Rust and Phelan 1997). However, their results are sensitive to assumptions about the valuation and accrual patterns of RHI and several studies rely on imputed data. Other studies exploit longitudinal data available in the Health and Retirement Survey (HRS) to examine individual's behavior over time. Blau and Gilleskie (2001) condition on employer-sponsored health insurance and compare men with and without retiree coverage. They find that RHI increases the retirement rate by 26-80 percent, with no differential effects by age or health status. French and Jones (2004) find that the positive effect of RHI on retirement rises and then falls with age, peaking at age 61.

Given the range of empirical strategies and data sources, this body of work provides relatively strong evidence that health insurance has a large, positive effect on retirement. However, the generalizability of the findings are somewhat limited given that the vast majority of these studies look only at men and some focus on non-representative sub-populations. Furthermore, retiree health insurance benefits have become much less

[9] These papers estimate that the availability of COBRA coverage increases the early retirement hazard by about 32 percent. One would expect the magnitude of the effect to be smaller than that of RHI since the length of COBRA coverage is limited and individuals must pay the full cost. COBRA availability varies by state and year, due to the timing of state and federal policies.

generous since the 1980's and early 1990's when the data used in these studies were collected.

This paper contributes to this literature on a number of fronts. Detailed individual-level data on pension benefits helps account for the associated retirement incentives that plagued the early reduced-form studies.[10] I also use a nationally representative sample that includes both men and women and a longer time series than others who have analyzed this issue using the HRS. I use a number of specification checks, including propensity score weighting, to address potential correlation between retiree health insurance and individual preferences for retirement.

2.2 Health Insurance and Health

The vast majority of existing research that examines the effect of health insurance on the health of older individuals focuses on the effect of Medicare. One set of studies identifies the effect of Medicare on health by comparing the near-elderly and the elderly (age 65+), before and after Medicare implementation in 1966 (Dow 2002, Finkelstein and McKnight 2005, Meara, Cook, and Landrum 2005).[11] They find increases in hospitalization rates, a 40 percent decline in out-of-pocket spending for those in the top quartile of the out-of-pocket spending distribution, and improvements in some measures of health status, but small negative effects or no effects on mortality. Using more current data, several studies identify Medicare's effect by comparing adults on either side of the eligibility discontinuity at age 65 (McWilliams, et al. 2003, Card, Dobkin, and Maestas 2004). They find a reduction in differences in preventive screening rates and an increase

[10] The correlation coefficient between RHI and having a pension is only 0.0472 in my sample (i.e., conditional on having employer-sponsored health insurance).

[11] Lichtenberg's (2002) descriptive analysis showed that utilization of ambulatory and inpatient care increase markedly at age 65 and that Medicare is associated with a sizeable increase in survival rates.

in hospital admissions, but no significant changes in treatments for specific diseases or evidence of a discrete shift in mortality rates.[12]

To my knowledge, Cutler and Vigdor (2005) is the only paper looking at the effect of health insurance on health among the elderly that does not study Medicare. They find large effects of health insurance on self-reported health and difficulty with activities of daily living, but not on mortality, among the near-elderly.[13] Most of the evidence from studies that focus on non-elderly populations suggests small, positive effects of insurance coverage on health outcomes among "marginal" populations: infants, the sick, and the poor. However, there is also evidence to suggest that health insurance may not cause measurable improvements in health (Levy and Meltzer 2004).

My analysis focuses on the effect of retiree health insurance on health outcomes, health insurance coverage, health care utilization, and out-of-pocket medical spending for the near-elderly population. To my knowledge, the effect of RHI on these outcomes has not been studied. The wide range of health outcomes available in the Health and Retirement Survey allows me to address the effects on morbidity and quality of life and to examine a more diverse set of health care utilization outcomes than previous research. Since a primary purpose of health insurance is to protect against the risk of high medical expenses, I also estimate the effects of retiree health insurance coverage on out-of-pocket spending and estimate individuals' willingness-to-pay for this type of coverage. Examining both the health and financial protection effects provides a more

[12] Several of these papers also consider whether the effects of Medicare vary by race/ethnicity, socioeconomic status, or previous health insurance status (Dow 2002, Card, Dobkin, and Maestas 2004, Meara, Cook, and Landrum 2005). Mixed results provide evidence of larger benefits from health insurance among disadvantaged groups on some outcomes (barriers to care, high blood pressure) but not on others (hospitalization rates).

[13] The increased morbidity is associated with lower rates of medical services among the uninsured.

comprehensive picture of how RHI affects older individuals than examining the impacts on health alone.

Section 3: Methods

3.1 Conceptual framework

The expectation that retiree health insurance will affect the retirement decision comes out of more general models of labor market mobility when health insurance is tied to employment. "Job-lock" refers to the phenomenon of workers staying in jobs with relatively low marginal utility or marginal product of labor in order to maintain their health insurance benefits. Conditions that should give rise to job-lock include when employers are unable to discern the value employees assign to health insurance, when employees' individual valuation of insurance is heterogeneous, and when health insurance is not portable between jobs (Gruber 2000).

In the case of retirement, an individual with retiree health insurance will have health insurance benefits whether they retire or not, whereas an individual without retiree health insurance will not have health insurance benefits in retirement, unless they purchase their own coverage or qualify for some form of public coverage. As in the labor mobility case, an individual without RHI may continue working until they become eligible for Medicare at age 65, even if they have low marginal utility or marginal productivity of labor and a high valuation of leisure time. Models of labor productivity that include deferred compensation to induce optimal effort among workers note that a worker's wage will exceed her marginal productivity in the later years of her working life. From the employer's perspective, additional mechanisms (mandatory retirement, pension benefit design or retiree health insurance) are therefore needed to induce efficient retirement where the worker's marginal productivity equals her wage (Lazear 1979). If individuals remain at work beyond this point due to differences in health insurance coverage in working and retired states, "retirement-lock" may result in inefficiencies.[14]

In addition to impacting labor force participation and health insurance coverage, RHI may affect how individuals interact with the health care system (Figure 1.1). Retiree health insurance coverage tends to be more generous than individual coverage purchased on the private market, and certainly should improve access to care relative to being uninsured. More generous coverage can be expected to increase health care utilization and may decrease out-of-pocket spending on health care services. Over time at least, we can expect that increased access to health care services should have a positive impact on health.[15]

[14] This is also true if workers remain beyond the point where their marginal productivity equals their marginal utility of leisure.

[15] To the extent that some health care utilization may have iatrogenic effects, these may cancel out any positive impact of increased access on health. This may be especially true for healthier individuals as well as for any increase in use of health care services with low marginal benefit (i.e., moral hazard).

Figure 1.1

RHI Offer

↓

Retirement

↓

RHI Coverage

↓

Medical Care Use
and Spending

↓

Health

3.2 Data: The Health and Retirement Survey

I use data from the Health and Retirement Survey (HRS) from 1992-2002. This survey collects information from respondents who are age 51-61 in 1992, as well as from their spouse. Respondents are interviewed every two years, yielding a 10-year panel on these individuals. The HRS covers physical and mental health, insurance coverage, financial status, family support systems, labor market status, and retirement planning.[16]

The key independent variable, RHI offer, is a binary indicator for whether the respondent is able to continue their current employer-sponsored health insurance coverage in retirement.[17] I define retirement (as an outcome) as retired full-time based on the labor force status variable constructed by the RAND Center for the Study of Aging (St. Clair, et al., 2004). If the respondent is not working and not looking for work and there is any mention of retirement, labor force status is defined as retired.[18] I measure health outcomes by self-reported health (5-point scale), the change in self-reported health between waves, and the change in the number of activities of daily living (ADLs) with which the respondent indicates experiencing some difficulty. Activities of daily living include dressing, bathing, eating, using the toilet, and walking across a room, among others.

Health insurance coverage is categorized as employer-sponsored coverage from either the respondent's or spouse's employer, public coverage (Medicare, Medicaid or

[16] I use Stata's "survey" commands to account for the complex survey design in the empirical analysis.

[17] I include respondents who report RHI offer consistently in 1992 and 1994 and those for whom I can identify a reason why their answers are different between waves (i.e., changed source of coverage, changed job). These additional restrictions drop 6.5% of the sample, but this group is not significantly different based on observable characteristics.

[18] A mention of retirement can be made either through the employment status question or the question that asks the respondent whether he/she considers himself retired.

VA/Champus coverage) or uninsured.[19] The variables used to measure health care utilization include any doctors visit, number of doctors visits, any hospital stay, number of hospital stays, number of nights in the hospital, outpatient surgery, regular use of prescription drugs, and any dental visit.[20] Questions about preventive care use are asked beginning in 1996, and I group these into gender-specific measures of any preventive care use in the previous two years.[21]

My base sample consists of respondents who report having employer-sponsored health insurance coverage and are ages 47-63 in 1992.[22] This coverage may come from their own or their spouse's employer. I include years from 1994 to 2002 in which the respondent is under age 65 in the analysis.[23]

3.3 Econometric Model

As Figure 1.1 suggests, estimating the effect of RHI offer on retirement is relatively straightforward. In order to understand the effect of retiree health insurance on health, health care utilization and health care spending however, I would like to estimate the following structural model:

$$\text{Coverage}_{it} = \alpha + \beta_1 \text{Offer}_{i1} + \beta_2 \text{Year}_t + \beta_3 X_{it} + \varepsilon_{it} \qquad (1)$$

$$Y_{it} = \alpha + \delta_1 \text{Coverage}_{it} + \delta_2 \text{Year}_t + \delta_3 X_{it} + \varepsilon_{it} \qquad (2)$$

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[19] The 18 percent of observations where the respondent reports both employer-sponsored and public coverage are coded as having employer-sponsored coverage.

[20] In 1992, these utilization measures referenced the previous year, but from 1994 onward, they ask about care used in the previous two years.

[21] Cholesterol test, flu shot, breast exam, mammogram and Pap smear for women; cholesterol test, flu shot and prostate exam for men.

[22] Ages 47-63 are the middle 95% of the age distribution. Sometimes the respondent is the spouse of the HRS age-eligible individual.

[23] The HRS is conducted every two years. The sample size is 23,590 person*year observations, with an average of 4 observations per individual.

where Y_{it} is the outcome variable of interest (i.e., health status), $Offer_{i1}$ is a binary indicator for whether individual *i* has an RHI offer at baseline, $Coverage_{it}$ is a binary indicator for whether individual *i* has RHI coverage at time *t*, $Year_t$ is a set of year fixed effects, and X_{it} is a vector of covariates. This empirical analysis is complicated, however, by the endogeneity of coverage with respect to retirement. We only observe coverage for those individuals who are already retired and it is possible that individuals' retirement behavior is correlated with expectations about their health care use and/or health status. While I considered several instruments for retirement (including spouse's age, number of dependents, pension plan characteristics, and baseline assets interacted with the stock market level in each year), none convincingly satisfied the exclusion restriction. For example, conditional on the respondent's age, their spouse's age is likely to be correlated with their health. There also may be a direct income effect of assets on health care utilization and health.

Retiree health insurance *offer* is observed for everyone (both retired and not retired) and, due to high takeup rates, serves as a good proxy for coverage.[24] Furthermore, employer offer of RHI is a margin that policymakers have some control over, so understanding the intent-to-treat (ITT) effect of offer on these outcomes is important in and of itself. While retiree health insurance offer is not randomly assigned, the identifying assumption here is that, conditional on having employer sponsored health insurance coverage at baseline and controlling for other important covariates, RHI offer

[24] Among respondents in my sample (ages 47-63 with employer-sponsored coverage in 1992) who retire before age 65 and report an RHI offer in the current or any previous year, take-up rates range from 96% in 1994 to 82% in 2002. While a change in the survey questionnaire in 1996 makes the time series difficult to interpret, fairly high rates in all years suggest minimal selection in the take-up decision.

status is not based on characteristics that are correlated with the outcome of interest. That is, in the reduced-form model:

$$Y_{it} = \alpha + \beta_1 \text{Offer}_{it} + \beta_2 \text{Year}_{t} + \beta_3 X_{it} + \varepsilon_{it} \quad (3)$$

$$\text{Corr}(\text{Offer}_{i1} \mid \text{ESI}_{i1}=1, \varepsilon_{it}) = 0$$

I conduct several robustness checks, both on subsamples and using propensity score weighting, to test this assumption. I also use plausibly exogenous changes in health to assess whether the effects of RHI offer vary by health status.

Table 1.2 provides support of the identifying assumption. Conditional on having employer-sponsored health coverage at baseline, the group offered RHI looks very similar to the group not offered on a number of demographic characteristics, including age, sex, education, and household income. With the exception of those with an RHI offer being slightly more likely to have been diagnosed with an acute illness, there are no significant differences in health status or out-of-pocket spending between those with and without a RHI offer at baseline. This is true for both subjective measures, such as self-reported health, and for more objective measures of the health and longevity of immediate family members. There are some significant differences between the two groups, but in some cases the direction would bias my estimates toward zero. For example, those offered coverage have a longer average tenure at their current job, suggesting no large scale movement by respondents into jobs that offer RHI coverage as they near retirement. Lastly, we see statistically significant differences in baseline labor force participation and marital status, which I control for in the regression analysis.

Table 1.2
Summary Statistics 1992
Ages 47-63

	Total with ESI in wave 1		ESI plan covers retirees		ESI plan doesn't cover retirees		p-value
	Mean	Std. Err.	Mean	Std. Err.	Mean	Std. Err.	
Respondent Demographics							
age	55.2	0.051	55.4	0.061	54.8	0.098	<0.01
female	53%	0.006	51%	0.008	55%	0.013	0.03
white	89%	0.003	90%	0.004	86%	0.007	<0.01
black	7%	0.002	7%	0.003	8%	0.005	
hispanic	3%	0.002	3%	0.002	6%	0.005	
education	12.9	0.033	13.0	0.039	12.7	0.068	<0.01
total household income	$57,501	641	$58,190	739	$55,658	1416	0.11
Marital Status							
married	83%	0.005	85%	0.006	77%	0.011	<0.01
partnered	2%	0.002	1%	0.002	2%	0.003	
separated/divorced	10%	0.004	8%	0.005	14%	0.009	
widowed	3%	0.003	3%	0.003	5%	0.006	
never married	3%	0.002	3%	0.003	3%	0.004	
Census Region							
northeast	23%	0.004	21%	0.005	27%	0.010	<0.01
midwest	27%	0.004	28%	0.005	23%	0.009	
south	32%	0.004	32%	0.005	33%	0.010	
west	19%	0.004	19%	0.004	17%	0.009	
Labor Force Status							
works full-time	65%	0.006	61%	0.008	73%	0.011	<0.01
works part-time	10%	0.004	9%	0.005	11%	0.008	
unemployed	1%	0.001	1%	0.002	1%	0.002	
partly retired	4%	0.003	5%	0.003	2%	0.004	
fully retired	12%	0.004	14%	0.006	5%	0.006	
disabled	1%	0.001	1%	0.001	1%	0.003	
not in labor force	8%	0.004	9%	0.004	7%	0.006	
tenure at current job*	14	0.161	15	0.206	12	0.276	<0.01
have applied for DI/SSDI	0%	0.000	0%	0.000	0%	0.001	0.79
have received DI/SSDI	2%	0.002	2%	0.002	2%	0.003	0.03
covered by union	26%	0.007	28%	0.008	21%	0.012	<0.01
number employees at location	660	41	702	49	582	85	0.22
total number employees	30,613	1,466	31,948	1,741	26,043	2,763	0.07

Significance tests for continuous and binary variables were performed with an adjusted Wald test (approximate F statistic) and for categorical variables with a Pearson Chi-Sq test adjusted for survey design. P-values for Chi-Sq tests for industry (13 categories) and occupation (17 categories) are both <0.01.
Standard errors are adjusted for survey design and are clustered at the individual level.
* Subsamples: race/ethnicity (6345), tenure at current job (5022), mother alive (6391), father alive (6318), mother age at death (3526), father age at death (5186).

Table 1.2, continued
Summary Statistics 1992
Ages 47-63

	Total with ESI in wave 1		ESI plan covers retirees		ESI plan doesn't cover retirees		p-value
	Mean	Std. Err.	Mean	Std. Err.	Mean	Std. Err.	
Self-reported Health Status							
excellent	27%	0.006	27%	0.007	26%	0.011	0.57
very good	33%	0.006	33%	0.007	33%	0.012	
good	27%	0.006	28%	0.007	27%	0.011	
fair	10%	0.004	9%	0.005	11%	0.008	
poor	3%	0.002	3%	0.003	4%	0.004	
health limits work	14%	0.004	14%	0.005	13%	0.008	0.31
Other Health Status Measures							
chronic illness	56%	0.006	57%	0.008	54%	0.013	0.053
acute illness	15%	0.005	16%	0.006	12%	0.008	<0.01
number of ADLs difficult	0.027	0.003	0.024	0.004	0.029	0.006	0.44
mother alive*	45%	0.006	45%	0.008	44%	0.013	0.56
father alive*	18%	0.005	18%	0.006	18%	0.010	0.48
mother age at death*	69	0.254	69	0.310	69	0.501	0.74
father age at death*	68	0.198	68	0.241	68	0.378	0.42
out-of-pocket health spending	$1,311	62	$1,292	75	$1,381	130	0.55
Obs	6445		4606		1839		

Significance tests for continuous and binary variables were performed with an adjusted Wald test (approximate F statistic) and for categorical variables with a Pearson Chi-Sq test adjusted for survey design.
P-values for Chi-Sq tests for industry (13 categories) and occupation (17 categories) are both <0.01.
Standard errors are adjusted for survey design and are clustered at the individual level.
* Subsamples: race/ethnicity (6345), tenure at current job (5022), mother alive (6391), father alive (6318), mother age at death (3526), father age at death (5186).

The history of RHI benefits is also supportive of the assumption that RHI offer is conditionally exogenous. Retiree health insurance benefits began in the 1950's and 1960's and were used by employers to attract workers in competitive labor markets and to encourage early retirement (Atkins 1994).[25] However, medical care and health insurance costs were lower, life expectancies shorter, and there were few retirees relative to current workers at this time. As a result, RHI benefits cost very little and were an afterthought to pensions, viewed as a "throwaway" benefit and a "good-will gesture" (Rappaport and Malone 1994). By the early 1980's, 86 percent of medium and large employers were offering active employees some retiree health insurance benefits (Schieber 2002). As health care costs increased, the current worker-to-retiree ratio declined, and new accounting rules took effect in the late-1980's and early-1990's, many employers dropped future medical benefits for active employees. Most were reluctant to drop benefits for current retirees and for older employees near retirement (Schieber 2002), which is relevant for the cohort I examine in this paper.

While RHI benefits are somewhat concentrated among large, profitable firms with unionized workers with long tenures, 1988 data show that some firms in nearly every industry offer RHI benefits (Warshawsky, et al., 1993). Firm-level data from 1992 shows that only 30-40 percent of the cross-sectional variation in RHI offer is accounted for by firm size, industry and geographic region.[26] This suggests there is a significant amount of

[25] Using RHI benefits as an incentive for early retirement became even more important after the 1967 Age Discrimination in Employment Act made age discrimination and mandatory retirement (for most occupations) illegal.

[26] Author's calculations using data from the 1992 Annual Survey of Employer-Sponsored Health Benefits, conducted by the Health Research and Educational Trust. The range depends on whether county or standard metropolitan statistical area is used as the measure of geographic region.

variation in RHI offer that is unrelated to these factors which may be correlated with worker preferences for retirement or their health expectations.

While it is reasonable to expect that workers consider health insurance availability when choosing jobs, it is less obvious that they think about the coverage that will be available to them when they retire. In the 1992 HRS, the average tenure at the current job among older workers ages 47-63 is 14 years, suggesting that most workers have been in their job for a long time before approaching retirement. In fact, 22 percent of respondents ages 40-44 did not know whether retiree health insurance was offered by their employer, and 12 percent of workers ages 55-64 did not know.[27] This indicates a fair amount of myopia in terms of retirement benefits even in middle age. Gustman and Steinmeier (2004) find similar incomplete and incorrect knowledge about employer-sponsored pension benefits.

Section 4: Estimated effect of RHI offer on retirement

4.1 Econometric specification

I estimate the effect of retiree health insurance offer as the average difference in full-time retirement rates between those offered and not offered RHI, conditional on having employer-sponsored coverage at baseline and other covariates. I use a probit model:

$$\Pr(\text{FullRet}_{it}) = \Phi(\alpha + \beta_1 \text{Offer}_{i1} + \beta_2 \text{Year}_t + X_{it}\beta_3 + \varepsilon_{it}) \qquad (4)$$

where FullRet_{it} is a binary indicator for whether individual *i* is retired full-time in year *t*, Offer_{i1} is a binary indicator for whether individual *i* is offered RHI at baseline, and Year_t represents a set of year fixed effects. I restrict the sample to respondents who are not

[27] Respondents who do not know if they are offered RHI are not included in the analysis.

retired full-time at baseline. The covariates X_{it} include a full set of age dummies and controls for race/ethnicity (white, black, Hispanic), and education. Baseline covariates are marital status, self-reported health, and spousal characteristics (age, education, and health status). I also control for household income, household assets, pension characteristics (none, defined benefit, defined contribution, or both and age of vesting), and industry and occupation dummies for the current job (all at baseline).[28] I cluster the standard errors at the individual level.

The normative interpretation of the large effects I find of RHI offer on retirement will depend, in part, on whether this effect varies by the respondent's health status. Furthermore, before estimating the effects of RHI offer on health outcomes in retirement, I must first assess whether there is differential selection into retirement with respect to health. To address these issues, the key econometric concern is to find variation in health status that is plausibly exogenous to insurance status (RHI offer). I argue that health shocks are plausibly exogenous since, while individuals may be able to predict the probability that they will have a health shock (and select insurance coverage accordingly), it is unlikely that they can predict the timing or severity.[29] As a result, conditional on baseline health status, the timing and severity of the health shock is unlikely to be correlated with RHI offer. Furthermore, if the health shock takes place before RHI coverage is in effect (i.e., before retirement), then RHI offer and the health shock should not be correlated via the causal pathway from coverage to health. The assumption therefore is:

[28] There are 6 education categories, 7 for marital status, 5 for self-reported health, 13 for industry and 17 for occupation.

[29] This approach is also used by Cutler, McClellan, and Newhouse (2000) and Cutler and Vigdor (2005) to deal with selection into health insurance coverage.

$$Y_{it} = \alpha + \beta_1 \text{Offer}_{it} + \beta_2 \text{HealthShock}_{it} + \beta_3 \text{Year}_t + \beta_4 X_{it} + \varepsilon_{it} \quad (5)$$

$$\text{Corr}(\text{HealthShock}_{it} \mid \text{HealthStatus}_{i1}, \varepsilon_{it}) = 0$$

Based on the subsample of individuals who are not retired full-time in 1992, I estimate:

$$\Pr(\text{FullRet}_{it}) = \Phi(\alpha + \beta_1 \text{Offer}_{i1} + \beta_2 \text{NewChronic}_{it} + \beta_3 \text{Offer}_{i1}{*}\text{NewChronic}_{it} + \beta_4 \text{Year}_t + X_{it}\beta_5 + \varepsilon_{it}) \quad (6)$$

where FullRet_{it}, Offer_{i1}, Year_t, and X_{it} are defined as above. NewChronic is a binary indicator for whether individual *i* receives a *new* diagnosis of the following conditions in year *t* (before retirement): congestive heart failure, high blood pressure, diabetes, lung disease, arthritis or a psychiatric illness.[30] I estimate a similar model using acute health shocks instead of chronic ones. NewAcute is a binary indicator for a new diagnosis of a heart attack, angina, stroke or cancer in year *t*.[31] β_3 is the coefficient of interest, estimating the incremental effect of RHI offer, conditional on experiencing a pre-retirement health shock. I also estimate models that include lagged health shocks, and their interaction with RHI offer, in order to estimate the effect of a health shock in year *t-2* on retirement in year *t*.

4.2 Results: Full-Time Retirement

Table 1.3 presents results from estimating equation 4 on a binary indicator for full-time retirement, conditional on not being retired full-time in 1992. Retiree health insurance offer has a large and significant positive effect on the probability of retiring before age 65. Without controlling for covariates, it increases the probability of

[30] Fifty-six percent of respondents report a chronic condition at baseline and an average of fifteen percent report a new chronic health shock each year between 1994 and 2002.

[31] Fifteen percent of respondents report an acute condition at baseline and an average of three percent report a new acute health shock each year between 1994 and 2002.

retirement by 8 percentage points, or by 40 percent (bottom of column a). Controlling for covariates, including pension characteristics that are likely to be important determinants of early retirement, decreases the estimated effect only slightly, to 7 percentage points (35 percent, column b).[32] The effect size is the same for men and women (columns d and f). Those in good, fair or poor health are significantly more likely to be retired before age 65 than those in excellent health. Hispanic women, women with higher education and divorced women are less likely to be retired than white women, women who did not finish high school and married women, respectively.[33]

[32] The fact that the covariates change the magnitude of the offer coefficient so slightly also supports the assumption that RHI offer is conditionally exogenous to other covariates that impact early retirement.

[33] Including all covariates except the pension vesting age dummies, the pension type covariates are positive and statistically significant (DB: .0396 [.0110]; DC: .0280 [.0132]; Both: .0625 [.0329]).

Table 1.3
Estimated Effect of Retiree HI Offer on Full-Time Retirement
Employer-Sponsored Health Insurance and Not Full-Time Retired in 1992
Less Than Age 65

Full-Time Retired

	Total		**Men**		**Women**	
	a	**b**	**c**	**d**	**e**	**f**
Offer	0.0762 ***	0.0698 ***	0.0877 ***	0.0653 ***	0.0650 ***	0.0730 ***
	0.0087	0.0102	0.0133	0.0144	0.0112	0.0133
female	-0.0214 *	-0.0026				
	0.0102	0.0133				
Race/Ethnicity						
black	-0.0075	-0.0243	-0.0060	-0.0318	0.0034	-0.0090
	0.0128	0.0144	0.0217	0.0230	0.0161	0.0182
hispanic	-0.0330	-0.0347	-0.0094	-0.0006	-0.0471 *	-0.0620 *
	0.0172	0.0191	0.0266	0.0285	0.0226	0.0247
Education						
high school	0.0002	-0.0019	0.0217	0.0436 *	-0.0172	-0.0503 *
	0.0126	0.0155	0.0185	0.0203	0.0167	0.0215
some college	-0.0202	-0.0268	-0.0256	-0.0046	-0.0206	-0.0636 **
	0.0140	0.0170	0.0201	0.0227	0.0186	0.0227
college	-0.0218	-0.0344	0.0007	0.0165	-0.0291	-0.0793 **
	0.0175	0.0206	0.0259	0.0302	0.0234	0.0249
more than college	-0.0078	-0.0165	0.0021	0.0493	0.0025	-0.0574 *
	0.0179	0.0230	0.0264	0.0340	0.0247	0.0287
Marital Status						
married, sp. absent	-0.0315	-0.0097	0.0128	0.0197	-0.0721	0.0261
	0.0512	0.0598	0.0721	0.0748	0.0593	0.0872
partnered	-0.0291	-0.0456	-0.0226	-0.0462	-0.0547	-0.0552
	0.0261	0.0287	0.0351	0.0368	0.0362	0.0385
separated	0.0182	0.0551	-0.0509	0.0914	0.1025	-0.0691
	0.0577	0.0785	0.0588	0.0965	0.1180	0.0872
divorced	0.0431	0.0322	0.0819	0.1481	0.0232	-0.1354 *
	0.0501	0.0631	0.0619	0.0838	0.0905	0.0605
widowed	0.0768	0.0289	0.2520 *	0.1929	0.0822	-0.1170
	0.0577	0.0663	0.1132	0.1127	0.1027	0.0606
never married	0.0596	0.0674	-0.0147	0.0629	0.1187	-0.0578
	0.0574	0.0739	0.0612	0.0899	0.1124	0.0871

* p<0.05; ** p<0.01; *** p<0.001
Marginal effects from probit models.
Standard errors are adjusted for survey design and are clustered at the individual level.
Mean percent retired full-time (if not retired full-time at baseline): 20% total, 21% men, 20% women.
Omitted categories are white, less than high school, married, excellent health, no pension, and 1994.
Year dummies are in all models and are all positive and statistically significant.
Columns a, c, and e include age dummies and controls for the spouse's age, education and health status. Columns b, d, and f add industry (13) and occupation (17) dummies, as well as dummies for the earliest age at which the respondent can receive their pension.

Table 1.3, continued
Estimated Effect of Retiree HI Offer on Full-Time Retirement
Employer-Sponsored Health Insurance and Not Full-Time Retired in 1992
Less Than Age 65

Full-Time Retired

	Total		**Men**		**Women**	
	a	**b**	**c**	**d**	**e**	**f**
Self-Reported Health						
health very good	0.0020	-0.0009	0.0121	0.0033	-0.0079	0.0000
	0.0112	0.0124	0.0170	0.0175	0.0144	0.0165
health good	0.0502 ***	0.0451 **	0.0714 ***	0.0543 **	0.0288	0.0332
	0.0125	0.0145	0.0190	0.0204	0.0157	0.0186
health fair	0.1614 ***	0.1510 ***	0.1632 ***	0.1005 **	0.1547 ***	0.1943 ***
	0.0197	0.0237	0.0302	0.0328	0.0255	0.0328
health poor	0.1545 ***	0.1111 *	0.1993 ***	0.1532 *	0.1166 **	0.0555
	0.0351	0.0501	0.0537	0.0648	0.0445	0.0658
Financial Variables						
household income	0.0002 *	0.0001	-0.0001	-0.0002	0.0005 ***	0.0004 **
	0.0001	0.0001	0.0002	0.0002	0.0001	0.0001
household assets	0.0000 *	0.0000	0.0000	0.0000	0.0000 *	0.0000
	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
DB pension		0.2007		0.0520		0.0169
		0.1370		0.0597		0.0386
DC pension		0.2178		0.0331		0.0217
		0.1614		0.0570		0.0339
Both pension		0.2555		0.0132		0.0919
		0.1761		0.0697		0.0652
cons	-1.8144 **	-6.6127 ***	-6.4444 ***	-6.1782	-1.9493 **	-5.6292 ***
	0.5575	0.5758	0.4753	.	0.6448	0.7903
Offer w/o covars	0.0805 ***		0.0895 ***		0.0731 ***	
	0.0085		0.0131		0.0112	
N	19904	13747	8757	6486	11147	7261

* p<0.05; ** p<0.01; *** p<0.001
Marginal effects from probit models.
Standard errors are adjusted for survey design and are clustered at the individual level.
Mean percent retired full-time (if not retired full-time at baseline): 20% total, 21% men, 20% women.
Omitted categories are white, less than high school, married, excellent health, no pension, and 1994.
Year dummies are in all models and are all positive and statistically significant.
Columns a, c, and e include age dummies and controls for the spouse's age, education and health status. Columns b, d, and f add industry (13) and occupation (17) dummies, as well as dummies for the earliest age at which the respondent can receive their pension.

The age dummies between ages 53 and 64 tend to be statistically significant and increasing with age (not shown). They range in magnitude from a 5 percentage point increase in the probability of early retirement at age 53 (relative to age 48) to an increase of about 35 percentage points at age 64. The magnitude and significance levels of these effects are similar for men and women.[34] In general, the industry and occupation dummies are not statistically significant.

These estimates are similar to those from the other studies discussed in section 2.1. Most notably, my estimate of 40 percent is only slightly larger than the 32 percent estimated effect of COBRA coverage, which provides the most convincing source of exogenous variation in health insurance coverage availability of the previous studies. One would expect the effect of RHI offer to be larger, since individuals have to pay the full cost of COBRA coverage and RHI coverage extends for a longer period.

4.3 Results: Differential Retirement by Health Status

Table 1.4 shows that the positive main effect of RHI offer on full-time retirement remains large and significant, controlling for both chronic and acute health shocks (6-8 percentage points). Those who experience a chronic health shock are nearly 4 percentage points less likely to retire while the large positive direct effect of an acute health shock is only seen in the lagged variable (16 percentage points).[35] There is no evidence of significant differential retirement as a function of health status, conditional on having an

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[34] I also ran a separate model that included interactions of the age dummies with RHI offer to investigate whether RHI offer has different effects at different ages. The main effects of age are positive and significant at the 5% level for ages 59 to 64, ranging in magnitude from 4 to 24 percentage points. The offer*age coefficients are positive and significant only for ages 60 to 64, ranging from 8 to 17 percentage points.

[35] The negative coefficient on a chronic health shock could be due to respondents with these conditions needing to maintain their incomes to pay ongoing medical bills.

RHI offer.[36] These results suggest that, conditional on being in poor health, retirement behavior is not significantly different among those who have an RHI offer. This allays some concerns about the equity implications of RHI with respect to older workers who become sick. This also provides the first piece of evidence to support looking at health outcomes among retirees.

[36] This is also true conditional on not having an RHI offer.

Table 1.4
Estimated Effect of Retiree HI Offer on Full-Time Retirement: Health Shocks
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Chronic Health Shock		Acute Health Shock	
	a	**b**	**c**	**d**
Chronic Health Shock				
Offer	0.0649 ***	0.0843 ***	0.0672 ***	0.0818 ***
	0.0107	0.0135	0.0104	0.0126
Health Shock	-0.0368 *	-0.0390	0.0562	0.0571
	0.0185	0.0231	0.0439	0.0553
Offer*Shock	0.0258	-0.0183	-0.0393	-0.0800
	0.0253	0.0288	0.0401	0.0455
Health Shock (2 yr lag)		0.0320		0.1559 *
		0.0245		0.0620
Offer*Lagged Shock		-0.0373		-0.0670
		0.0256		0.0501
N	12,366	9,516	12,085	9,387

* p<0.05; ** p<0.01; *** p<0.001
Marginal effects from probit models.
Standard errors are adjusted for survey design and are clustered at the individual level.
Included sample is not retired full-time at baseline.
Controls: sex, age, race, education, marital status, baseline health status, spouse's age, education and baseline health, pension, pension vesting age, industry, occupation, household income and assets, and year.

Section 5: Estimated effect of RHI offer on health, health insurance coverage and health care utilization

5.1 Econometric specification

To estimate the effect of RHI offer on health and health care utilization, I use equation 4 for the binary outcomes (fair/poor health, any doctor's visit, prescription drug use, and outpatient surgery) and use a similar OLS model for continuous outcome variables (change in self-reported health and change in ADLs). Covariates in the health outcomes regressions include a full set of age dummies and controls for sex, race/ethnicity, education, and baseline health status. The health care utilization models add census division (9 categories) to control for health care practice patterns that vary geographically.

To estimate the effect of RHI offer on health insurance coverage, I use a multinomial logit model so that the log odds of individual *i* having health insurance *j* relative to having employer-sponsored coverage (*j* = 1) in time *t* is a linear function of RHI offer, year fixed effects and other covariates.

$$\log\left(\Pr(HI_{itj})/\Pr(HI_{it1})\right) = \alpha + \beta_1 \, Offer_{i1} + \beta_2 \, Year_{t} + X_{it}\,\beta_3 + \varepsilon_{it} \qquad (7)$$

For the health insurance variable HI_{itj}, $j = 3$: employer-sponsored coverage, public coverage and uninsured. The covariates are age, sex, race/ethnicity, education, and baseline health and marital status.

There is an obvious concern with looking at the effect of RHI offer on health outcomes in retirement, since I have just shown that RHI offer has a large and statistically significant effect on retirement itself. Any differential selection into retirement based on health, or any other omitted variable that is correlated with RHI offer and health, will bias estimates based on the retired subsample. I combine several pieces of suggestive evidence to address this concern.

First, the results in Table 1.4 show no evidence of significant differential selection into retirement based on health status. Second, I estimate the effects of RHI offer on health outcomes for the total sample (both retired and not retired) and scale those estimates by the percent retired. This provides a sense of the magnitude of the effect among the retired group without estimating it on a selected sample. Third, I split the sample into retired and non-retired individual-years and estimate the effects separately. Estimating the model using the not retired subsample serves as a placebo test, since RHI coverage does not take effect until after retirement. If we do not see significant effects of RHI offer among the not retired subsample, this provides additional evidence that the estimates from the retired subsample are free from significant selection bias.

5.2 Results: Health Status

Table 1.5 shows the estimated effects of RHI offer on three measures of health status: a binary indicator for self-reported fair or poor health, the change in self-reported health status between waves (ranging from -4 to 4), and the change in the number of ADLs performed with some difficulty (ranging from -5 to 5). For all of these outcomes, a negative coefficient indicates an improvement in health. For all three measures, we see no significant effect of RHI offer on health for the total sample after controlling for covariates (columns a, d, and g). Scaling these effect sizes by the average percent retired between 1994 and 2002 (24.6 percent), I calculate implied effects of RHI offer among retirees to be a nearly 4 percentage point decline in the probability of being in fair/poor health (37 percent), no effect on the change in self-reported health, and decline of .0124 in the change in number of ADLs performed with difficulty (38 percent).

Table 1.5
Estimated Effect of Retiree HI Offer on Health Outcomes
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Fair/Poor Health (not f/p health in 1992)			**Change in Self-Reported Health (ranges -4 to 4)**		
	Total	**Not Retired**	**Retired**	**Total**	**Not Retired**	**Retired**
	a	**b**	**c**	**d**	**e**	**f**
Offer	-0.0083	-0.0089	-0.0308	0.0001	-0.0038	-0.0038
	0.0068	0.0069	0.0179	0.0079	0.0091	0.0262
female	-0.0026	0.0051	-0.0314 *	-0.0149 *	-0.0081	-0.0437 *
	0.0062	0.0064	0.0138	0.0073	0.0089	0.0205
black	0.0368 ***	0.0331 **	0.0494 *	0.0516 ***	0.0495 ***	0.0601
	0.0103	0.0109	0.0204	0.0111	0.0139	0.0322
hispanic	0.0528 **	0.0575 **	0.0547	0.0427 **	0.0574 **	-0.0049
	0.0176	0.0182	0.0390	0.0161	0.0202	0.0558
high school	-0.0547 ***	-0.0458 ***	-0.0924 ***	-0.0625 ***	-0.0746 ***	-0.0358
	0.0079	0.0085	0.0159	0.0116	0.0145	0.0318
some college	-0.0606 ***	-0.0518 ***	-0.0996 ***	-0.0738 ***	-0.0893 ***	-0.0362
	0.0072	0.0076	0.0159	0.0126	0.0157	0.0362
college	-0.0690 ***	-0.0571 ***	-0.1198 ***	-0.1042 ***	-0.1162 ***	-0.0951 *
	0.0069	0.0074	0.0155	0.0153	0.0186	0.0423
more than college	-0.0758 ***	-0.0696 ***	-0.1091 ***	-0.1120 ***	-0.1352 ***	-0.0508
	0.0068	0.0070	0.0158	0.0138	0.0169	0.0384
health very good	0.0482 ***	0.0443 ***	0.0693 **	-0.1272 ***	-0.1456 ***	-0.0642 *
	0.0096	0.0100	0.0209	0.0088	0.0104	0.0280
health good	0.1748 ***	0.1593 ***	0.2240 ***	-0.2625 ***	-0.2959 ***	-0.1712 ***
	0.0114	0.0121	0.0239	0.0097	0.0119	0.0286
health fair				-0.3624 ***	-0.4154 ***	-0.2614 ***
				0.0146	0.0207	0.0342
health poor				-0.5088 ***	-0.5650 ***	-0.4097 ***
				0.0268	0.0430	0.0491
cons	-6.8972 ***	-6.9820 ***	-5.2450 ***	0.3188 *	0.3690 *	0.0834
	0.2318	0.4359	0.5056	0.1492	0.1588	0.0450
Offer w/o covars	-0.0887 *			0.0011		
	0.0418			0.0085		
Offer scaled by percent retired	-0.0366			0.0002		
R squared				0.0389	0.0442	0.0319
N	18870	14249	4285	21504	15691	5294

* p<0.05; ** p<0.01; *** p<0.001
Marginal effects from probit models (fair/poor health), OLS models for change in self-reported health and change in ADLS
Standard errors are adjusted for survey design and are clustered at the individual level.
Rates of fair/poor health (not in fair/poor health at baseline): 10% total, 8% not retired, 15% retired.
Mean change in self-reported health: .081 total, .086 not retired, .071 retired. Mean change in ADLs: .032 total, .025 not retired, .049 retired. Mean percent retired: 22.6% if not in fair/poor health at baseline, 24.6% total. Age and year dummies are included in all models. Omitted categories are white, less than high school, excellent health, and 1994.

Table 1.5, continued
Estimated Effect of Retiree HI Offer on Health Outcomes
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Change in ADLs Performed with Some Difficulty (ranges -5 to 5)		
	Total **g**	**Not Retired** **h**	**Retired** **i**
Offer	-0.0030	0.0026	-0.0396
	0.0049	0.0053	0.0223
female	0.0063	0.0080	0.0019
	0.0043	0.0048	0.0139
black	0.0152	0.0096	0.0232
	0.0082	0.0085	0.0254
hispanic	0.0005	0.0077	-0.0283
	0.0116	0.0111	0.0434
high school	-0.0073	0.0025	-0.0381
	0.0077	0.0087	0.0232
some college	-0.0011	0.0099	-0.0424
	0.0082	0.0091	0.0255
college	0.0028	0.0011	0.0055
	0.0093	0.0091	0.0309
more than college	0.0004	0.0005	-0.0005
	0.0089	0.0089	0.0293
health very good	0.0055	0.0091 *	-0.0067
	0.0036	0.0041	0.0130
health good	0.0260 ***	0.0216 ***	0.0310 *
	0.0051	0.0061	0.0152
health fair	0.0917 ***	0.0814 ***	0.0774 **
	0.0132	0.0159	0.0288
health poor	0.1549 ***	0.1356 **	0.1621 *
	0.0328	0.0518	0.0703
cons	0.0517	0.0428	0.0758 *
	0.0622	0.0671	0.0348
Offer w/o covars	-0.0058		
	0.0050		
Offer scaled by percent retired	-0.0124		
R squared	0.0082	0.0080	0.0189
N	21496	15684	5296

* p<0.05; ** p<0.01; *** p<0.001
Marginal effects from probit models (fair/poor health), OLS models for change in self-reported health and change in ADLS.
Standard errors are adjusted for survey design and are clustered at the individual level. Rates of fair/poor health (not in fair/poor health at baseline): 10% total, 8% not retired, 15% retired.
Mean change in self-reported health: .081 total, .086 not retired, .071 retired. Mean change in ADLs: .032 total, .025 not retired, .049 retired. Mean percent retired: 22.6% if not in fair/poor health at baseline, 24.6% total. Age and year dummies are included in all models. Omitted categories are white, less than high school, excellent health, and 1994.

The precisely estimated null effects of RHI offer in columns b, e, and h provide evidence of no significant selection into retirement based on health. Turning to the retired group, the effect of RHI offer on health should occur through the causal channel of increased access to medical care. While the results are not statistically significant at the 95 percent level due to smaller sample sizes, they mirror the scaled effects and are suggestive of RHI offer improving health outcomes. RHI offer reduces the probability of being in fair/poor health by 3 percentage points, or 21 percent (column c), and reduces the change in the number of ADLs reported by about 81 percent of the average (column i).[37]

[37] In a separate analysis that examines respondents ages 65-74, I find a statistically significant effect of RHI offer on the change in self-reported health among retirees (-0.0948, se 0.0329). The estimates of the effects on self-reported health and change in the number of ADLs reported are not significant.

5.3 Results: Health Insurance Coverage

Table 1.6 shows that RHI offer increases the probability of having employer-sponsored coverage by 5 percentage points, or about 6 percent (column a), and decreases the probability of having public health insurance coverage by 2 percentage points, or 60 percent (column d).[38] On net, the probability of being uninsured declines by 3 percentage points, or 55 percent (column g). Scaling these effects by the average percent retired implies a 21 percentage point increase in the probability of having employer-sponsored coverage, a 10 percentage point decline in the probability of having public coverage, and an 11 percentage point decline in the probability of being uninsured among retirees. We see significant effects of RHI offer on coverage among both the not retired and retired subsamples and much larger effects among retirees, though they are not quite as large as the scaled estimates.

[38] In separate models, I find precisely-estimated zero effects of RHI offer on DI application, DI receipt and Medicare coverage. Since I am restricting my analysis to adults under age 65, respondents who indicate Medicare coverage are assumed to be DI recipients. Control variables include age dummies, sex, race, education, baseline health status, baseline marital status, household income, household assets, pension characteristics, and fixed effects for industry, occupation and year.

Table 1.6
Estimated Effect of Retiree HI Offer on Health Insurance Coverage
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	ESI			**Public**		
	Total	**Not Retired**	**Retired**	**Total**	**Not Retired**	**Retired**
	a	**b**	**c**	**d**	**e**	**f**
Offer	0.0516 ***	0.0307 ***	0.1369 ***	-0.0240 ***	-0.0133 ***	-0.0589 ***
	0.0060	0.0057	0.0173	0.0044	0.0040	0.0120
age	-0.0039 ***	-0.0038 ***	-0.0009	0.0025 ***	0.0027 ***	0.0005
	0.0008	0.0009	0.0015	0.0006	0.0007	0.0010
female	-0.0104 **	-0.0109 *	-0.0081	0.0080 **	0.0084 **	0.0055
	0.0041	0.0044	0.0077	0.0028	0.0031	0.0050
Race/Ethnicity						
black	-0.0108	-0.0066	-0.0152	0.0091	0.0007	0.0230 *
	0.0069	0.0073	0.0123	0.0053	0.0049	0.0103
hispanic	-0.0175	-0.0152	-0.0454 *	0.0014	0.0035	-0.0012
	0.0092	0.0100	0.0225	0.0054	0.0066	0.0096
Education						
high school	0.0178 ***	0.0155 **	0.0214 *	-0.0104 ***	-0.0099 **	-0.0123 *
	0.0047	0.0054	0.0086	0.0031	0.0036	0.0053
some college	0.0193 ***	0.0148 **	0.0305 ***	-0.0101 **	-0.0103 **	-0.0105
	0.0049	0.0054	0.0089	0.0033	0.0036	0.0057
college	0.0217 ***	0.0182 **	0.0291 **	-0.0149 ***	-0.0148 ***	-0.0123 *
	0.0053	0.0060	0.0103	0.0031	0.0034	0.0060
more than college	0.0296 ***	0.0259 ***	0.0314 **	-0.0144 ***	-0.0110 **	-0.0189 ***
	0.0050	0.0055	0.0103	0.0034	0.0039	0.0060
Self-Reported Health						
health very good	0.0068	0.0010	0.0185	-0.0050	-0.0026	-0.0067
	0.0051	0.0055	0.0103	0.0036	0.0040	0.0067
health good	-0.0002	-0.0019	0.0027	0.0004	0.0013	0.0010
	0.0054	0.0059	0.0106	0.0038	0.0043	0.0073
health fair	-0.0230 *	-0.0099	-0.0297	0.0172 *	0.0078	0.0248
	0.0094	0.0093	0.0169	0.0076	0.0071	0.0136
health poor	-0.0613 ***	-0.0711 **	-0.0380	0.0421 **	0.0514 **	0.0270
	0.0188	0.0248	0.0258	0.0152	0.0197	0.0207
Marital Status						
married, sp. absent	-0.0263	-0.0457	0.0314	0.0034	0.0212	-0.0329 ***
	0.0415	0.0529	0.0295	0.0219	0.0341	0.0037
partnered	-0.0032	-0.0013	-0.0322	-0.0140 **	-0.0090	-0.0207 **
	0.0145	0.0129	0.0553	0.0049	0.0063	0.0077
separated	-0.0016	-0.0064	0.0016	-0.0035	-0.0023	0.0002
	0.0174	0.0210	0.0266	0.0115	0.0167	0.0156
divorced	-0.0096	-0.0024	-0.0587 *	0.0018	-0.0005	0.0151
	0.0083	0.0080	0.0246	0.0051	0.0053	0.0109
widowed	-0.0417 **	-0.0224	-0.1320 ***	0.0182	0.0168	0.0329
	0.0156	0.0161	0.0409	0.0104	0.0122	0.0226
never married	0.0059	0.0172	-0.0427	-0.0065	-0.0109	0.0069
	0.0109	0.0105	0.0280	0.0067	0.0077	0.0133
Offer w/o covars	0.0689 ***			-0.0306 ***		
	0.00677			0.0046		
Offer scaled by percent retired	0.20985			-0.0974		
N	20,760			20,760		

* p<0.05; ** p<0.01; *** p<0.001
Standard errors are adjusted for the survey design and clustered at the individual level.
Marginal effects from a multinomial logit model where the outcome variable takes on 3 values: ESI, Public, and Uninsurec
Mean employer coverage rates: 89% total, 91% not retired, 82% retired. Mean public coverage rates: 4% total, 3% not retired, 6% retired. Mean uninsured rates: 5% total, 5% not retired, 7% retired. Omitted categories are white, less than high school, married, excellent health, and 1994.

Table 1.6, continued
Estimated Effect of Retiree HI Offer on Health Insurance Coverage
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Uninsured		
	Total g	**Not Retired h**	**Retired i**
Offer	-0.0277 ***	-0.0174 ***	-0.0781 ***
	0.0041	0.0040	0.0129
age	0.0014 ***	0.0011 *	0.0003
	0.0004	0.0005	0.0010
female	0.0024	0.0025	0.0026
	0.0027	0.0029	0.0056
Race/Ethnicity			
black	0.0017	0.0060	-0.0078
	0.0042	0.0051	0.0067
hispanic	0.0161 *	0.0117	0.0466 *
	0.0067	0.0068	0.0197
Education			
high school	-0.0073 *	-0.0056	-0.0091
	0.0032	0.0036	0.0063
some college	-0.0091 **	-0.0045	-0.0200 ***
	0.0033	0.0038	0.0062
college	-0.0068	-0.0034	-0.0168 *
	0.0041	0.0048	0.0073
more than college	-0.0153 ***	-0.0150 ***	-0.0124
	0.0033	0.0034	0.0082
Self-Reported Health			
health very good	-0.0018	0.0017	-0.0117
	0.0034	0.0036	0.0075
health good	-0.0002	0.0006	-0.0037
	0.0036	0.0039	0.0071
health fair	0.0059	0.0021	0.0049
	0.0052	0.0055	0.0094
health poor	0.0192	0.0197	0.0110
	0.0105	0.0159	0.0146
Marital Status			
married, sp. absent	0.0230	0.0245	0.0015
	0.0298	0.0328	0.0293
partnered	0.0172	0.0102	0.0529
	0.0132	0.0109	0.0533
separated	0.0051	0.0087	-0.0018
	0.0126	0.0131	0.0192
divorced	0.0078	0.0029	0.0435 *
	0.0061	0.0058	0.0215
widowed	0.0235 *	0.0057	0.0992 **
	0.0111	0.0090	0.0342
never married	0.0007	-0.0062	0.0358
	0.0087	0.0072	0.0247
Offer w/o covars	-0.0383 ***		
	0.0048		
Offer scaled by percent retired	-0.1125		
N	20,760		

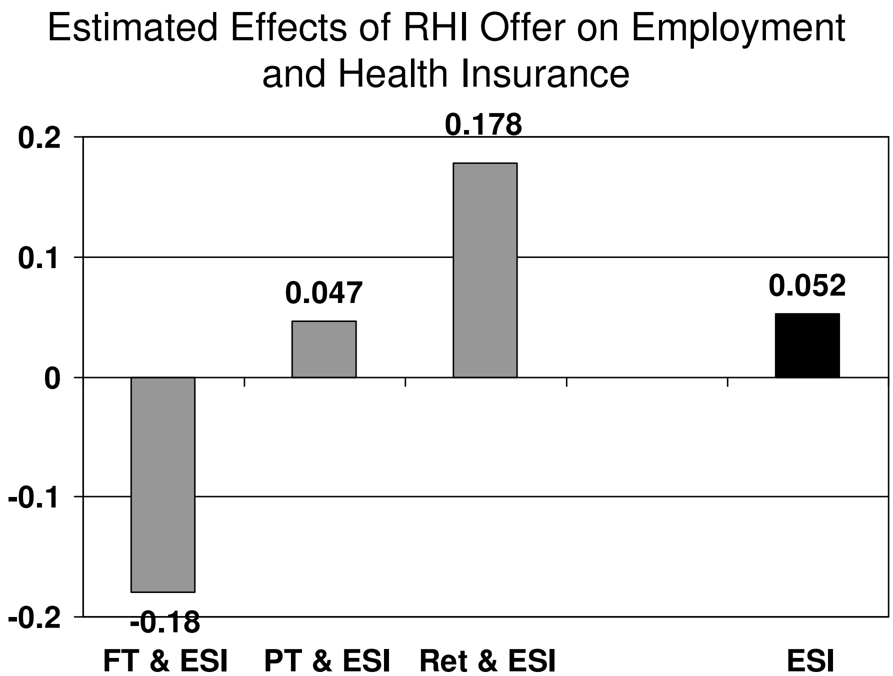
* p<0.05; ** p<0.01; *** p<0.001
Standard errors are adjusted for the survey design and clustered at the individual level.
Marginal effects from a multinomial logit model where the outcome variabl takes on 3 values: ESI, Public, and Uninsured.
Mean employer coverage rates: 89% total, 91% not retired, 82% retired. Mean public coverage rates: 4% total, 3% not retired, 6% retired. Mean uninsured rates: 5% total, 5% not retired, 7% retired. Omitted categories are white, less than high school, married, excellent health, and 1994.

To understand the effects of RHI offer on retirement and coverage jointly, I estimate a multinomial logit model similar to equation 7 with a dependent variable that combines labor force status and health insurance coverage.[39] Figure 1.2 presents selected results from this analysis. The grey bars on the left show the estimated effects of RHI offer on three of the joint labor force/health insurance outcomes. Those with an RHI offer are 18 percentage points less likely to be working full-time with employer-sponsored coverage, 5 percentage points more likely to be working part-time with ESI and 18 percentage points more likely to be retired full-time with ESI. Combined, these results are equivalent to the 5 percentage point increase in the probability of having ESI shown in Table 1.6. These results suggest that though the net effect of RHI offer on ESI coverage is relatively small, there are large differences in how individuals get this coverage. Those with an RHI offer are more likely to be retired and have retiree coverage from their employer while those without an RHI offer are more likely to remain at work and have current worker ESI.[40]

[39] The dependent variable takes on 9 values: employer-sponsored coverage (ESI) and full-time work, public coverage and full-time work, uninsured and full-time work, ESI and part-time work, public and part-time work, uninsured and part-time work, ESI and full-time retired, public and full-time retired, and uninsured and full-time retired. Control variables include age dummies, sex, race, education, baseline health status and year fixed effects.

[40] There are not significant differences in the estimated effects of RHI offer on public coverage or uninsurance across labor force participation categories.

Figure 1.2



Marginal effects from multinomial logit models. All coefficients are statistically significant at p<.001.
The left three bars are from a model where the dependent variable has 9 work/HI values. The right-hand bar is from a model where the dependent variable has 3 HI values.

5.4 Results: Health Care Utilization

I used measures for several different types of health care utilization, including outpatient, inpatient, and preventive care. For outpatient services, such as doctor's visits, outpatient surgery and prescription drug use, the coefficients on RHI offer are positive and significant for the total sample (Table 1.7). For example, the probability of having a doctor's visit in the past 2 years increases by 1 percentage point (1 percent, column a), the probability of using prescription medications increases by 3 percentage points (4 percent, column d), and the probability of outpatient surgery increases by 2 percentage points (11 percent, column g). Scaling these effects by the average percent retired implies effect sizes among retirees of increases of 5 percentage points, 12 percentage points, and 8 percentage points, respectively. The coefficients on RHI offer are relatively precisely estimated zeros for the measures of inpatient care and preventive services for both men and women (not shown).

Table 1.7
Estimated Effect of Retiree HI Offer on Health Care Utilization
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Any Doctors Visit			Regularly Use Prescription Drugs		
	Total **a**	**Not Retired** **b**	**Retired** **c**	**Total** **d**	**Not Retired** **e**	**Retired** **f**
Offer	0.0120 *	0.0116	0.0148	0.0289 *	0.0200	0.0290
	0.0056	0.0061	0.0117	0.0120	0.0132	0.0214
female	0.0483 ***	0.0456 ***	0.0516 ***	0.1654 ***	0.1780 ***	0.1260 ***
	0.0052	0.0057	0.0094	0.0108	0.0123	0.0173
black	0.0150 *	0.0175 *	0.0048	-0.0222	-0.0217	-0.0096
	0.0067	0.0074	0.0135	0.0163	0.0186	0.0256
hispanic	-0.0135	-0.0096	-0.0309	-0.0912 ***	-0.0903 ***	-0.0742
	0.0108	0.0118	0.0214	0.0247	0.0273	0.0421
high school	0.0329 ***	0.0341 ***	0.0326 ***	0.0629 ***	0.0778 ***	0.0355
	0.0061	0.0068	0.0102	0.0153	0.0177	0.0233
some college	0.0451 ***	0.0485 ***	0.0362 ***	0.0964 ***	0.1177 ***	0.0426
	0.0055	0.0060	0.0100	0.0163	0.0183	0.0266
college	0.0564 ***	0.0537 ***	0.0651 ***	0.1349 ***	0.1539 ***	0.0960 ***
	0.0051	0.0059	0.0076	0.0179	0.0201	0.0286
more than college	0.0574 ***	0.0596 ***	0.0512 ***	0.1109 ***	0.1283 ***	0.0781 **
	0.0047	0.0050	0.0088	0.0172	0.0193	0.0285
health very good	0.0158 **	0.0190 **	0.0048	0.0812 ***	0.0870 ***	0.0540 *
	0.0058	0.0063	0.0111	0.0126	0.0142	0.0211
health good	0.0366 ***	0.0344 ***	0.0385 ***	0.1855 ***	0.1883 ***	0.1536 ***
	0.0052	0.0059	0.0094	0.0123	0.0141	0.0200
health fair	0.0534 ***	0.0607 ***	0.0366 ***	0.2727 ***	0.2829 ***	0.2203 ***
	0.0047	0.0047	0.0096	0.0113	0.0141	0.0163
health poor	0.0660 ***	0.0680 ***	0.0583 ***	0.2967 ***	0.3076 ***	0.2384 ***
	0.0034	0.0041	0.0074	0.0111	0.0153	0.0152
cons	0.1689	0.4764	-7.1434	-1.1754	-1.1397	-6.5868
	0.3779	0.4103	0.4089	0.3500	0.3707	0.5011
Offer w/o covars	0.0135 *			0.0262 *		
	0.0056			0.0127		
Offer scaled by percent retired	0.0488			0.1176		
N	21482	15629	5258	21762	15779	5370

* p<0.05; ** p<0.01; *** p<0.001
Marginal effects from probit models. Standard errors are adjusted for the survey design and are clustered at the individual level. All survey questions for the health care utilization measures refer to use in the past two years.
Average percent retired: 24.6%. Rates of doctor visits: 93% total, 93% not retired, 94% retired. Rates of prescription drug use: 67% total, 64% not retired, 76% retired. Rates of outpatient surgery: 19% total, 18% not retired, 21% retired. Age and geographic region dummies are included in all models. Omitted categories are white, less than high school, excellent health, and 1994.

Table 1.7, continued
Estimated Effect of Retiree HI Offer on Health Care Utilization
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Any Outpatient Surgery		
	Total **g**	**Not Retired** **h**	**Retired** **i**
Offer	0.0205 *	0.0191 *	0.0156
	0.0080	0.0091	0.0169
female	0.0167 *	0.0126	0.0295 *
	0.0077	0.0091	0.0142
black	-0.0624 ***	-0.0548 ***	-0.0804 ***
	0.0099	0.0123	0.0170
hispanic	-0.0571 ***	-0.0523 ***	-0.0531
	0.0143	0.0161	0.0325
high school	0.0141	0.0208	0.0011
	0.0117	0.0140	0.0200
some college	0.0426 **	0.0504 **	0.0233
	0.0140	0.0168	0.0242
college	0.0659 ***	0.0799 ***	0.0382
	0.0179	0.0213	0.0344
more than college	0.0663 ***	0.0804 ***	0.0411
	0.0168	0.0201	0.0285
health very good	0.0089	0.0130	0.0041
	0.0100	0.0113	0.0200
health good	0.0470 ***	0.0446 ***	0.0463 *
	0.0112	0.0129	0.0214
health fair	0.0823 ***	0.0851 ***	0.0687 *
	0.0182	0.0220	0.0300
health poor	0.1529 ***	0.1602 ***	0.1279 **
	0.0288	0.0395	0.0404
cons	-0.9168	-1.0128	-0.5198
	0.3964	0.1949	0.7819
Offer w/o covars	0.0232 **		
	0.0081		
Offer scaled by percent retired	0.0833		
N	15497	10877	4219

* p<0.05; ** p<0.01; *** p<0.001
Marginal effects from probit models. Standard errors are adjusted for the survey design and are clustered at the individual level.
All survey questions for the health care utilization measures refer to use in th past two years. Average percent retired: 24.6%. Rates of doctor visits: 93% total, 93% not retired, 94% retired. Rates of prescription drug use: 67% total, 64% not retired, 76% retired. Rates of outpatient surgery: 19% total, 18% not retired, 21% retired.
Age and geographic region dummies are included in all models. Omitted categories are white, less than high school, excellent health, and 1994.

The estimated effects of RHI offer on any doctor's visit and prescription drug use are somewhat larger in magnitude among the retired subsample than the not retired group, but larger standard errors result in estimates that are only suggestive. Significant effects in the not retired group suggest there may be some selection into retirement based on preferences for outpatient surgery. It is also possible that those with an RHI offer also have more generous *current* employee health insurance benefits, allowing them to consume more health care services while they are still working. This fits with Card et al.'s (2004) finding that rates of elective surgeries increase more with Medicare eligibility for those with higher rates of health insurance coverage before 65, which they attribute this to the generosity of coverage. If this is the case, we cannot interpret the coefficient on RHI offer as "causal" per se, but the effect is likely driven by the combination of two different types of insurance coverage (ESI and RHI) as opposed to differences in individual characteristics.

5.5 Results: Differential Effects by Health Status

Though we see only suggestive effects on health and outpatient utilization rates, this may be an artifact of the majority of the sample being in fairly good health. To the extent that causal effects of RHI offer on health and health care use exist, we are more likely to see them in the subgroup in poor health. I use health shocks as described in section 4.1 as an exogenous source of variation in health status to test for differential effects of RHI offer.

For the health outcomes, the main effect of RHI offer remains zero while the main effects of the health shocks themselves worsen health (Table 1.8, columns a, b, and c). Conditional on experiencing a health shock, there is no significant evidence of a differential effect of RHI offer on health. Columns d, e, and f show that controlling for a health shock, RHI offer decreases the probability of being uninsured by about 2 percentage points, while the effects on ESI and public coverage are somewhat larger than in Table 1.6 (+8 and -6 percentage points, respectively). The health shocks themselves have no significant effect on health insurance coverage, nor is there an incremental effect of RHI offer conditional on a health shock. For the health care utilization outcomes, RHI offer has a positive effect on the probability of prescription drug use and outpatient surgery (columns g, h, and i). Since the survey questions on health care utilization refer to the 2 years prior to the interview, the results for the lagged health shocks are most informative and tend to show positive effects on outpatient utilization. In the two years following a chronic health shock, the probability of a doctor's visit increases by 3 percentage points and the probability of using prescription drugs increases by 7 percentage points.[41] These results suggest no incremental effect of RHI offer on health outcomes or health insurance coverage for those in poor health, but incremental effects for some utilization outcomes.

[41] The negative and very large coefficients on current acute health shocks interacted with RHI offer are likely driven by a few observations, given the small number of people who experience acute health shocks (N=256).

Table 1.8
Estimated Effect of Retiree HI Offer on Health, Health Insurance Coverage and Health Care Utilization: Health Shocks
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Health			**Health Insurance Coverage**		
	F/P Health **a**	**Change self-reported health** **b**	**Change ADLs** **c**	**Employer-Sponsored Coverage** **d**	**Public Coverage** **e**	**Uninsured** **f**
Chronic Health Shock						
Offer	-0.0119	-0.0003	-0.0050	0.0779 ***	-0.0549 ***	-0.0230 ***
	0.0090	0.0131	0.0070	0.0088	0.0076	0.0042
Chronic health shock	0.0463 **	0.2195 ***	0.0471	0.0104	-0.0054	-0.0051
	0.0166	0.0451	0.0249	0.0103	0.0086	0.0054
Offer*Shock	0.0224	0.0023	0.0253	-0.0046	-0.0009	0.0055
	0.0182	0.0545	0.0297	0.0150	0.0118	0.0090
Lagged chronic shock	0.0592 ***	-0.0845	-0.0097	-0.0061	0.0009	0.0052
	0.0169	0.0471	0.0268	0.0110	0.0091	0.0057
Offer* Lagged shock	0.0035	0.0618	-0.0084	0.0075	-0.0066	-0.0009
	0.0160	0.0563	0.0317	0.0130	0.0112	0.0063
N	11948	13310	13303	16666	16666	16666
Acute Health Shock						
Offer	-0.0081	0.0051	-0.0049	0.0778 ***	-0.0557 ***	-0.0221 ***
	0.0080	0.0112	0.0063	0.0080	0.0069	0.0039
Acute health shock	0.2626 ***	0.7464 ***	0.2146	0.0163	-0.0134	-0.0028
	0.0554	0.1374	0.1144	0.0191	0.0143	0.0116
Offer*Shock	-0.0079	-0.1097	-0.1269	0.0059	0.0016	-0.0075
	0.0281	0.1590	0.1222	0.0284	0.0247	0.0131
Lagged acute shock	0.1908 ***	-0.2136	-0.0391	0.0017	-0.0042	0.0025
	0.0535	0.1215	0.1242	0.0240	0.0191	0.0130
Offer* Lagged shock	-0.0100	-0.0215	0.0269	-0.0198	0.0060	0.0138
	0.0316	0.1501	0.1301	0.0378	0.0300	0.0220
N	11818	13129	13124	16403	16403	16403

* p<0.05; ** p<0.01; *** p<0.001
Columns a, g, h, and i use probit models. The estimates for health insurance coverage are marginal effects from a multinomial logit model.
Standard errors are adjusted for the survey design and are clustered at the individual level.
Fair/poor health regression is conditional on not being in fair/poor health at baseline.
Health controls: sex, age, race, education, baseline health status, and year.
Health insurance coverage controls: sex, age, race, education, marital status, baseline health status, and year.
Health care utilization controls: sex, age, race, education, baseline health status, geographic region, and year.

Table 1.8, continued
Estimated Effect of Retiree HI Offer on Health, Health Insurance Coverage and Health Care Utilization: Health Shocks
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Health Care Utilization		
	Doctor's Visits	**Rx Use**	**Outpatient Surgery**
	g	**h**	**i**
Chronic Health Shock			
Offer	0.0046	0.0229	0.0203 *
	0.0070	0.0151	0.0095
Chronic health shock	0.0392 ***	0.1350 ***	-0.0178
	0.0093	0.0194	0.0197
Offer*Shock	0.0176	0.0415	0.0402
	0.0155	0.0281	0.0275
Lagged chronic shock	0.0082	0.0817 ***	0.0359
	0.0116	0.0220	0.0214
Offer* Lagged shock	0.0331 **	0.0671 *	0.0061
	0.0110	0.0273	0.0239
N	13126	13297	13290
Acute Health Shock			
Offer	0.0092	0.0334 *	0.0285 ***
	0.0067	0.0145	0.0086
Acute health shock	0.0805 ***	0.2177 ***	0.2646 ***
	0.0032	0.0358	0.0576
Offer*Shock	-0.9308 ***	-0.1699 *	-0.0888 **
	0.0025	0.0832	0.0291
Lagged acute shock	0.0616 ***	0.1820 ***	0.0777
	0.0081	0.0424	0.0538
Offer* Lagged shock	-0.0461	-0.0674	-0.0536
	0.1034	0.0865	0.0440
N	12954	13117	13111

* p<0.05; ** p<0.01; *** p<0.001
Columns a, g, h, and i use probit models. The estimates for health insurance coverage are marginal effects from a multinomial logit model.
Standard errors are adjusted for the survey design and are clustered at the individual level. Fair/poor health regression is conditional on not being in fair/poor health at baseline. Health controls: sex, age, race, education, baseline health status, and year. Health insurance coverage controls: sex, age, race, education, marital status, baseline health status, and year. Health care utilization controls: sex, age, race, education, baseline health status, geographic region, and year.

Section 6: Robustness

6.1 Robustness specifications

Additional analyses presented here test the identifying assumption that RHI offer is conditionally random. I present results that are based on a subsample of respondents with baseline job tenure at or greater than the median (12 years), since respondents who report current job tenure between 12 and 48 years at ages 47-63 are unlikely to have moved into jobs with an RHI offer as they neared retirement age. I also present results based on a subsample of individuals who are at least four years away from retirement when they report their RHI offer status, since those who are close to retirement may be more aware of their retiree benefits.

To address concerns about other control variables that differ across the groups offered and not offered RHI (i.e., marital status), I also rerun the main analysis using propensity score weighting which balances observed characteristics very closely between groups (McWilliams, et al. 2003). This approach is based on treatment assignment being unconfounded with potential outcomes conditional on a set of observed covariates (Rosenbaum and Rubin 1983). I calculate the propensity score with a probit model to predict RHI offer using all of the covariates in Table 1.2.[42] Individual weights equal to the probability of belonging to the opposite RHI offer group are calculated based on the propensity score. After adjustment using the new weights, all observed covariates are balanced by RHI offer status.

6.2 Robustness results

[42] Current job tenure, industry and occupation are not included in the propensity score model since they are only observed for those working at baseline. Similarly, parents' age at death is not included since it is only observed for a subsample of respondents.

Row 2 of Table 1.9 shows that the estimated effects of RHI offer based on the subsample of respondents with longer job tenure are similar to the baseline specification. The effects of RHI offer on retirement are slightly larger (9 percentage points) as are the effects on health outcomes, with the change in the number of ADLs performed with difficulty declining by 0.016 for those with an RHI offer, compared to a mean increase of 0.025. The estimated effects on health insurance coverage and health care utilization are close to the baseline results.

Table 1.9
Estimated Effect of Retiree HI Offer on Full-Time Retirement, Health, Health Insurance Coverage and Health Care Utilization: Alternative Specifications
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Full-Time Retirement			Health		
	Total	Men	Women	F/P Health	Change self-reported health	Change ADLs
1. baseline specification						
Offer	0.0698 ***	0.0653 ***	0.0730 ***	-0.0083	0.0001	-0.0030
	0.0102	0.0144	0.0133	0.0068	0.0079	0.0049
N	13,747	6,486	7,261	18,870	21,504	21,496
2. tenure>=median						
Offer	0.0900 ***	0.0883 ***	0.0854 ***	-0.0155	0.0039	-0.0164 *
	0.0152	0.0209	0.0219	0.0110	0.0134	0.0083
N	7,593	4,030	3,563	7,871	8,665	8,665
mean of outcome	0.250	0.253	0.246	0.092	0.084	0.025
3. not retired waves 1-3						
Offer	0.0477 ***	0.0454 *	0.0504 **	-0.0085	0.0320 *	-0.0105
	0.0145	0.0210	0.0188	0.0103	0.0151	0.0093
N	5,584	2,540	3,044	7,230	7,785	7,786
mean of outcome	0.194	0.214	0.177	0.110	0.107	0.015
4. propensity score weighted						
Offer	0.0696 ***	0.0657 ***	0.0743 ***	-0.0084	0.0008	-0.0003
	0.0109	0.0152	0.0138	0.0068	0.0080	0.0051
N	13,406	6,269	7,137	18,494	21,022	21,014

* p<0.05; ** p<0.01; *** p<0.001
Probit models are used for retirement, fair/poor health and utilization outcomes. The estimates for health insurance coverage are marginal effects from a multinomial logit model.
Standard errors are adjusted for the survey design and are clustered at the individual level.
Full-time retirement regressions are conditional on not being full-time retired at baseline.
Fair/poor health regression is conditional on not being in fair/poor health at baseline.
Full-time retirement controls: sex, age, race, education, marital status, baseline health status, pension, industry, occupation, household income and assets, and year.
Health controls: sex, age, race, education, baseline health status, and year.
Health insurance coverage controls: sex, age, race, education, marital status, baseline health status, and year.
Health care utilization controls: sex, age, race, education, baseline health status, geographic region, and year.
Median tenure at baseline is 12.3 years. Models in row 3 are restricted to individuals who are not full-time retired in waves 1-3 and to years 1998, 2000, and 2002.
The propensity score is calculated using a probit model to predict RHI offer based the following control variables (at baseline): age, sex, race, education, household income, marital status, census region, labor force participation, self-reported health, whether health limits work, DI application, DI receipt, any chronic illness, any acute illness, and parents' mortality.

Table 1.9, continued
Estimated Effect of Retiree HI Offer on Full-Time Retirement, Health, Health Insurance Coverage and Health Care Utilization: Alternative Specifications
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Health Insurance Coverage			Health Care Utilization		
	Employer-Sponsored Coverage	**Public Coverage**	**Uninsured**	**Doctor's Visits**	**Rx Use**	**Outpatient Surgery**
1. baseline specification						
Offer	0.0516 ***	-0.0240 ***	-0.0277 ***	0.0120 *	0.0289 *	0.0205 *
	0.0060	0.0044	0.0041	0.0056	0.0120	0.0080
N	20,760	20,760	20,760	21,482	21,762	15,497
2. tenure>=median						
Offer	0.0400 ***	-0.0234 ***	-0.0165 ***	0.0100	0.0131	0.0319 *
	0.0082	0.0062	0.0050	0.0091	0.0192	0.0126
N	8,389	8,389	8,389	8,675	8,773	6,272
mean of outcome	0.918	0.025	0.039	0.934	0.673	0.185
3. not retired waves 1-3						
Offer	0.0431 ***	-0.0292 ***	-0.0139 **	0.0065	0.0105	0.0140
	0.0092	0.0079	0.0048	0.0072	0.0157	0.0106
N	7,425	7,425	7,425	7,796	7,901	7,898
mean of outcome	0.871	0.047	0.047	0.939	0.722	0.184
4. propensity score weighted						
Offer	0.0468 ***	-0.0184 ***	-0.0285 ***	0.0096	0.0198	0.0210 *
	0.0054	0.0033	0.0042	0.0056	0.0123	0.0083
N	20,340	20,340	20,340	20,995	21,270	15,156

* p<0.05; ** p<0.01; *** p<0.001
Probit models are used for retirement, fair/poor health and utilization outcomes. The estimates for health insurance coverage are marginal effects from a multinomial logit model.
Standard errors are adjusted for the survey design and are clustered at the individual level.
Full-time retirement regressions are conditional on not being full-time retired at baseline.
Fair/poor health regression is conditional on not being in fair/poor health at baseline.
Full-time retirement controls: sex, age, race, education, marital status, baseline health status, pension, industry, occupation, household income and assets, and year.
Health controls: sex, age, race, education, baseline health status, and year.
Health insurance coverage controls: sex, age, race, education, marital status, baseline health status, and year.
Health care utilization controls: sex, age, race, education, baseline health status, geographic region, and year.
Median tenure at baseline is 12.3 years. Models in row 3 are restricted to individuals who are not full-time retired in waves 1-3 and to years 1998, 2000, and 2002.
The propensity score is calculated using a probit model to predict RHI offer based the following control variables (at baseline): age, sex, race, education, household income, marital status, census region, labor force participation, self-reported health, whether health limits work, DI application, DI receipt, any chronic illness, any acute illness, and parents' mortality.

Row 3 shows the estimates based on the subsample of respondents who are not retired full-time in 1992-1996. Although the magnitudes of the estimated effects on retirement are smaller than the baseline results, a 25 percent increase in the probability of early retirement is still substantial. The estimated effects of offer on health insurance coverage are very similar to the baseline results, as are those for two of the three health outcomes. The estimated effects on health care utilization are somewhat smaller and no longer statistically significant.

Row 4 shows the results using propensity score weighting. The estimated effects all match the baseline results very closely, with the exception of the statistical significance of the effects on any doctor's visit and prescription drug use. This suggests that the baseline estimates suffer from minimal bias due to unbalanced covariates across the two groups.

These robustness checks strengthen the conclusions from the main analysis that an RHI offer significantly increases the probability of early retirement and employer-sponsored health insurance coverage. It decreases the probability of public health insurance coverage and of being uninsured, and has no statistically significant impact on health status. The original finding of a positive effect on outpatient health care utilization is somewhat weaker.

Section 7: Risk Protection and Out-of-Pocket Medical Spending

7.1 Econometric Specification

Next I investigate whether those without RHI spend more out-of-pocket for medical care services, in essence experiencing more financial loss to protect their health.

Out-of-pocket medical expenses include amounts paid for care related to hospitals, nursing homes, doctors, dentists, outpatient surgery, prescription drugs, home health care, and special facilities. Figure 1.3 shows the distributions of total out-of-pocket medical spending by RHI offer and Figure 1.4 shows these distributions for retirees. While the spending distributions for the total sample appear to track each other closely, the mean difference in the top 10 percent is $ 786 and is $1,243 in the top 5 percent. Among retirees, the differences are much larger and clearly visible in the graph. The average difference is $3,833 in the top 10 percent and $5,516 in the top 5 percent.

Figure 1.3

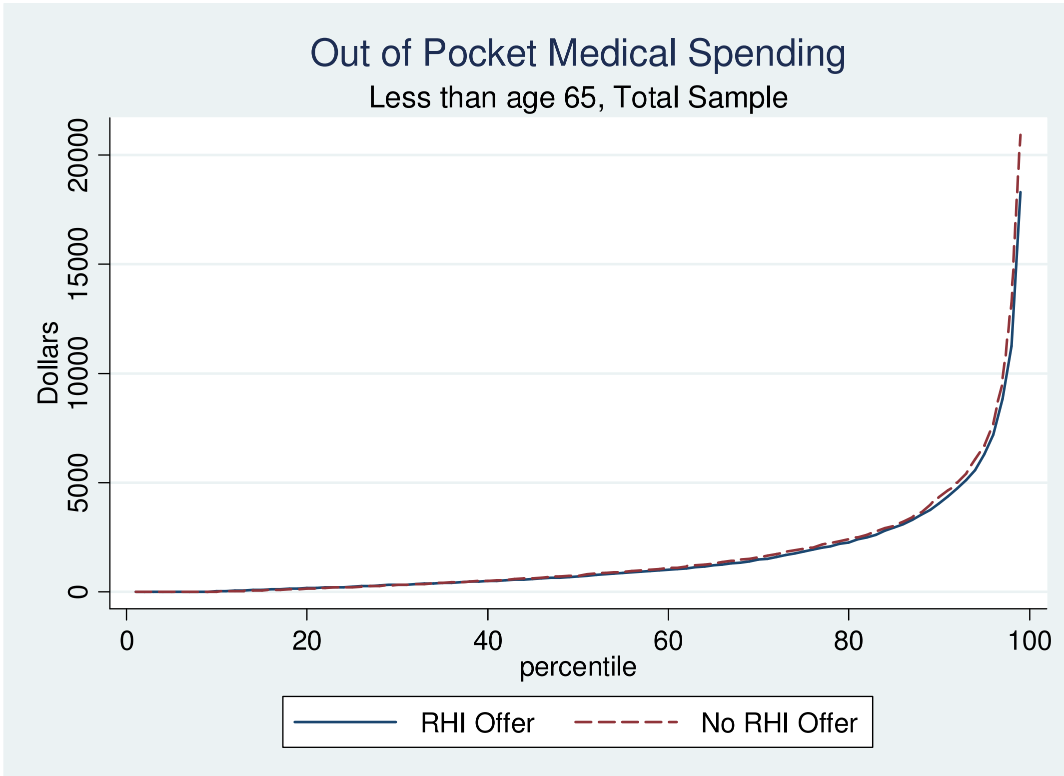
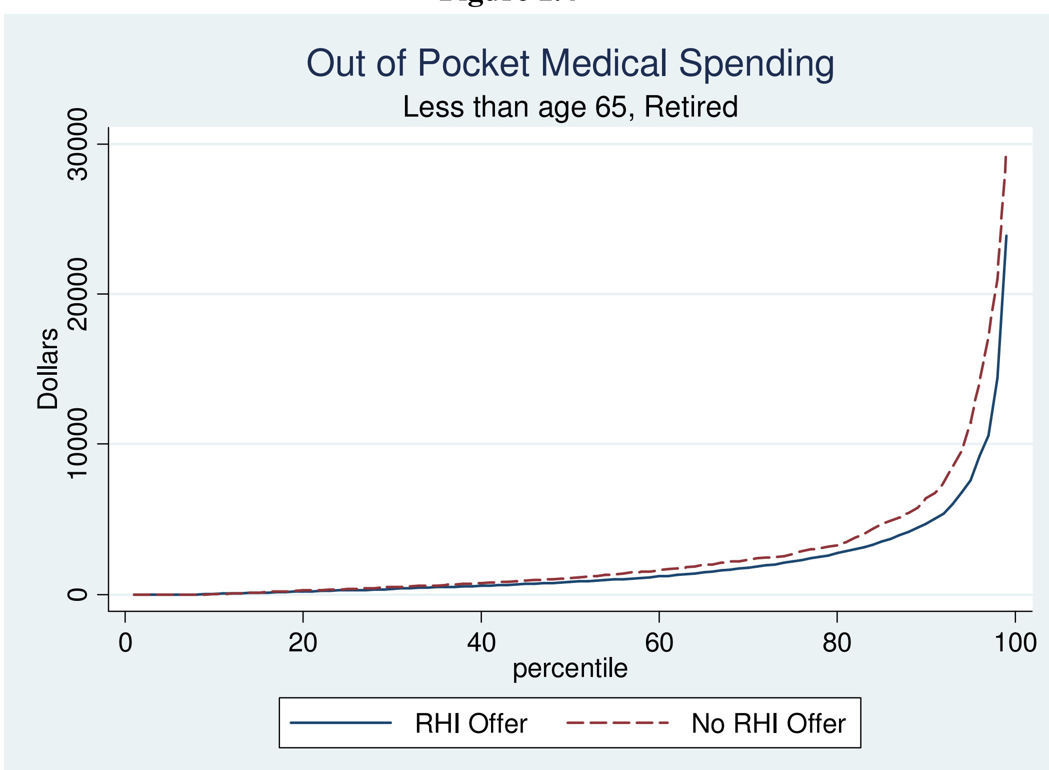


Figure 1.4



Because of the skewed distribution of medical expenditures, it is useful to examine the effects of RHI offer throughout the distribution, instead of just the effect at the mean or median (Finkelstein and McKnight 2005). I calculate the residual out-of-pocket spending after controlling for individual demographics (age, sex, race/ethnicity, education, and baseline health status)

$$\text{OOP Spending}_{it} = \alpha + X_{it}\,\beta_1 + \varepsilon_{it} \qquad (8)$$

and then calculate the difference in residual spending between those with and without a RHI offer at each percentile *p* of the distribution.

$$\Delta_p = \{\text{spend}_p(\text{offer} = 1) - \text{spend}_p(\text{offer} = 0)\} \qquad (9)$$

7.2 Results: Out-of-Pocket Medical Spending

Figure 1.5 shows this centile treatment effect and its 95% confidence interval, calculated using the empirical standard deviation of 200 bootstrap replications of the centile treatment estimates. Until the 60^{th} percentile, the difference between the spending residuals is not statistically significant. In the top 40 percent of the residual spending distribution, however we can see an negative, increasing (in absolute value) and statistically significant effect of RHI offer on out-of-pocket medical spending. In the top 40 percent of the distribution, those with an RHI offer spend about $275 (6 percent) less per year out-of-pocket for medical care than their counterparts without an RHI offer (Table 1.10).

Figure 1.5

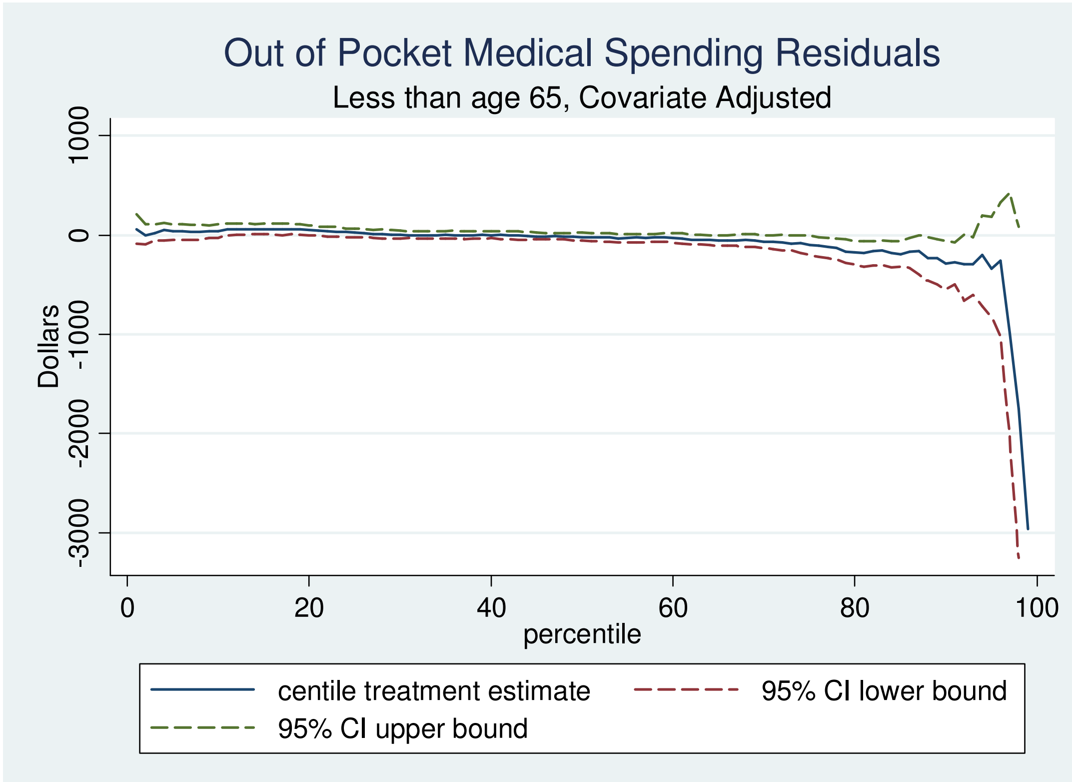

Out of Pocket Medical Spending Residuals
Less than age 65, Covariate Adjusted
Dollars
1000
0
-1000
-2000
-3000
0
20
40
60
80
100
percentile
centile treatment estimate
95% CI lower bound
95% CI upper bound

Table 1.10
Estimated Effect of Retiree HI Offer on Out-of-Pocket Medical Spending
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Total		Not Retired		Retired	
Centile	Average Out-of-Pocket Medical Spending	Centile Treatment Estimate	Average Out-of-Pocket Medical Spending	Centile Treatment Estimate	Average Out-of-Pocket Medical Spending	Centile Treatment Estimate
60	$1,000	-$28	$960	-$54 *	$1,260	-$131
61	$1,046	-$34	$1,000	-$54 *	$1,320	-$146
62	$1,098	-$48	$1,000	-$45	$1,380	-$163
63	$1,130	-$53	$1,040	-$47	$1,440	-$196
64	$1,180	-$49 *	$1,081	-$47	$1,488	-$245
65	$1,220	-$58 *	$1,120	-$49	$1,524	-$236 *
66	$1,272	-$55 *	$1,168	-$57	$1,620	-$245 *
67	$1,320	-$54	$1,212	-$64	$1,680	-$231 *
68	$1,387	-$47	$1,248	-$62	$1,740	-$256 *
69	$1,450	-$57	$1,300	-$57	$1,820	-$260 *
70	$1,500	-$66 *	$1,360	-$45	$1,917	-$229 *
71	$1,560	-$69 *	$1,430	-$57	$1,980	-$204 *
72	$1,640	-$77	$1,500	-$64	$2,025	-$223 *
73	$1,720	-$86 *	$1,536	-$89	$2,144	-$226
74	$1,788	-$80 *	$1,620	-$96	$2,210	-$246
75	$1,870	-$101 *	$1,710	-$99 *	$2,320	-$288
76	$1,970	-$111 *	$1,768	-$112 *	$2,404	-$308
77	$2,000	-$121 *	$1,860	-$97 *	$2,500	-$370
78	$2,120	-$128 *	$1,969	-$101 *	$2,600	-$394 *
79	$2,216	-$166 *	$2,000	-$138	$2,728	-$391 *
80	$2,312	-$172 *	$2,120	-$134 *	$2,880	-$457 *
81	$2,440	-$184 *	$2,220	-$139 *	$3,000	-$519 *
82	$2,527	-$159 *	$2,312	-$139	$3,138	-$591 *
83	$2,672	-$156 *	$2,450	-$131 *	$3,245	-$736 *
84	$2,800	-$178 *	$2,550	-$120	$3,480	-$769 *
85	$2,980	-$198 *	$2,700	-$125	$3,680	-$854 *
86	$3,100	-$171 *	$2,840	-$126	$3,920	-$874 *
87	$3,340	-$161 *	$3,000	-$81	$4,120	-$1,042 *
88	$3,573	-$232 *	$3,200	-$123	$4,400	-$1,025 *
89	$3,809	-$234 *	$3,478	-$142	$4,680	-$1,231 *
90	$4,100	-$289 *	$3,700	-$247 *	$5,000	-$1,494 *
91	$4,440	-$272 *	$4,000	-$264	$5,300	-$1,632 *
92	$4,824	-$297	$4,400	-$198	$5,860	-$1,500 *
93	$5,200	-$294 *	$4,800	-$156	$6,500	-$1,736 *
94	$5,760	-$202	$5,222	$10	$7,200	-$2,347 *
95	$6,449	-$338	$5,860	-$141	$8,400	-$3,633 *
96	$7,350	-$263	$6,600	-$219	$9,700	-$5,371 *
97	$9,117	-$985	$7,735	-$194	$12,000	-$5,787 *
98	$12,000	-$1,752	$10,200	-$781	$16,058	-$7,936 *
99	$19,280	-$2,966	$15,915	$84	$24,780	-$8,505

* $p<0.05$
Covariate adjusted (age, sex, race, education, and baseline health status).

Figures 1.6 and 1.7 show the results for the not retired and retired groups, respectively. The treatment estimates are generally not significant for the not retired group, so without evidence of significant selection effects we turn to the centile treatment estimates among retirees. In the top 40 percent of the residual spending distribution, retirees with an RHI offer spend more than $1,300 less per year on average than those without an RHI offer, which is 21 percent of average spending.

Figure 1.6

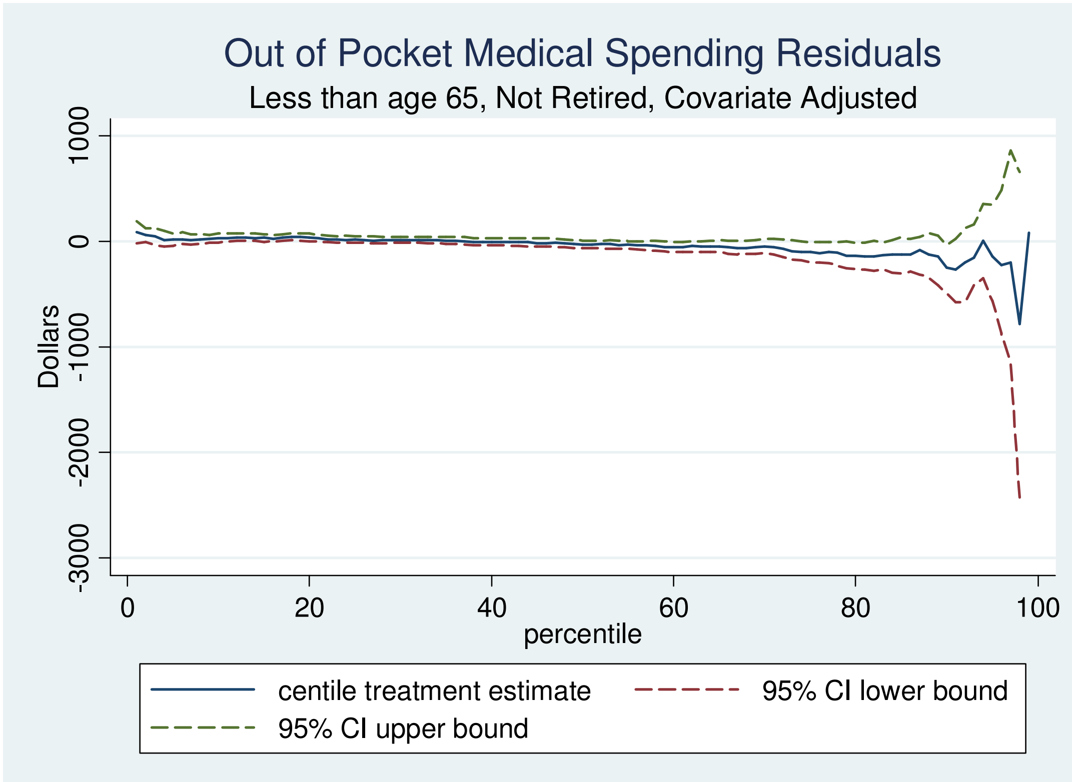
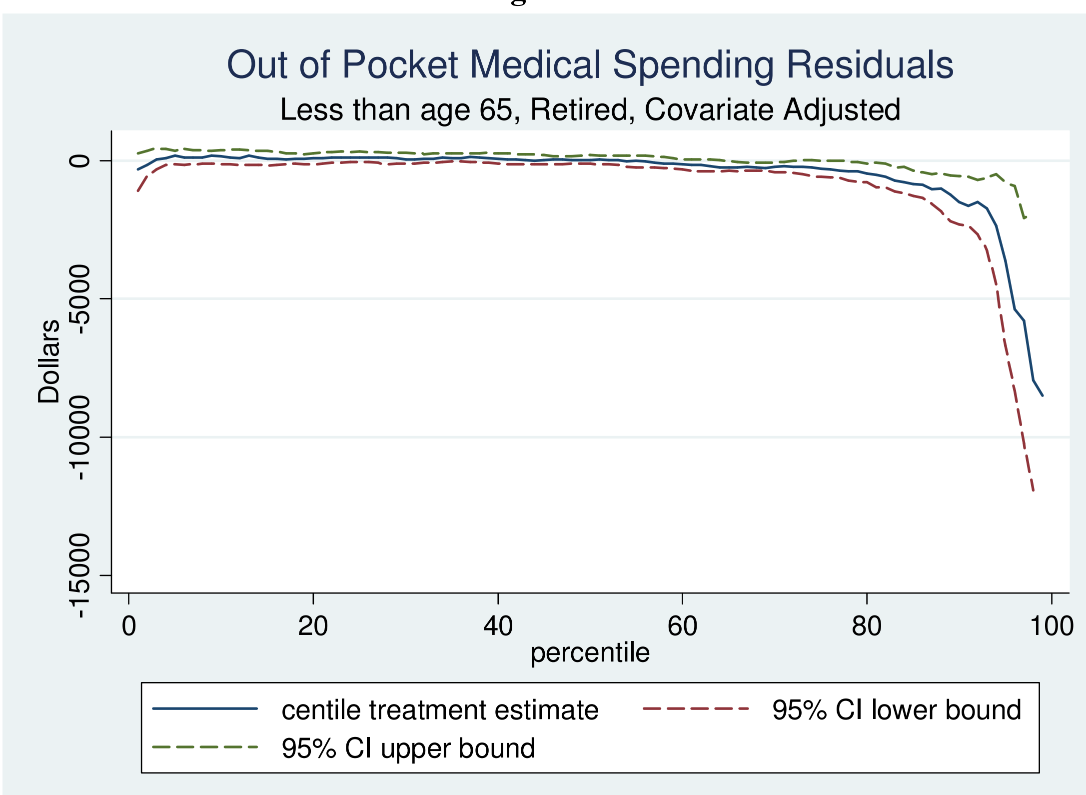


Figure 1.7



These quantitatively large and statistically significant effects suggest that RHI offers significant protection from high out-of-pocket medical costs. Even though respondents with an RHI offer use more of some types of health care services on average, they spend less in the top 40 percent of the out-of-pocket spending distribution. These results also suggest that part of the reason that we do not observe a significant effect of RHI offer on health is because those without RHI spend more out-of-pocket to get the medical care that they need. Given that we do not see significant effects of RHI on spending for non-retirees, the effects we see for retirees are likely due to the causal effects of insurance rather than selection.

7.3 The value of retiree health insurance

In order to understand the value of this risk protection to individuals, I conduct a utility analysis using the empirical distributions of out-of-pocket spending among retirees with and without an RHI offer. I examine men and women separately and focus on the 60-64 age group, since we can expect less selection into retirement with respect to health and medical spending than for younger individuals. For each individual, I take 50 random draws from the out-of-pocket medical spending distribution of those offered RHI. I repeat this process using the distribution for the not offered group.[43] Expected utility is given by:

$$EU_{igo} = 1/50\, \Sigma\, u\, (y_{ig} - m_{go}) \qquad \text{and} \qquad EU_{ign} = 1/50\, \Sigma\, u\, (y_{ig} - m_{gn})$$

[43] To avoid draws from the spending distribution that are very high (or even higher than) income, I cap out-of-pocket spending at 90 percent of income. Capping at 80 or 95 percent of income generates lower (higher) willingness-to-pay estimates, but they are of the same order of magnitude as those presented here. Many of the capped draws are likely to be legitimate spending amounts (I do not allow for savings or borrowing in my model, which many people will draw on to pay medical bills). The results presented here are therefore likely to be an underestimate of the willingness to pay for retiree health insurance.

where y_{ig} is household income for individual *i* in age/sex group *g*, m_{go} is out-of-pocket medical spending drawn from the offered distribution for group *g*, and m_{gn} is out-of-pocket spending drawn from the not offered distribution for group *g*. I calculate each individual's certainty equivalent under the offered and not offered spending distributions. The certainty equivalent is the amount of money that makes the individual as well of as facing the risk of out-of-pocket medical expenditures.

$$u(CE_{igo}) = 1/50 \, \Sigma \, u(y_{ig} - m_{go}) \qquad \text{and} \qquad u(CE_{ign}) = 1/50 \, \Sigma \, u(y_{ig} - m_{gn})$$

The risk premium, or willingness to pay, for full insurance under each risk distribution is the difference between the individual's household income and their certainty equivalent. The risk premium for retiree health insurance is therefore the difference in risk premia between two distributions. I use a constant relative risk aversion utility function with a preferred coefficient of relative risk aversion equal to 3. Due to the inherent uncertainty of the value of this parameter, I also present a range of results using coefficients of risk aversion of 1 and 5.[44]

[44] The results are substantively similar using a constant absolute risk aversion utility function and a central risk-aversion parameter of 0.00021 (Manning and Marquis, 1996).

Table 1.11 provides the mean certainty equivalent and mean risk premia across individuals. Retired men ages 60-64 are willing to pay $8,929 per year for full insurance against the risk in the not offered distribution and $5,101 for the offered distribution. This suggests that the value of RHI coverage for this subgroup is $3,828. Retired women ages 60-64 are willing to pay $3,797 for insurance against the differential risk between the two distributions. After accounting for the fact that those with an RHI offer spend less on average ($465 for the men and $673 for the women), men's value of the risk reduction associated with RHI is about $3,400 per year and women's is about $3,100 per year.[45]

[45] An alternative procedure that averages household income and out-of-pocket spending draws across years within individuals yields similar, though slightly smaller, estimates: $2,400 for men and $2,100 for women.

Table 1.11
Estimates of Willingness-to-Pay for Retiree Health Insurance Coverage
Employer-Sponsored Health Insurance in 1992
Less Than Age 65

	Retired Men 60-64			Retired Women 60-64		
Coefficient of relative risk aversion	1	3	5	1	3	5
Certainty Equivalents						
Not offered mean	$49,796	$44,008	$38,854	$43,726	$37,741	$32,713
Offered mean	50,742	47,836	44,844	45,011	41,539	38,221
Willingness-to-Pay						
Not offered mean	3,141	8,929	14,083	3,826	9,810	14,839
Offered mean	2,195	5,101	8,093	2,540	6,013	9,331
WTP difference	946	3,828	5,990	1,285	3,797	5,508
N		1911			1894	
Spending Distributions						
Capped	**Mean**	**SE**		**Mean**	**SE**	
Not Offered	2,288	4,674		2,762	5,008	
Offered	1,824	4,020		2,089	4,051	
Difference	465			673		
Total						
Not Offered	5,920	63,861		3,790	14,646	
Offered	1,953	4,677		2,357	6,920	
Difference	3,967			1,433		

Spending draws are capped at 90 percent of income.
The certainty equivalent and willingness-to-pay estimates are based on the capped spending distribution. Information on the total (uncapped) distribution is provided only for comparison.

Given the capping procedure, these estimates should be treated as conservative estimates of the value of retiree health insurance. The small probability of catastrophic medical expenses is exactly the type of risk that health insurance is designed to prevent against, and these very large expenses are effectively eliminated from the capped distributions.[46]

These sizeable estimates of the willingness-to-pay for retiree health insurance suggest that policies aimed at correcting market imperfections in the individual health insurance market could increase coverage rates for the near-elderly. Approximately 12 percent of retirees aged 55-64 currently purchase insurance coverage on their own (United States General Accounting Office 2001b). Average premiums for individuals aged 50-64 are $3,300 and are nearly $4,000 for those aged 60-64.[47] Loading fees in the individual insurance market are approximately 40 percent (Pauly, Percy, Herring 1999). My willingness-to-pay estimates for 60-64 year old retirees suggest that decreasing loading fees in the individual market to 20 percent would bring the average premium in line with the average willingness-to-pay, which could substantially increase take-up rates by the near-elderly. Policies that enable individuals to buy into public insurance programs at actuarially fair rates could provide another venue for the near-elderly to obtain coverage.

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[46] For men, the difference in mean spending is nearly 9 times as large in the total distribution as in the capped distribution, and the standard error of the not offered spending distribution is more than an order of magnitude larger in the total distribution than in the capped. For women, the differences are not as large but are still substantial.

[47] "Average" premiums are difficult to calculate, due to the wide variety of deductibles, upper payment limits, exclusions and coinsurance policies for individual market health insurance policies as well as varying state regulations. These figures come from averaging results from a number of studies that calculate average premiums for actual individual market policies or for quotes for standardized individuals and plan types (America's Health Insurance Plans 2005, Collins, et al. 2006, Gabel, et al. 2002, Hadley and Reschovsky 2002, Musco and Wildsmith 2002, Simantov, et al. 2001, U.S. GAO 1998).

Section 8: Conclusions and Policy Implications

This paper examines the effects of retiree health insurance offer on labor force and health outcomes in the near-elderly population. The results imply that the large impact of RHI offer on retirement is unlikely to be driven by actual health consequences. That is, while those facing early retirement without an RHI offer may not expect an immediate negative impact on health, the results suggest that they experience significant risk aversion with respect to health shocks and high out-of-pocket expenditures. Early retirees without an RHI offer may manage to protect their health, but may spend a lot more to do so. Though the health and medical expenditure risks are relatively small, they are important enough for this age group to result in a relatively large willingness-to-pay for retiree health insurance.

I find that RHI offer increases the probability of early retirement by about 40 percent for both men and women. This effect is robust to adding important covariates including individual and spousal demographics, job characteristics, income and pension information, as well as controlling for a health shock. There is no evidence of significant differential early retirement based on health status.

The analysis suggests that RHI offer has little, if any, effect on health, either among the entire near-elderly population or among the group most likely to be affected – individuals who experienced health shocks. RHI offer decreases the probability of being uninsured by 55 percent, while increasing the probability of employer-sponsored coverage and decreasing the probability of public coverage. Considering labor force participation and health insurance coverage jointly, those with an RHI offer are significantly more likely to have employer-sponsored coverage in retirement and those without an offer are more likely to maintain their employer-sponsored coverage by continuing to work full-time. There is suggestive evidence that individuals with an RHI offer are more likely to visit the doctor, use prescription drugs on a regular basis and have outpatient surgery. There is no evidence of differences in inpatient or preventive care utilization.

The estimates of the effect of RHI offer on out-of-pocket medical spending are a key part of the analysis. Those with an RHI offer are better protected from high out-of-pocket medical costs, on average spending $1,300 less in the top 40 percent of the spending distribution (21 percent). This suggests that part of the reason we see no significant effect of RHI offer on health may be because the near-elderly without retiree coverage spend more out-of-pocket to get the care they need. This also highlights the role of health insurance in protection from high medical expenditures and that the effects of health insurance on both health and financial risk protection need to be examined to understand the full scope of effects. The utility analysis suggests that men ages 60-64 are willing to pay approximately $3,400 for the risk protection that RHI provides, and women in the same age group are willing to pay about $3,100.

In order to quantify the implications of the decline in employer RHI offer rates, I calculate a weighted average of the decline across different firm size categories.[48] This average decline in offer rates of 43 percent suggests a 15 percent decline in early retirement rates, assuming no other changes. Labor force participation rates among men aged 55-64 have declined by 19 percent since 1960 and as the cohorts of workers affected

[48] I use data on the decline in RHI offer rates by firm size from the Kaiser/HRET Employer Health Benefits Annual Survey, 1988-2005 and calculate the share of adult employment (ages 20-50) from the 2001 Current Population Survey.

by the decline in RHI offer rates age, as much as 80 percent of this decline in labor force participation could be reversed. Among women in this age group, labor force participation rates have increased by 50 percent, and this change could be magnified by as much as 30 percent.[49]

The change in early retirement rates will have an uncertain effect on the efficiency of retirement patterns, depending on how productive older workers are, whether employers can adjust on other margins such as wages, and how employers use other mechanisms (i.e., pension design) to induce efficient retirement in the face of rising wage profiles. From the perspective of individual workers, my results suggest that the decline in RHI offer rates may keep some individuals at work beyond the point where their marginal utility of leisure time exceeds their marginal productivity. The decline in early retirement rates may have some positive effects from the standpoint of social efficiency however, as individuals continue to pay into the Social Security system, rather than draw from it. Future research that addresses these issues will be valuable in understanding the normative implications of the effects of the decline in RHI offer rates on early retirement.

My results also suggest that the decline in RHI offer rates will result in significant changes in insurance coverage and financial risk protection for the near elderly. The 43 percent decline in offer rates implies a 29 percent increase in the probability of being uninsured and a 5 percent increase in the probability of being in the top 35 percent of the out-of-pocket spending distribution, ceterius paribus. The sizeable willingness-to-pay estimates for RHI coverage suggest that much of the projected increase in uninsured rates

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[49] Labor force participation rates are from the Statistical Abstract of the United States, 1980, 2000, and 2006. The labor force participation rate among men aged 55-64 declined from 85 percent in 1960 to 69 percent in 2004. For women aged 55-64, rates increased from 37 percent in 1960 to 56 percent in 2004.

among the near-elderly could be averted if loading fees for individual market policies were reduced and prices were closer to actuarially fair. Decreasing loading fees by half could bring average individual market premiums to the level of average willingness-to-pay for the near-elderly, which could have a significant impact on take-up rates of individual insurance. Policies that stabilize employer-based group coverage, address imperfections in the individual insurance market and deal with the insurance needs of the near-elderly through public insurance programs will be important as the impacts of the decline in RHI offer rates become manifest in the coming years.

Medicaid's Effect on Single Women's Labor Supply: Evidence from the Introduction of Medicaid*

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Section 1: Introduction

The Medicaid program has provided health insurance coverage to low-income adults and children in the U.S. for over 40 years. Eligibility expansions to the program are currently a central strategy to increase rates of health insurance coverage, even while the program is a major and growing cost component of Federal and state budgets. As policymakers try to balance the costs and benefits of Medicaid, it is worthwhile to note that the direct costs for young women and their children are relatively low.[1] There may be indirect costs, however, through distortions to labor supply and family structure, increases in caseloads for cash assistance programs, and crowd-out of private health insurance. While the relevant incentives driving the first two of these distortions have been weakened by the decoupling of Medicaid and cash assistance in the 1980's, concern remains about the disincentives associated with providing health insurance coverage to the poor (Cannon, 2005).

This paper examines the impact of the introduction of the Medicaid program on labor supply decisions among single women in the late 1960's and early 1970's. Medicaid was closely tied to welfare receipt at this time, which generates a clear theoretical prediction of a negative labor supply response based on a budget constraint analysis. However, if health insurance coverage has positive effects on health for

[1] In 2003, adults (primarily poor working parents) and children comprised 75 percent of Medicaid beneficiaries and 30 percent of Medicaid spending (approximately $70 billion). Per capita spending was approximately $1,410 per child and $1,799 per non-disabled adult (Kaiser Commission on Medicaid and the Uninsured, 2007).

mothers or their children, there could be positive impacts on labor supply that might outweigh the negative incentives in the budget constraint.[2]

Little is known about the effects of this major public health insurance program on labor supply at its inception.[3] While other research has used Medicaid expansions in the 1980's or measures of the value of Medicaid to estimate the effect on labor supply, this analysis looks specifically at the implementation of the Medicaid program. Using a different natural experiment to identify the effect of Medicaid contributes to the generalizability of the existing results. I use a difference-in-differences-in-differences methodology to estimate the effect of Medicaid on eligible women's labor force participation, taking advantage of variation in implementation timing across states that is plausibly exogenous to individual's labor supply decisions. I compare my estimates to the effects predicted by economic theory and those found by previous empirical research.

The Medicaid program has changed significantly since 1965, by raising the income limits for eligibility, modifying the family structure requirements, and severing the tie with welfare receipt. Still, a better understanding of how the original eligible population responded to the associated labor supply incentives may enhance our understanding of how low-income families value health insurance coverage. As many children and adults currently remain eligible but not enrolled in Medicaid, this knowledge could improve the design of contemporary programs to increase coverage rates in this population.

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[2] Decker (1993) and Currie and Gruber (1996a and 1996b) find negative effects of Medicaid coverage on rates of low-birthweight among low-income infants and on infant and child mortality, as well as positive effects on health care utilization among children.

[3] Decker (1993) focuses primarily on the effect of Medicaid implementation on AFDC enrollment, but also fails to find significant effects on labor force participation.

Section 2: The Medicaid Program at Implementation

Title XIX, the Federal law authorizing the Medicaid program, was passed in 1965 and states implemented the program between 1966 and 1970. This joint Federal-State program provided health insurance to the poor, a benefit not provided by any pre-existing government programs and that low-wage, low-skill jobs were unlikely to offer. Medicaid is administered by the states and state expenditures were matched by Federal dollars at the rate of 50-83%, which varied inversely with state per capita income. In order to qualify for these matching funds, states were required to provide a basic benefits package to all beneficiaries (including inpatient and outpatient hospital services, laboratory and X-ray services, and physician services) with the option to provide more generous benefits. Until 1987, eligibility was very closely tied to the states' eligibility criteria for Aid to Families with Dependent Children (AFDC), the major welfare program that was already in place. Specifically, to be eligible for Medicaid, one had to meet their state's AFDC income and asset requirements and also meet the categorical eligibility requirements for Medicaid, which essentially consisted of being a female head of household with children under age 21.

In addition to varying in their implementation timing and in their eligibility generosity as a function of other "welfare" programs such as AFDC, Aid to the Blind and Totally Disabled and to the poor elderly, states varied as to whether they covered the so-called "medically needy" under their Medicaid programs. These were individuals whose income was above the level which would make them eligible for AFDC, but whose medical expenses were so high that their income net of medical expenses was below some level. The allowed income and asset levels to qualify as medically needy were more generous than those required for AFDC eligibility in some states and less generous in others (Table 2.1). The medically needy were further classified as categorically-related or non-categorically related, where the former group met the categorical requirements for the relevant welfare program (i.e., family structure for AFDC) but their income was above the limit for eligibility while the latter group met neither the categorical or income requirements. Title XIX offered states federal matching funds for both medical and administrative costs for the categorically-related medically needy but only for administrative costs for the non-categorically related medically needy. While states were required to cover individuals who received or were eligible for welfare assistance, they were allowed to include both medically needy groups, neither group, or only the categorically-related medically needy under their Medicaid programs. Table 2.2 describes coverage of the medically needy by the states included in this analysis.

Table 2.1
AFDC Need Standards and Income Cutoffs
for Medically Needy Programs

	Annual Standard of Need for an AFDC Family of 4 in 1974	Annual Income that may be Retained for a Medically Needy Family of 4 in 1970	Difference between Medically Needy and AFDC Eligibility Income Levels
ALMS	$2,700/$3,324		
CA	$4,164	$3,600	-$564
CT	$3,868	$4,400	$532
DC	$4,188	$3,560	-$628
FL	$2,671		
IL	$3,216	$3,600	$384
IN	$3,158		
MIWI	$4,368/$5,472	$3,540/$3,100	-$828/-$2,372
NJ	$4,272		
NY	$3,096	$5,000	$1,904
OH	$4,668		
PA	$4,188	$4,000	-$188
TX	$2,244		
NEWENG			
ME	$4,188		
MA	$3,648	$4,176	$528
NH	$2,652	$4,056	$1,404
RI	$3,588	$4,300	$712
VT	$4,572	$3,420	-$1,152

Source: Need Standards from "Characteristics of State Plans for Aid to Families with Dependent Children Under the Social Security Act, Title IV-A: Need, Eligibility, Administration," U.S. Dept of Health, Education, and Welfare, Social and Rehabilitation Service, Assistance Payments Administration, 1974 Edition; Medically Needy Income Levels from "Characteristics of State Medical Assistance Programs Under Title XIX of the Social Security Act," U.S. Dept of Health, Education, and Welfare, Social and Rehabilitation Service, Assistance Payments Administration, 1970 Edition.

Table 2.2
Coverage of the Medically Needy as of 1970

	Federal Cost Sharing in Medical and Administrative Expenditures			Federal Cost Sharing in Administrative Expenditures Only	
	Categorically Related Needy	Categorically Related Medically Needy	Medically Needy Under 21 Years	General Assistance Recipients	Medically Needy Between Ages 21 and 64
ALMS	1				
CA	1	1			
CT	1	1	1		
DC	1	1	1		
FL	1				
IL	1	1			
IN	1				
MIWI	1	1	1		
NJ	1				
NY	1	1	1	1	1
OH	1				
PA	1	1	1	1	1
TX	1				
NEWENG					
ME					
MA	1	1	1		
NH	1	1			
RI	1	1			
VT	1	1	1		

Source: "Characteristics of State Medical Assistance Programs Under Title XIX of the Social Security Act, U.S. Dept of Health, Education, and Welfare, Social and Rehabilitation Service, Assistance Payments Administration, 1970 Edition.

Section 3: Theoretical Effects of Medicaid on Labor Supply

The AFDC and Medicaid programs provided a guaranteed level of income and health insurance for beneficiaries who had income and assets below a certain level. This eligibility cutoff and any earned income that would be disregarded was determined by, and varied across, states. Figure 2.1 shows the budget constraint ADE and the tax rate τ faced by a person not eligible for AFDC. Prior to the implementation of the Medicaid program, an individual who met the categorical and financial eligibility criteria would face the budget constraint AFDE, where AF is the size of the maximum AFDC cash benefit. As beneficiaries worked and earned income, their AFDC benefits were taxed away at a rate τ_{AFDC} until they reached a level of income, $Y_{breakeven}$, which reduced their benefits to zero.

Figure 2.1: Budget Constraints Under AFDC and Medicaid Programs Without Medically Needy Provisions

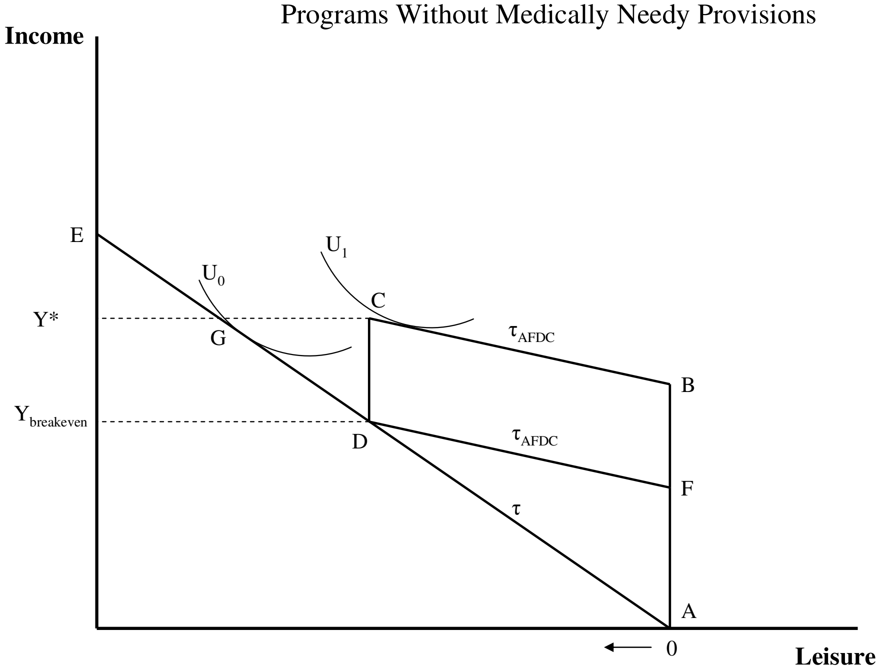
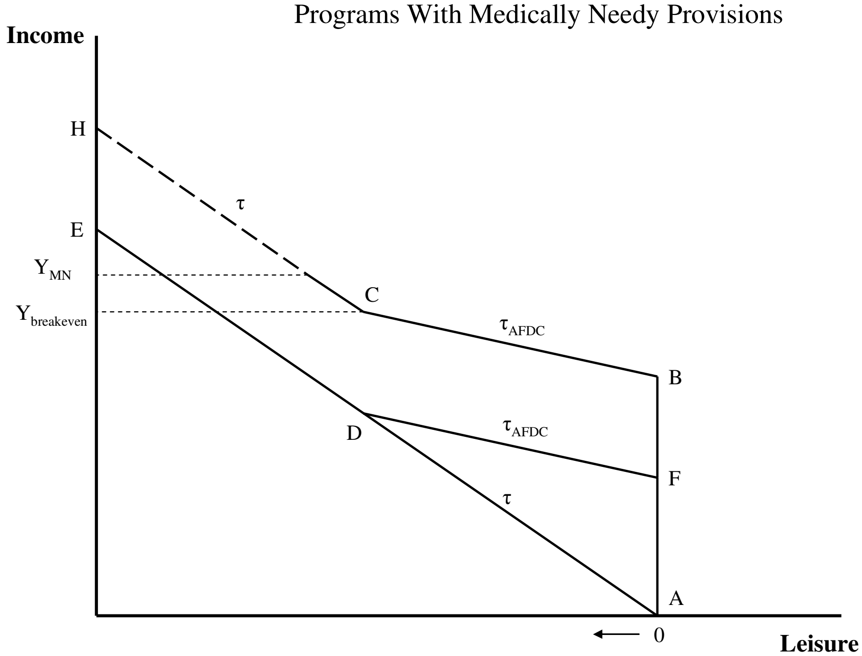


Figure 2.2: Budget Constraints Under AFDC and Medicaid Programs With Medically Needy Provisions



When Medicaid was implemented, AFDC beneficiaries in states that did not have a provision to cover the medically needy faced the budget constraint ABCDE, with segment FB representing the value of the Medicaid benefit. Beneficiaries maintained their eligibility for the entire package of Medicaid benefits and services and faced the τ_{AFDC} marginal tax rate as long as their earnings fell below $Y_{breakeven}$. Once beneficiaries earned above the breakeven level however, they became ineligible for AFDC and all their Medicaid benefits were lost. This discontinuous drop in benefits is known as the "Medicaid notch" and it creates a dominated part of the budget set represented by segment DG. An individual would be better off earning less money and maintaining eligibility for AFDC and Medicaid (remaining on segment BC at U_1) than earning between $Y_{breakeven}$ and Y* and locating on the DG segment of the budget set at U_0. As Yelowitz (1995) pointed out, in order to know what region of work hours or income is dominated, one would need to know what value the beneficiary places on the Medicaid benefits received.

In states that did cover the medically needy under their Medicaid programs and had medically needy cutoffs above their AFDC cutoffs, an individual would have faced the budget constraint ABCH, shown in Figure 2.2. Even if they had income and assets high enough to be ineligible for AFDC (greater than $Y_{breakeven}$), they could obtain Medicaid coverage by "spending down" to the medically-needy eligibility level, Y_{MN}. That is, if an individual's income and assets net of medical expenses met the eligibility level, they could receive Medicaid coverage for all additional medical expenses within a designated time period. Since residents in states with coverage for the medically needy did not face a fixed income level at which they lost all benefits associated with Medicaid coverage, they did not face the same non-convexity in their budget constraints nor any dominated portion of their budget sets.

These models provide some basic predictions with respect to the creation of the "Medicaid notch" and labor supply decisions. The new health insurance provided under Medicaid increased the well-being of those who were eligible and didn't work. As a result, the model predicts that those who were not participating in the labor force before Medicaid would remain out of the labor force. In addition, some who were working may have chosen to stop working to receive AFDC and Medicaid benefits – low wage workers without health insurance, for example. The theoretical effect of the budget constraint incentives associated with Medicaid on the participation decision is unambiguously negative. However, unlike a tax credit or other monetary transfer, the in-kind benefit of health insurance could have positive health impacts on mothers and/or children that would enable increased labor force participation.[4] As long as the earner stayed below the breakeven level, Medicaid benefits would be maintained. For example, better chronic disease management, for themselves or their children, may allow some single mothers to enter the labor force.

In states with Medically Needy programs, the theoretical predictions for labor supply are somewhat different. Though this program design does not eliminate the negative labor supply incentives entirely, the ability for individuals to spend-down to the eligibility level in the event of significant medical expenditures would smooth the "notch" and lessen the incentive to leave the labor force among those already working. The theoretical prediction of the effect of Medicaid on the labor force participation

[4] No data on health status are available in the 1962-75 CPS, but previous research using variation in Medicaid coverage due to the 1980's expansions has found positive effects of that coverage on health (Currie and Gruber, 1996a and 1996b).

decision is therefore negative in states without a medically needy provision and a smaller or no change in states with such a program. [5]

Section 4: Previous Research

A few studies have examined how the Medicaid program affects labor supply decisions. The well-designed studies indicate that changes in the Medicaid program do not have a significant effect on single women's labor supply, contrary to the theoretical predictions.

Blank (1989) and Winkler (1991) examined Medicaid's impact on the labor supply decision of female household heads in terms of participation, hours worked and AFDC participation, using variation in Medicaid's market value by state ("generosity") to identify the effect.[6] Blank reports statistically insignificant effects of average Medicaid value and the presence of Medically Needy programs on AFDC participation and labor supply. Winkler finds Medicaid has a small negative impact on an average female head's probability of being employed. Estimates based on these measures of Medicaid "generosity" are difficult to interpret, since averages based on state-level expenditures mask its true value to recipients which will vary according to their demographic characteristics, income level and health status.

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[5] Medicaid could also affect labor supply on the intensive margin of hours worked. This analysis is complicated by differences in the reference period for CPS survey questions on labor force participation and hours worked.

[6] Blank calculated a mean state-specific value of Medicaid for each household by summing insurance values reported in the state for an adult (controlling for disability status) and three children. Winkler calculated the market value using annual Medicaid expenditures for AFDC recipients divided by the unduplicated number of Medicaid recipients over the course of the year and annual Medicaid expenditures for AFDC recipients divided by the average monthly number of AFDC recipients.

Moffitt and Wolfe (1992) used expected medical expenditures and health conditions to introduce individual and family heterogeneity into the value of Medicaid. They also measured the availability of insurance should a female-headed family be off welfare. They estimate that increases in expected Medicaid benefits have a small negative effect on the likelihood of working, but it is only a minority of families (those with high expected medical expenditures) that alters their AFDC participation or employment decisions in response to Medicaid. They estimate that the response for the majority of female heads is essentially zero.[7] The fundamental problem here is that health status is likely to have an important direct effect on labor supply. It is unclear whether the labor supply response is due to poor health status or high valuation of Medicaid.

Decker (1993) estimated the impact of Medicaid implementation on AFDC participation and labor supply. She found a significant increase in AFDC caseloads but no change in labor force participation or real wages, concluding that most of the increase in AFDC participation was due to eligible households increasing participation, not decreasing work hours to become eligible.

Yelowitz (1995) assessed the impact of losing public health insurance on labor market decisions by examining Medicaid eligibility expansions in the 1980's that raised income limits and severed the tie between AFDC and Medicaid eligibility. His identification strategy is based on how much a state's new income limit increased over its previous AFDC level, noting the impact should be less in states that were already generous. Exploiting variation in the budget constraint for mothers with children of

[7] They also find that private health insurance is likely to have stronger effects on work incentives than Medicaid.

different ages within a state, across states, and over time, he found a small positive effect of the fully-phased in Medicaid expansion on labor force participation of single mothers. In examining the differential impact of the reform by marital status and education level for single women, Yelowitz found positive effects on labor force participation for ever married women and those with a high school diploma but little to no effect on never married women and those with more than a high school education. He suggests that the health insurance reform may be a way for short-term participants to exit welfare, but is unlikely to have an effect on long-term participants.

Several more recent papers have highlighted flaws in Yelowitz's analysis. Meyer and Rosenbaum (2000, 2001) and Ham and Shore-Sheppard (2005) point out that his findings are sensitive to the parameterization of the key variable, which measures how much the Medicaid eligibility limit increased and is restricted to be zero or positive. He also constrains the AFDC and Medicaid income limits to have effects of equal magnitudes with opposite signs. When corrections are made to the calculation of the AFDC eligibility limit and the effects of Medicaid expansions are estimated separately from the effects of changes in AFDC, both pairs of authors find no statistically significant effect of the Medicaid expansions on labor force or welfare participation. The evidence suggests that changes in AFDC and employer-provided health insurance were much more important to employment changes among single mothers in this period.

This research contributes to this literature in a number of ways. Using variation in the timing of Medicaid implementation across states provides more convincing identification than the early studies that used measures of the value of Medicaid benefits. It also avoids calculating eligibility limits, which can introduce error into the analysis.

Much of the more convincing work described above focuses on the effects of Medicaid expansions in the 1980's. By examining the effect of implementation in the 1960's, I use a different natural experiment which enhances the generalizability of the results of the literature as a whole. I also focus on a different population, since those originally eligible for Medicaid were much poorer than those to whom the 1980's expansions applied.

Section 5: Methods

5.1 Empirical Strategy

The empirical analysis aims to identify the effect of the Medicaid program on labor supply using variation in the timing of implementation across states and differences in eligibility across demographic groups. I focus on labor force participation, since the predicted effect of the budget constraint analysis was unambiguously negative. I use a "difference-in-differences" (DD) approach to estimate changes in labor supply in states that had implemented the program relative to states that had not, controlling for time-invariant state differences in labor supply and labor supply time trends common to all states. I also employ a "difference-in-differences-in-differences" (DDD) approach to estimate the relative changes in labor supply of eligible and non-eligible individuals in states that had implemented Medicaid, relative to states that had not. The first strategy relies on the timing of Medicaid implementation being exogenous to other factors affecting trends in labor supply. The second relies on there being no shock that affects the relative labor supply of the eligible and non-eligible groups in the same state-years as Medicaid implementation (Gruber 1994).

As discussed in Section 2, Medicaid eligibility was initially a function of meeting a state's income and asset requirements for AFDC and of categorical requirements (i.e., family structure). Since the former are endogenous to labor supply, I define treatment and control groups based on demographic characteristics that are consistent with the categorical eligibility criteria. Specifically, single women with children were by far the largest group initially eligible for Medicaid.[8]

In addition to family structure, I explore other demographic characteristics that predicted receipt of public assistance payments in 1966 and are exogenous to labor force participation.[9] Among single women ages 20-50, the strongest predictors of public assistance receipt were previous marital status, having at least one child under age 18, and age (based on univariate regressions). Interactions between age and living in an urban area and between age and previous marital status also contain some explanatory power. Table 2.3 shows that after controlling for age, single women with children were nearly 5 times as likely to receive public assistance as single women without children (column a). Never married women were nearly 13 times less likely to receive assistance as married women whose spouse is absent. Given the importance of these factors, I define the treatment and control groups as single women with and without children,

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[8] Among adults ages 18-64, 2.9 percent of women received public assistance while only 1.4 percent of men did. Since men's labor force participation patterns were very different from women's, and were likely affected by different factors, I do not consider them a viable control group. Goldin (1990) notes that married women's labor force participation was increasing rapidly over this period due to both labor demand/supply and cultural reasons. Since the underlying trends in labor force participation likely differed between single and married women during this period, and may have been responding to different factors, I focus on single women. 7.7 percent of single women ages 18-64 received public assistance payments in 1966 while 1 percent of married women did.

[9] The earliest available data in the CPS on receipt of public assistance payments is from 1966. The survey asks about public assistance payments under the AFDC, Aid to the Blind and Totally Disabled, and Old Age Assistance programs. There are no survey questions regarding public assistance eligibility. While a few Medicaid programs were already in place at this time (CA, HI, IL, MN, ND, OK, and PA), they would only have been in effect for a few months before the March survey, minimizing any effect on labor force participation and/or public assistance receipt by the Medicaid program.

respectively, and do several sub-analyses splitting the sample by previous marital status. Non-white single women in 1966 were about 5 times less likely to receive public assistance, controlling for other covariates, while urban residence and education level generally did not have significant effects. I control for these factors in the regression analysis.

Table 2.3
Predicting Welfare Receipt in 1966 among Single Women Ages 20-50

	a	b	c	d	e
kids	0.4369 ***	0.4786 ***	0.4596 ***	0.4758 ***	0.5212 ***
	0.0991	0.1014	0.1186	0.1130	0.1119
separated	-0.1469	-0.0094	-0.2119	-0.2276	-0.0591
	0.0869	0.0970	0.6952	0.6774	0.6980
widowed	-0.8096 ***	-0.8705 ***	5.0638 ***	5.3131 ***	5.3735 ***
	0.1048	0.1117	0.6667	0.8224	0.8836
divorced	0.2774 ***	0.2171 *	-0.2061	-0.1960	-0.3103
	0.0718	0.0938	0.7960	0.7970	0.7840
never married	-1.2217 ***	-1.3506 ***	-1.2230 ***	-1.2193 ***	-1.4134 ***
	0.1466	0.1653	0.2613	0.2606	0.2658
urban		-0.0607		0.0571	0.1193
		0.0563		0.1366	0.1340
minority		-0.5162 ***			-0.5234 ***
		0.1055			0.1005
_cons	-1.1076 ***	-1.6843 ***	-1.0768 *	-1.1099 *	-1.7406 ***
	0.2957	0.4389	0.4989	0.5225	0.5033
age dummies	y	y	y	y	y
education dummies		y			y
age*marital status			y	y	y
age*urban				y	y
pbar	0.0970	0.0970	0.1156	0.1156	0.1156
r2_p	0.2102	0.2395	0.2067	0.2210	0.2488
N	4439	4439	3719	3719	3719

* p<.05, **p<.01, ***p<.001
Omitted category is married with spouse absent.
Coefficients are marginal effects from probit models.
Standard errors are below the coefficients and are clustered at the state level.

I begin by estimating the change in labor force participation of single women, before and after Medicaid implementation, relative to single women in states that did not implement Medicaid in the same year. I control for changes in labor force participation common across all states with year fixed effects and for fixed differences in participation between states with state fixed effects. I estimate the following probit equation on the sample of single women:

$$\text{prob}(\text{LFP}_{ist}) = \Phi(\beta_0 + \beta_1\text{Year}_t + \beta_2\text{State}_s + \beta_3\text{Mcaid}_{st} + \beta_4\text{X}_i + \varepsilon_{ist}) \qquad (1)$$

where LFP_{ist} is equal to one if woman *i* in state *s* in year *t* reported working full- or part-time in the previous week, Year_t and State_s are vectors of fixed effects, and Mcaid_{st} is an indicator equal to one if the observation is from a state and year with a Medicaid program in effect. The vector X_i includes age, age squared, education, education squared, and indicator variables for marital status (married with spouse absent, separated, widowed, divorced, never married), being nonwhite, and living in a central city. The coefficient of interest β_3 estimates the impact of the Medicaid program on the labor force participation of single women relative to the change in their participation in the absence of Medicaid. I also estimate this model separately for single women with and without children, both to set up the DDD analysis and to test for a significant "effect" among women not eligible for the program.

The "difference-in-differences-in-differences" approach compares the change in labor force participation among eligible single women with children, before and after Medicaid implementation, relative to the change for single women without children, who would not have been eligible based on the categorical requirements. I estimate:

$$\text{prob}(\text{LFP}_{ist}) = \Phi(\alpha_0 + \alpha_1\text{Year}_t + \alpha_2\text{State}_s + \alpha_3\text{Mcaid}_{st} + \alpha_4\text{Kids}_i + \alpha_5\text{Mcaid}_{st}{*}\text{Kids}_i + \alpha_6 X_i + \varepsilon_{ist}) \qquad (2)$$

where Year_t, State_s, Mcaid_{st} and X_i are as defined above and Kids_i is an indicator variable for having one or more children under age 18. Interaction terms for $\text{State}_s{*}\text{Kids}_i$ and $\text{Year}_t{*}\text{Kids}_i$ are also included and, together with Mcaid_{st}, control for characteristics of the treatment group that are constant over time within states, changes over time for the treatment group across all states, and changes over time within states.

Similar analyses are conducted comparing women in states with medically needy provisions to those without in order to estimate the effect of generosity and program design on labor supply. Given the difference in the budget constraints facing Medicaid recipients in states that do and do not have provisions to cover the medically needy (figures 2.1 and 2.2), I predicted a smaller labor supply response to Medicaid in states with the provision. The empirical analysis is constrained by the fact that variation in generosity as measured by the medically needy program is nearly equivalent to the variation in implementation timing. All the states that implemented in 1966 and 1967, except Ohio, had medically needy programs while all the states that implemented in 1968-70 did not. As a result, I confine the analysis to comparing the effect of the Medicaid program in Ohio to New York and Michigan/Wisconsin.[10] I estimate:

$$\text{prob}(\text{LFP}_{ist}) = \Phi(\gamma_0 + \gamma_1\text{Year}_t + \gamma_2\text{State}_s + \gamma_3\text{MN}_{st} + \gamma_4 X_i + \varepsilon_{ist}) \qquad (3)$$

where MN is equal to one if the observation is from New York or Michigan/Wisconsin (states with Medically Needy programs) after Medicaid implementation (1967).

[10] All three states implemented Medicaid in 1966. I selected New York and Michigan/Wisconsin due to geographic proximity and demographic characteristics among the adult population (men and women ages 18-64) that are similar to Ohio.

I also use the DDD methodology with these three states to test whether single women with children have a different labor supply response than single women without children, conditional on being in a state with a Medically Needy program, after Medicaid implementation. I estimate:

$$\text{prob}(LFP_{ist}) = \Phi(\eta_0 + \eta_1 Year_t + \eta_2 State_s + \eta_3 MN_{st} + \eta_4 Kids_i \qquad (4)$$

$$+ \eta_5 MN_{st}{*}Kids_i + \eta_6 X_i + \varepsilon_{ist})$$

where the variables are defined as above and interaction terms for $State_s{*}Kids_i$ and $Year_t{*}Kids_i$ are also included. Survey data are weighted to be nationally representative and standard errors in all models are clustered at the state level.

5.2 Data

Repeated cross-sectional data on individual demographic characteristics and labor supply come from the Current Population Survey (CPS) March supplements from 1963-1975. These years cover the 1966-70 implementation period as well as a window of time on either side. Implementation dates for each state's Medicaid program come from the U.S. Department of Health and Human Services "Characteristics of State Medicaid Programs Under Title XIX" publications. Only 11 individual states are identified in these years of the CPS data[11], so I also use observations that are coded as either Michigan or Wisconsin, as Alabama or Mississippi and as Maine, Massachusetts, New Hampshire, Rhode Island and Vermont combined. The key condition to using these states in this "combined" form is that they started their Medicaid programs in the same year, which is the case for Michigan/Wisconsin and Alabama/Mississippi. The states that comprise

[11] California, Connecticut, District of Columbia, Florida, Illinois, Indiana, New Jersey, New York, Ohio, Pennsylvania, and Texas.

New England implemented their Medicaid programs within a 12-month period (July 1966-July 1967). Individuals in these states are likely to be similar demographically and in their labor force participation patterns due to geographic proximity. Table 2.4 presents the Medicaid implementation date in each state and the first year it was considered in effect for the purposes of this analysis.

Table 2.4
Implementation Dates for Medicaid

	Implementation Date	First Year Effective
California	3/1/1966	1966
Connecticut	7/1/1966	1967
District of Columbia	7/1/1968	1969
Florida	1/1/1970	1970
Illinois	1/1/1966	1966
Indiana	1/1/1970	1970
New Jersey	1/1/1970	1970
New York	5/1/1966	1967
Ohio	7/1/1966	1967
Pennsylvania	1/1/1966	1966
Texas	9/1/1967	1968
Alabama	1/1/1970	1970
Mississippi	1/1/1970	1970
Michigan	10/1/1966	1967
Wisconsin	7/1/1966	1967
Maine	7/1/1966	1967
Massachusetts	9/1/1966	1967
New Hampshire	7/1/1967	1967
Rhode Island	7/1/1966	1967
Vermont	7/1/1966	1967

Note: Since the CPS data is from March, a program is considered in effect if it was implemented before March 1st of that year.
Source: Characteristics of State Medicaid Programs under Title XIX, US Department of Health and Human Services.

Table 2.5
Summary Statistics, Single Women Ages 20-50

	Total	Pre Medicaid	Post Medicaid	p value
Age	31.7	32.7	31.2	<.001
	0.144	0.166	0.164	
Mean Years Education	11.8	11.4	12.0	<.001
	0.099	0.122	0.093	
High School	70.7%	64.7%	73.1%	<.001
	0.014	0.016	0.013	
Minority	21.2%	21.4%	21.2%	0.822
	0.020	0.024	0.019	
Marital Status				
Married, Spouse Absent	6.0%	8.2%	5.0%	<.001
	0.004	0.007	0.003	
Separated	14.4%	14.4%	14.5%	
	0.009	0.009	0.010	
Widowed	9.2%	10.7%	8.6%	
	0.004	0.005	0.003	
Divorced	18.4%	17.0%	19.0%	
	0.023	0.024	0.023	
Never Married	51.9%	49.6%	53.0%	
	0.021	0.027	0.019	
Children Age 18 and Under	45.9%	42.9%	47.3%	<.001
	0.012	0.016	0.011	
Number of Children	1.0	1.0	1.0	0.547
	0.037	0.057	0.032	
Urban	47.3%	46.6%	47.6%	0.630
	0.043	0.050	0.041	
Labor Force Participation	65.1%	66.4%	64.4%	0.131
	0.005	0.011	0.006	
N	54,782	13,276	41,506	
N for education	52,556	11,050	41,506	

Education is missing in 1963.
Standard errors are below the coefficients and are clustered at the state level.

Table 2.5, continued
Summary Statistics, Single Women Ages 20-50

	With Children				Without Children			
	Total	Pre Medicaid	Post Medicaid	p value	Total	Pre Medicaid	Post Medicaid	p value
Age	31.1	31.8	30.8	<.001	32.2	33.3	31.6	<.001
	0.223	0.290	0.208		0.196	0.180	0.235	
Mean Years Education	11.3	10.9	11.4	<.001	12.3	11.8	12.5	<.001
	0.098	0.138	0.093		0.109	0.121	0.108	
High School	63.4%	56.9%	65.8%	<.001	77.0%	70.5%	79.7%	<.001
	0.017	0.023	0.016		0.013	0.013	0.013	
Minority	27.2%	27.1%	27.3%	0.917	16.1%	17.1%	15.7%	0.101
	0.025	0.031	0.023		0.017	0.019	0.016	
Marital Status								
Married, Spouse Absent	9.3%	14.1%	7.3%	<.001	3.2%	3.7%	2.9%	<.001
	0.006	0.011	0.005		0.002	0.004	0.001	
Separated	22.3%	21.8%	22.6%		7.7%	8.8%	7.2%	
	0.015	0.014	0.017		0.006	0.008	0.006	
Widowed	11.5%	13.4%	10.8%		7.3%	8.8%	6.6%	
	0.006	0.010	0.005		0.003	0.004	0.003	
Divorced	24.1%	21.3%	25.2%		13.6%	13.9%	13.5%	
	0.029	0.029	0.029		0.017	0.019	0.016	
Never Married	32.8%	29.5%	34.2%		68.2%	64.8%	69.9%	
	0.019	0.021	0.019		0.018	0.024	0.016	
Children Age 18 and Under	100.0%	100.0%	100.0%		0.0%	0.0%	0.0%	
	0	0	0		0	0	0	
Number of Children	2.3	2.4	2.2	0.006	0	0	0	
	0.035	0.061	0.030		0	0	0	
Urban	45.5%	44.0%	46.2%	0.332	48.7%	48.6%	48.8%	0.907
	0.041	0.045	0.040		0.046	0.055	0.042	
Labor Force Participation	54.4%	55.0%	54.2%	0.604	74.1%	75.0%	73.7%	0.196
	0.011	0.013	0.013		0.006	0.011	0.006	
N	25,458	5,805	19,653		29,324	7,471	21,853	
N for education	24,486	4,833	19,653		28,070	6,217	21,853	

Education is missing in 1963.
Standard errors are below the coefficients and are clustered at the state level.

The sample consists of 54,782 single women between the ages of 20 and 50, with and without children under age 18. Table 2.5 presents summary statistics for several subgroups. Though some differences are statistically significant between single women in state-years before and after Medicaid implementation, none save for high school graduation rates are particularly large in magnitude. Among single women with children, the difference in high school graduation rates before and after Medicaid persists, and we also see important differences in (previous) marital status. Among single women without children, we again see large and significant differences in high school graduation rates but not for other covariates.

Section 6: Results

6.1 First Differences

Before estimating the effect of Medicaid on labor supply decisions, I examine the time trends in labor force participation for single women with and without children. Figure 2.3 shows labor force participation rates among single women ages 20-50 from 1963-75. In the 20 states included in this analysis, about 55% of single women with children reported working in the previous two weeks while over 75% of single women without children did. The trends are fairly flat over this time period for both groups.

Figure 2.4 shows the trends for the two groups by time from Medicaid implementation. We see a slight increase in labor force participation rates among single women with children around the time of Medicaid implementation (t=0) and some divergence of the two trendlines 5 years after implementation. Among single women with children who were previously married (Figure 2.5), labor force participation was trending down before Medicaid implementation and rose in the few years after. We see an upward trend in the pre-period and a downward trend in the post-period for never married single women with children. This may be consistent with Yelowitz's (1995) assertion that having children and never having been married is a proxy for long-term welfare dependency. In the case where health insurance is tied to welfare receipt, they may be more sensitive to the labor supply incentives associated with Medicaid.

Figure 2.3

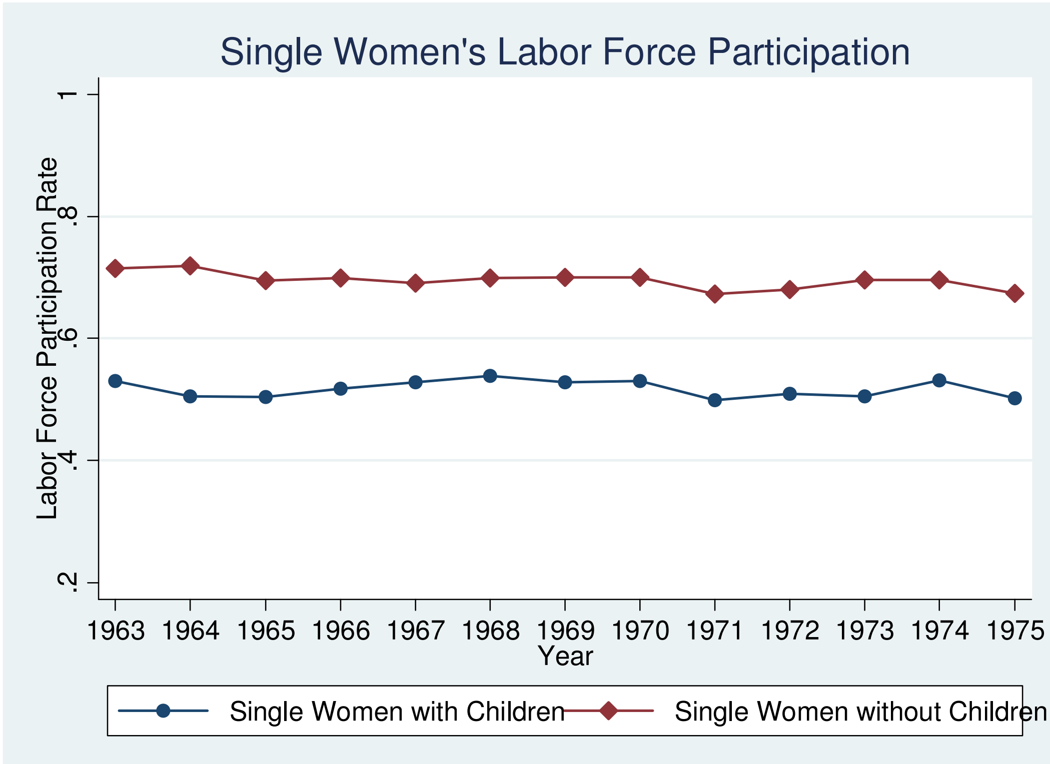
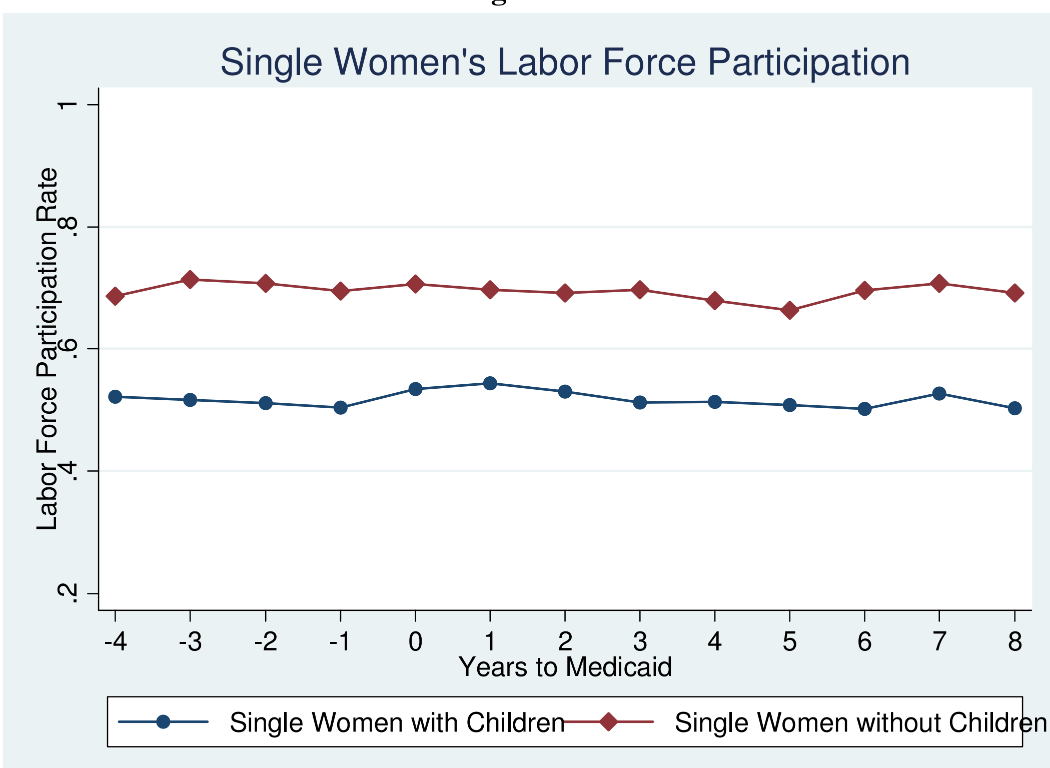


Figure 2.4



6.2 Estimated Effects of Medicaid

Table 2.6 presents the results from estimating equation 1. For all single women, the coefficient on Medicaid is positive but not statistically significant, indicating that the additional health insurance benefit did not have a significant effect on their labor force participation relative to their participation in states that had not yet implemented the program. Among single women with children (those eligible for Medicaid), we see larger but still statistically insignificant positive impacts of Medicaid on labor force participation (3 percentage points). For single women without children who were not eligible, we see no effect. Non-white single women are less likely to work, while the coefficients on age, education, urban residence are all significant and positive. Women who are separated, widowed, divorced, or never married are also more likely to work than married women with an absent spouse. The demographics have similar effects for single women with and without children, with the exception of urban residence which has a negative impact on labor force participation among women with children and a positive impact among women without children.

Table 2.6
Estimated Effect of Medicaid on Single Women's Labor Supply
Difference-in-Differences Estimate

	Single Women		**Single Women with Kids**		**Single Women without Kids**	
medicaid	0.0091	0.0110	0.0250	0.0297	0.0021	-0.0022
	0.0114	0.0108	0.0190	0.0206	0.0091	0.0086
minority		-0.1113 ***		-0.1105 ***		-0.0844 ***
		0.0161		0.0201		0.0138
age		0.0331 ***		0.0186 **		0.0472 ***
		0.0026		0.0065		0.0017
age sq		-0.0004 ***		-0.0002		-0.0006 ***
		0.0000		0.0001		0.0000
education		0.0578 ***		0.0618 ***		0.0557 ***
		0.0079		0.0086		0.0082
education sq		-0.0013 ***		-0.0013 ***		-0.0016 ***
		0.0003		0.0004		0.0003
urban		0.0210 **		-0.0196 *		0.0400 ***
		0.0074		0.0096		0.0108
separated		0.0705 ***		0.0975 ***		0.0525 **
		0.0127		0.0180		0.0184
widowed		0.0389 **		0.0327		0.0408 *
		0.0138		0.0202		0.0161
divorced		0.1888 ***		0.2510 ***		0.1013 ***
		0.0124		0.0178		0.0208
never married		0.2491 ***		0.2268 ***		0.1196 ***
		0.0107		0.0163		0.0197
_cons	0.5198 ***	-3.0594 ***	0.1353 *	-2.6158 ***	0.8127 ***	-3.3409 ***
	0.0487	0.1267	0.0676	0.1970	0.0530	0.1782
pbar	0.6506	0.6495	0.5440	0.5433	0.7411	0.7401
r2_p	0.0032	0.0725	0.0056	0.0702	0.0049	0.0511
N	54,779	52,553	25,457	24,485	29,322	28,068

* p<.05, **p<.01, ***p<.001
All models include state and year fixed effects.
Omitted category is married with spouse absent.
Coefficients are marginal effects from probit models, except for the constant term.
Standard errors are below the coefficients and are clustered at the state level.

I expanded equation 1 to include indicator variables for the years prior to and after Medicaid implementation. Figure 2.6 shows the pattern of the coefficients on these variables, relative to the omitted category of one year prior to implementation. The positive but declining coefficients prior to Medicaid implementation indicate that labor force participation among single women was increasing more slowly relative to states that were not approaching Medicaid implementation. This is the case for both single women with and without children. While the relative trend continues to decline for single women without children, for single women with children it reverses in the post-period and climbs upward. This suggests that the labor force participation rate of single women with children was growing at the same rate or faster in states that had implemented Medicaid relative to states that had not. Both Table 2.6 and Figure 2.6 suggest a positive, though not statistically significant, effect of Medicaid, which is not in the direction we would expect based on the budget constraint analysis.

Figure 2.7 graphs the same coefficients for single women with children, separately by previous marital status. For those who were previously married, the relative trend continues to go against the prediction: it is downward-sloping in the pre-period and then increases post-Medicaid. For never married women with children, the relative trend increases in the pre-period and then decrease post-Medicaid, though none of the coefficients are statistically significant. This is the labor supply response that the budget constraint analysis predicts and, if these women are indeed more dependent on welfare, they may react more strongly to the negative labor supply incentives created by the implementation of Medicaid.

Figure 2.6: Coefficients on Lag Variables, with Demographic Controls

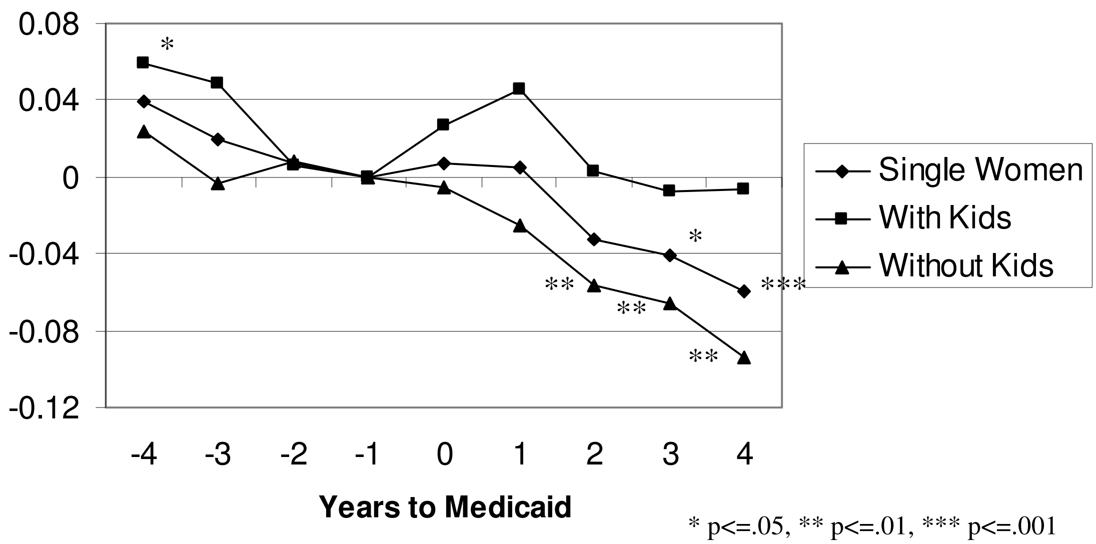
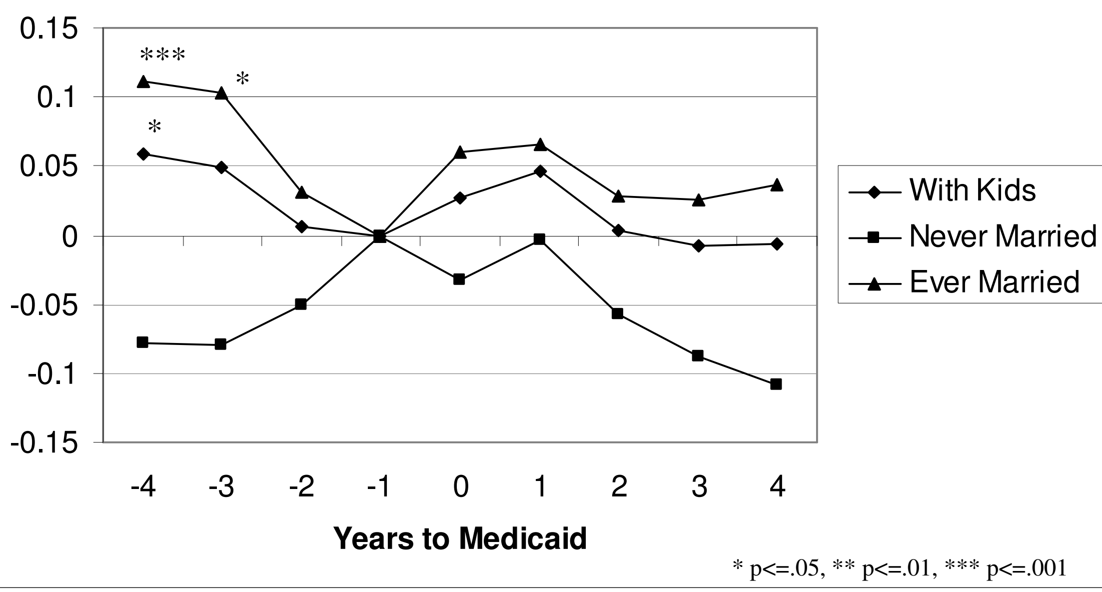


Figure 2.7: Coefficients on Lag Variables, with Demographic Controls, Single Women with Children



The results of the triple-difference analysis are reported in Table 2.7. The coefficient on the DDD estimator "kids*Medicaid" is positive and not statistically significant, indicating that, conditional on being in a state with its Medicaid program in effect, the change in labor force participation of single women with children was not statistically different from that of single women without children. Again, these results challenge our initial predictions since single women with children were eligible for Medicaid, and we therefore expected that they would have a negative labor supply response to the associated incentives.

Table 2.7
Estimated Effect of Medicaid on Single Women's Labor Supply
Triple Differences Estimate

	a	b
kids	-0.2478 ***	-0.1984 ***
	0.0253	0.0212
medicaid	0.0024	0.0008
	0.0104	0.0101
kids*medicaid	0.0207	0.0241
	0.0184	0.0189
minority		-0.1016 ***
		0.0161
age		0.0336 ***
		0.0030
age sq		-0.0004 ***
		0.0000
education		0.0618 ***
		0.0081
education sq		-0.0016 ***
		0.0003
urban		0.0137
		0.0080
separated		0.0741 ***
		0.0134
widowed		0.0446 ***
		0.0122
divorced		0.1860 ***
		0.0126
never married		0.1904 ***
		0.0108
_cons	0.8127 ***	-2.5969 ***
	0.0530	0.1180
pbar	0.6506	0.6495
r2_p	0.0380	0.0875
N	54,779	52,553

* p<.05, **p<.01, ***p<.001
All models include state, year, kids*state, and kids*year fixed effects.
Omitted category is married with spouse absent.
Coefficients are marginal effects from probit models, except for the constant term.
Standard errors are below the coefficients and are clustered at the state level.

6.3 Differences by Previous Marital Status

Splitting the sample of single women with children into those who were never married and ever married, Table 2.8 shows positive but statistically insignificant effects of Medicaid for both groups, though the estimate is much larger for the ever married group. The triple difference estimate is similarly not significant, but again larger and positive for the ever married group. These results are consistent with Yelowitz's hypothesis that never married women with children are more persistently tied to welfare than their ever married counterparts. Though the results suggest that both groups are increasing their labor supply after Medicaid implementation, the never married group increases less, hewing more closely to the negative incentives associated with the program.

Table 2.8
Estimated Effect of Medicaid on Single Women's Labor Supply
By Previous Marital Status

	Single Women with Kids		**Single Women**	
	Never Married	**Ever Married**	**Never Married**	**Ever Married**
	a	**b**	**c**	**d**
medicaid	0.0091	0.0453	0.0093	-0.0213
	0.0193	0.0285	0.0254	0.0252
kids			-0.0761 *	-0.3096 ***
			0.0358	0.0361
kids*medicaid			-0.0069	0.0643
			0.0295	0.0374
minority	-0.2575 ***	-0.0505 *	-0.1745 ***	-0.0463 *
	0.0339	0.0241	0.0175	0.0208
age	0.0530 ***	0.0172 *	0.0621 ***	0.0177 ***
	0.0127	0.0074	0.0027	0.0044
age sq	-0.0008 ***	-0.0002	-0.0009 ***	-0.0002 **
	0.0002	0.0001	0.0000	0.0001
education	0.1107 ***	0.0320 ***	0.0830 ***	0.0267 ***
	0.0131	0.0046	0.0101	0.0042
education sq	-0.0047 ***	0.0006 *	-0.0028 ***	0.0005 *
	0.0006	0.0003	0.0004	0.0002
urban	0.0290 *	-0.0462 ***	0.0534 ***	-0.0364 ***
	0.0136	0.0124	0.0111	0.0083
separated		0.1028 ***		0.0885 ***
		0.0184		0.0144
widowed		0.0330		0.0475 **
		0.0212		0.0144
divorced		0.2632 ***		0.2206 ***
		0.0201		0.0171
_cons	-2.9657 ***	-2.6312 ***	-3.8946 ***	-1.6133 ***
	0.4635	0.2462	0.2761	0.1967
pbar	0.5822	0.5242	0.7021	0.5928
r2_p	0.0531	0.0975	0.0814	0.0966
N	8,002	16,483	27,137	25,416

* p<.05, **p<.01, ***p<.001
All models include state and year fixed effects. Columns c and d add kids*state and kids*year fixed effects.
Omitted category is married with spouse absent.
Coefficients are marginal effects from probit models, except for the constant term.
Standard errors are below the coefficients and are clustered at the state level.

6.4 Estimated Effect of the Medically Needy Program

Results from estimating equation 3 are presented in Table 2.9. The estimated coefficient γ_3 on the Medically Needy program indicator shows that single women in states with a Medically Needy program in effect are 9 percentage points (14 percent) less likely to work than single women in states without the program (in this case, Ohio). We see a much larger estimated effect among single women with children (17 percentage points) and a smaller (and less statistically significant) effect among single women without children. These results are surprising, given that the budget constraint analysis suggested that women in states with Medically Needy programs would not face the same incentives to reduce their labor supply that women in states without such programs would.

Table 2.9
Estimated Effect of the Medically Needy Program on Single Women's Labor Supply
Difference-in-Differences Estimate

	Single Women		**Single Women with Kids**		**Single Women without Kids**	
NY	0.0425 ***	0.0640 ***	0.0517 ***	0.1191 ***	0.0200 *	0.0287 *
	0.0098	0.0091	0.0084	0.0042	0.0099	0.0136
MIWI	0.0438 ***	0.0656 ***	0.1070 ***	0.1413 ***	0.0054	0.0245
	0.0093	0.0099	0.0077	0.0028	0.0100	0.0134
MN program	-0.0663 ***	-0.0929 ***	-0.1358 ***	-0.1724 ***	-0.0198	-0.0423 *
	0.0142	0.0119	0.0116	0.0039	0.0148	0.0173
minority		-0.1230 **		-0.1355 **		-0.0850 **
		0.0396		0.0426		0.0313
age		0.0284 ***		0.0127		0.0450 ***
		0.0035		0.0076		0.0041
age sq		-0.0003 ***		-0.0001		-0.0006 ***
		0.0000		0.0001		0.0001
education		0.0548 ***		0.0711 ***		0.0470 ***
		0.0075		0.0067		0.0092
education sq		-0.0011 ***		-0.0014 ***		-0.0013 ***
		0.0003		0.0002		0.0003
urban		0.0221 **		-0.0029		0.0319 ***
		0.0068		0.0134		0.0063
separated		0.0555 ***		0.0890 **		0.0314
		0.0056		0.0318		0.0318
widowed		0.0297		0.0313		0.0156
		0.0261		0.0427		0.0151
divorced		0.1673 ***		0.2322 ***		0.0728 **
		0.0132		0.0320		0.0237
never married		0.2645 ***		0.2531 ***		0.0949 **
		0.0041		0.0064		0.0367
_cons	0.4434 ***	-2.6551 ***	-0.0383	-2.6090 ***	0.8068 ***	-3.0840 ***
	0.1249	0.2866	0.1965	0.4405	0.1038	0.2988
pbar	0.6441	0.6432	0.5121	0.5130	0.7481	0.7465
r2_p	0.0028	0.0790	0.0067	0.0809	0.0019	0.0425
N	17,225	16,516	7,702	7,408	9,523	9,108

* p<.05, **p<.01, ***p<.001
All models include year fixed effects.
Omitted categories are Ohio and married with spouse absent.
Coefficients are marginal effects from probit models, except for the constant term.
Standard errors are below the coefficients and are clustered at the state level.

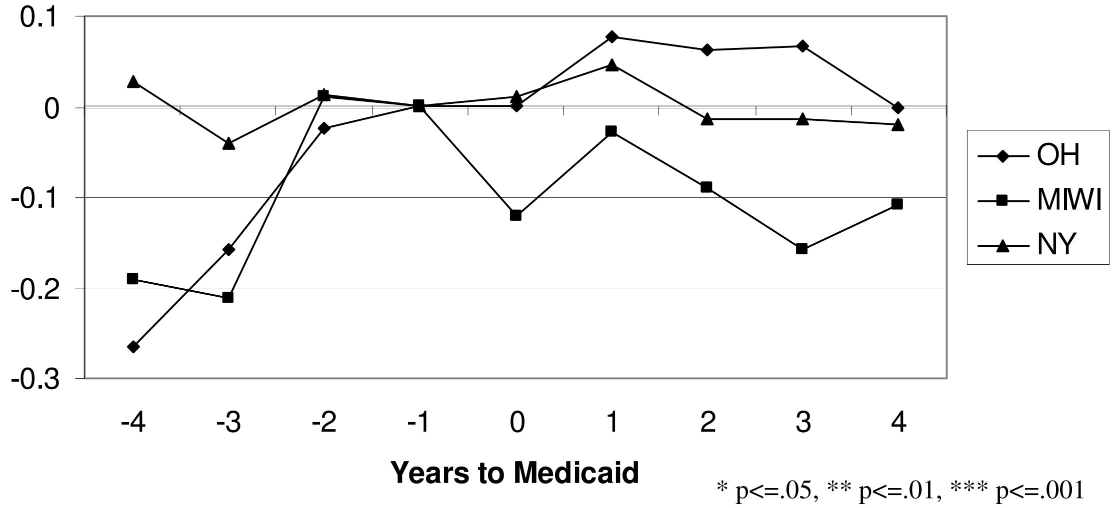
Table 2.10 presents results from the triple-differences analysis. There was an 11 percentage point fall in the relative labor force participation rate of single women with children in states with a Medically Needy program relative to the change in relative participation rates in Ohio. Again, this is not the predicted direction of the effect. Turning to Figure 2.8, we can see that the trend in labor force participation among single women with kids was flat or increasing in the pre-period in all three states relative to single women without kids. In the post-period, the relative trend continues to increase in Ohio, remains relatively flat in New York, and falls in Michigan/Wisconsin. Of all the 14 "states" included in the main analysis, Ohio is the only one with a marked, sustained increase in labor force participation rates among single women with children after Medicaid implementation. To the extent that there was something unusual about Ohio or there were other factors affecting labor supply decisions there during this period, the lack of a good control group complicates the analysis of the medically needy provisions.

Table 2.10
Estimated Effect of the Medically Needy Program on
Single Women's Labor Supply
Triple Differences Estimate

	a	b
kids	-0.3338 ***	-0.2916 ***
	0.0576	0.0232
MN program	-0.0230	-0.0449 *
	0.0172	0.0215
kids*MN program	-0.1060 ***	-0.1100 ***
	0.0070	0.0234
NY	0.0232 *	0.0385 *
	0.0115	0.0166
MIWI	0.0063	0.0219
	0.0116	0.0173
kids*NY	0.0246 ***	0.0530 **
	0.0039	0.0168
kids*MIWI	0.0896 ***	0.1005 ***
	0.0041	0.0156
minority		-0.1129 **
		0.0386
age		0.0285 ***
		0.0040
age sq		-0.0003 ***
		0.0001
education		0.0598 ***
		0.0072
education sq		-0.0014 ***
		0.0002
urban		0.0135
		0.0108
separated		0.0612 ***
		0.0080
widowed		0.0389
		0.0269
divorced		0.1654 ***
		0.0106
never married		0.1967 ***
		0.0074
_cons	0.8068 ***	-2.3709 ***
	0.1038	0.2772
pbar	0.6441	0.6432
r2_p	0.0502	0.0976
N	17,225	16,516

* p<.05, **p<.01, ***p<.001
All models include year and year*kids fixed effects.
Omitted categories are Ohio, kids*Ohio, and married with spouse absent.
Coefficients are marginal effects from probit models, except for the constant term.
Standard errors are below the coefficients and are clustered at the state level.

Figure 2.8: Coefficient on Single Women with Kids*Lag, with Demographic Controls



6.5 Robustness

I ran several robustness checks to test the sensitivity of the results. One at a time, I expanded the age range to 18-55, dropped Ohio and New England from the sample, defined single as separated, divorced, widowed or never married (dropping married with spouse absent), set labor force participation equal to one if the respondent was working, had a job but was not at work, or was unemployed in the previous week. The statistical significance and the magnitudes of the estimates do not change.

Section 7: Conclusions

This paper examines the effect of the implementation of the Medicaid program on single women's labor supply. Taking advantage of variation in the timing of implementation across states and selecting treatment and control groups based on demographic characteristics that predict welfare receipt (i.e., also being eligible for Medicaid), I estimate that Medicaid did not have a statistically significant impact on labor force participation among single women. The theoretical model predicted a negative labor supply response given the close tie between AFDC receipt and Medicaid eligibility. Based on the 95 percent confidence intervals of the estimated effects of Medicaid on single women's labor supply, I am able to rule out negative effects greater than the following magnitudes: 2 percent for the DD estimate among single women with children (1 percentage point); 2 percent for the DDD estimate (1 percentage point); 9 percent for the DDD estimate among never married women (6 percentage points); and 2 percent for the DDD estimate among ever married women (1 percentage point). As discussed above, never married women may be more sensitive to the labor supply incentives when Medicaid is tied to AFDC receipt.

While the results should be interpreted with caution given their lack of statistical significance, the divergence from the theoretical predictions raises some important, policy-relevant questions. If the positive health effects of having health insurance for a single mother and her children affect the extensive labor supply margin, they have the potential to partially offset the negative labor supply incentives. In her analysis of the health impacts of Medicaid's implementation, Decker (1993) found large and statistically significant declines in the probability of low-birthweight for the children of low-income, non-white and single women. Later analyses by Currie and Gruber (1996a and 1996b) found negative effects of Medicaid coverage on infant and child mortality and positive effects on health care utilization among children. These improvements in health status for children may have operated to increase labor supply, in conjunction with factors like the negative stigma attached to welfare receipt and changing norms regarding women's labor force participation. Future work that examines the health impacts on adults would help illuminate how important these effects are in explaining the labor force participation findings.

Alternatively, low-income single women with children may not be very responsive to labor supply incentives associated with public health insurance programs. AFDC earning limits were very low in the early 1970's (equivalent to $10,000 - $20,000 for a family of 4 in 2007 dollars)[12] and additional income may have been much more valuable than health insurance. Simply put, it may not have been worth it to be that poor

[12] The AFDC standards of need for Texas and Ohio in Table 1, adjusted using the inflation calculator at http://www.bls.gov/cpi/.

just to get Medicaid coverage. Health insurance was not as valuable as it is today, given the relative lack of sophistication and low cost of health care itself. Furthermore, low-income families likely had ways to get low-cost or free health care when they needed it, making the marginal value of Medicaid in terms of access to care relatively low.[13]

It is, of course, the value that beneficiaries place on Medicaid coverage that in large part determines their sensitivity to the associated incentives. As Gruber and Madrian's (2002) review of the economic literature on health insurance and labor supply highlights, some groups are quite responsive while others are not (early retirees vs. low-income single mothers). Based on this analysis and the previous work described in Section 4, there is an emerging consensus in the economics literature that low-income single mothers are not responsive to these incentives. Given the substantial health benefits, low direct costs, lack of significant labor supply distortions, and low rates of employer offer of health insurance in low-wage jobs, public provision of health insurance for low-income families appears to be a reasonable and efficient way to cover this population.

Further research is needed to better understand how low-income families think about, and value, health insurance coverage. Not only will this knowledge better inform economic research, it will also strengthen the design of social insurance programs and policies to reduce the number of uninsured in the U.S.

[13] Additional analysis using the CPS data suggests that the labor force participation rate of black single women with children in the South declined by 10 percentage points relative to other single women with children ($p=0.003$). Medicaid would have been more valuable to these women since it's likely they had less access to other sources of charity care. This higher valuation could help explain their differential labor supply response.

Racial/Ethnic Disparities in Outpatient Primary Care:
The Role of Physician-Patient Concordance[*]

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Section 1: Background

Racial/ethnic disparities in health and health care are well documented in both the academic literature and the popular press. Minority patients tend to receive lower quality and intensity of care than whites, both for diagnostic and therapeutic services. Perhaps most importantly, disparities in care have been linked to worse clinical outcomes and higher mortality rates among minorities. Despite all that we do know about the presence and pervasiveness of disparities, the mechanisms underlying these differences are not well understood (Smedley, et al. 2002).

Though disparities may arise in part because white and minority patients see different physicians, even when they see the same providers, they may be treated differently.[1] In this case, discrimination – operating through the mechanisms of prejudice, stereotypes, and uncertainty – may drive disparities (Balsa and McGuire 2003). Racial/ethnic concordance between physicians and patients may affect health care disparities by reducing discrimination. In concordant pairs, favorable prejudice (in-group favoritism), modification of negative stereotypes, and increased clinical certainty on the part of the physician are thought to be more likely. For patients, concordance may promote trust and increase compliance. Due in part to these beliefs regarding the impact of concordance, policies such as increasing the number of minority physicians have been advanced to reduce disparities in care and outcomes.

Empirical research has explored the extent of racial concordance in the United States, the reasons for it, and its impact on selected outcomes. While the analyses that

[1] If minority patients generally see different providers than white patients do (across-provider variation), then characteristics of these providers that are associated with worse outcomes may be an important component of observed differences. For example, if these providers are of lower quality or have fewer resources available for their patients, minority patients are likely to receive inferior care (Chandra and Skinner 2003; Bach, et al. 2004).

have been done to date have explored a number of important issues, gaps still exist and we remain far from a consensus on the importance and benefit of racial matching between physicians and patients. This paper contributes to this literature by addressing methodological weaknesses of most existing studies, expanding outcomes to include a set of quality of care measures that have not yet been studied, and examining racial/ethnic concordance among physicians who see both white and minority patients.

An understanding of the mechanisms underlying racial/ethnic disparities in health care is necessary to design effective policy interventions. For example, if observed disparities arise primarily because minority patients are geographically concentrated in areas with relatively few health care resources, policies that increase resources for all patients in those areas may be most effective. Alternatively, if poor communication, lack of understanding between patients and physicians, or discrimination is an important driver of disparate outcomes, taking steps to improve communication and understanding between people of different backgrounds or race/ethnicity may prove more effective than redistributing resources. Responses to differences in treatment that arise even when patients see the same providers include increasing information, awareness and cultural competence, creating and enforcing guidelines or rules for care delivery, and increasing the number of minority physicians.

Section 2: Related Literature

This section discusses the existing literature related to three main topics affecting the concordance debate: rates of concordance, the distribution of patients across physicians and reasons for concordance, and the role of concordance in quality of care outcomes. I also highlight the contributions of this analysis to each of these areas.

If patients were randomly distributed across physicians, we would expect the following concordance rates: 75 percent among white patients, 4 percent for blacks, 5 percent for Hispanics and 13 percent for Asians.[2] In a 2001 nationally-representative population survey, 85 percent of white respondents reported being concordant with their regular physician, compared to 25 percent of blacks, 28 percent of Hispanics and 45 percent of Asians (Saha, et al. 2003). A 1997 survey of managed care enrollees in the Washington, DC metropolitan area yielded a 67 percent concordance rate among whites and 43 percent among blacks (Cooper-Patrick, et al.1999). Clearly, there are mechanisms at work that increase concordance rates above what we would expect from random matching, particularly for minority groups. Due to the high percentage of white physicians, we can expect that concordance rates among white patients will be highest in almost all contexts, but important variation may exist by geographic location and characteristics such as insurance type.

While the concordance rates found in previous studies are based on national population percentages, understanding the role of concordance is especially important for groups that use health care services at higher rates. This analysis provides an important and unknown statistic: concordance rates in a sample that is representative of outpatient visits rather than representative of the population. Young children, the elderly, and those with chronic conditions will be over-represented in a sample of visits, relative to a population sample. The roles of discrimination and concordance in contributing to

[2] Race/ethnicity is known for 64% of physicians. 75% are white, 4% are black, 5% are Hispanic, 13% are Asian and 4% are in another category (Smart 2006).

disparate outcomes are of particular interest for subpopulations who have access to the health care system and interact with it most.

The evidence on the reasons behind racial concordance points to two factors that are likely to be important: geography and patient and physician preference. Stinson and Thurston (2002) find that after controlling for physician specialty, practice setting and location, large differences by physician race in the racial makeup of their patient panel are decreased. They argue that location plays perhaps the most important role in determining how many minority patients a doctor has. Clearly, physician preferences are involved here as well, since they will help determine which specialty a provider chooses and where he or she locates. Komaromy, et al. (1996) found that even after controlling for the racial makeup of the community, black physicians see more black patients and Hispanics see more Hispanics than other physicians, indicating that location is not the only important factor. Survey results have also highlighted the role of patient preference. Saha, et al. (2000) found that blacks and Hispanics seek care from physicians of their own race because of personal preference and language, not just geographic accessibility.

While direct measures of patient and physician preferences are not available in these data, I investigate how patients of different racial/ethnic groups are distributed across primary care physicians. I also analyze differences in physician, practice and geographic characteristics by patient race and the minority patient composition of physician's practices. These results contribute additional evidence to this strand of the literature examining how and why concordance occurs.

With respect to the contribution of concordance to quality of care outcomes, published work has focused on the effect of racial concordance between patients and their usual provider on outcomes such as patient satisfaction with care, patients' ratings of their physician, degree of participatory decision-making, and receipt of needed health services. While some studies have found a positive relationship[3], others have found no effect of racial/ethnic concordance on these outcomes.[4] For example, Cooper-Patrick, et al. (1999) found that patients in race-concordant relationships rated their visits as significantly more participatory than patients in race-discordant relationships and Saha, et al. (1999) found that blacks in concordant relationships were more likely to rate physicians as excellent and to be very satisfied with their physician. However, Saha, et al. (2003) found no effect of racial concordance on the quality of patient-physician interaction (cultural sensitivity), patient satisfaction, or the use of health care services.

While this variation in results may be due in part to differences in data sources or real changes over time, it is critical to note that it is uncertain whether these studies are clearly identifying the contribution of concordance. Specifically, they do not separately control for both physician and patient race in models that estimate the effect of concordance. Physician (patient) race is likely to be correlated with both the outcome of interest and with patient (physician) characteristics and the achievement of concordance, resulting in omitted variables bias in the estimated coefficient. Since concordance is essentially an interaction between patient and physician race, not including both of these main effects will confound the estimated effect of concordance.[5] This analysis separately identifies the contributions of physician race, patient race and concordance to disparate

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[3] Cooper-Patrick, et al. 1999; LaVeist and Nuru-Jeter 2002; LaVeist, et al. 2003; Saha, et al. 1999.

[4] Cooper, et al. 2003; Saha, et al. 2003.

[5] This has been less of a problem in the literature on the effect of gender concordance, some of which has been done using the NAMCS data. See Franks and Bertakis 2003 and Schmittdiel, et al. 2000.

outcomes in the outpatient setting, thereby separating any incremental impact of racial matching.

In addition to making methodological improvements, this analysis explores the impact of concordance on a set of clinical outcomes that have not been examined in the previous literature. Instead of using patients' subjective opinions about the quality of their care, I assess more objective provider reports of services ordered or provided at outpatient visits that are strong indicators of quality primary care and the length of those visits. By restricting my analysis to patient populations that should receive these services, these measures are reasonable indicators of the quality of care a physician provides during a visit.

Several other differences from the previous literature arise, since I am using data from the clinical encounter rather than national population surveys. I examine the effect of the race of the physician that actually interacted with the patient at the visit, instead of the race of the provider who the individual considers to be his or her usual provider. Again, while individuals' opinions about their usual provider are one component of quality care, the experiences of actual patients who access the health care system are another important measure. In these data, physician and patient race are reported by the physician office instead of by the patient, which is the case in population surveys.

Section 3: Methods

3.1 Conceptual Framework

The IOM attributes racial/ethnic differences in the quality of health care to three major factors: 1) clinical appropriateness, need, and patient preferences; 2) the operation of health care systems and the legal and regulatory climate; and 3) discrimination (Smedley, et al. 2002). Differences due to the latter two categories are defined as disparities. Identifying the role of discrimination in differences in quality of care requires controlling for both differences in clinical appropriateness and need, as well as the factors associated with the operation of the health care system, notably physician resources, patients' health insurance coverage, and variation in local practice patterns. As described above, racial/ethnic concordance between patients and physicians is hypothesized to affect quality of care by reducing discrimination. Accordingly, I seek to identify the role of concordance controlling for factors related to patient need and access to the health care system.

In an observational setting, physician-patient pairs may be concordant for various reasons, including physician preferences and selection regarding location, specialty and practice type, patient preferences and selection regarding location, sources of care and providers, as well as random matching. Concordance that arises due to geographic constraints may affect outcomes in very different ways and the patients and physicians are likely to be quite different than those in more racially mixed areas. By controlling for the geographic area in which the visit takes place, I can control for both health care system factors that contribute to disparities (i.e., regional practice variation) and for some factors that may affect reasons for, and achievement of, concordance.

Whether concordance arises due to selection or to random chance, assessing the importance of the discrimination mechanisms that operate through the physician-patient relationship requires comparisons between concordant and non-concordant pairs that are similar in other respects. This analysis tests whether racial matching has a positive effect on quality of care in the outpatient setting relative to unmatched pairs, *seen by the same provider*.

	Physician Race	
Patient Race	White	Black
White	Y_{WW}	Y_{WB}
Black	Y_{BW}	Y_{BB}

In a simplified example with two race categories, if racial concordance improves outcomes, we expect $Y_{WW} > Y_{BW}$ and $Y_{BB} > Y_{WB}$. After controlling for patient need, access to the health care system, and geographic factors, differences in outcomes are likely driven by the discrimination mechanisms, and it is the effect of concordance on these mechanisms that I aim to assess.

3.2 Data

I use a pooled sample (2001-2003) of the National Ambulatory Medical Care Survey (NAMCS), a nationally-representative sample of ambulatory care in physicians' offices stratified based on geography and physician specialty. The unit of observation is the patient visit and detailed information is provided on patient characteristics, length of the visit, diagnostic tests and procedures ordered or performed, diagnoses made, and drugs prescribed. Physician information such as specialty and practice setting are available in the public-use version of the dataset. Confidential data files that include additional physician information (race, gender, date of birth) and geographic identifiers were accessed through the National Center for Health Statistics' Research Data Center. County-level data from the 2000 Area Resource File (ARF), including the percent of the population in different racial/ethnic groups, median income and education variables, were merged with the NAMCS data (U.S. Department of Health and Human Services 2003). County-level fixed effects are used to make comparisons within relatively small geographic areas but the more detailed information from the ARF help inform which factors are correlated with high concentrations of minority patients and physicians.

I restrict the sample to primary care physicians: those whose specialty is general/family practice or internal medicine. I construct patient and physician race/ethnicity variables that take the following values: white, non-Hispanic; black, non-Hispanic; Hispanic (of any race); and Asian, non-Hispanic. Respondents who identify as another race category or whose race/ethnicity responses are missing or imputed are not included in this analysis.[6] Patient race is observed for 14,448 visits to primary care physicians, while both patient and physician race are observed for 8,160 visits. In this latter subsample, white physicians and patients each make up about 80 percent of visits, while Hispanic physicians and patients make up about 7 percent. Black patients account for 9 percent of visits while black physicians account for only 2 percent. Asian patients comprise 3 percent of visits and Asian physicians make up 13 percent.

I examine two types of outcomes in the primary care setting: preventive screening rates and the length of the visit in minutes. Preventive screening outcomes include whether blood pressure was checked, whether the physician knows if the patient uses tobacco, whether tobacco cessation counseling was ordered or occurred if the patient uses tobacco, and whether a cholesterol test was ordered or performed. The length of the visit

[6] Twenty-three percent of visits to primary care physicians have imputed or missing patient race/ethnicity.

is defined as the amount of time the physician spent with the patient. If the patient saw a nurse or other provider, the length of the visit is recorded as zero minutes.[7] Various restrictions on patient age and gender, physician specialty, and type of visit are used, depending on the outcome variable in question. For example, the analysis of preventive screenings is restricted to patient populations for which the test is indicated.[8] The length of visit analysis controls for the general purpose of the visit, whether the patient has been seen at that office before, and whether it is an initial or follow-up visit.

Table 3.1 describes patient, physician, practice and county characteristics, as well as mean outcomes, by patient race. There are no differences by race/ethnicity in patient age or gender, but black and Hispanic patients are significantly more likely to have Medicaid coverage than white patients. Similarly, there are not significant differences in physician age or gender by patient race. However, we see large and significant differences in physician race, with white physicians less likely to see minority patients, and black, Hispanic, and Asian physicians many times more likely to see patients of their own race/ethnicity than white patients. Physicians who see Hispanic and Asian patients are also much more likely to be graduates of foreign medical schools than those who see white and black patients. Weighted concordance rates are 84 percent for whites, 18 percent for blacks, 44 percent for Hispanics, and 59 percent for Asians. These rates are somewhat higher for Asians and Hispanics than those reported by Saha, et al (2003).

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[7] The length of the visit does not include the time a patient spent waiting to see the physician, time spent receiving care from someone other than the physician, or the physician's time spent preparing for a patient (reviewing the patient's medical records or test results before seeing the patient).

[8] Screening recommendations generally involve age and gender restrictions and come from the US Preventive Services Task Force, 2003 and the American Heart Association Guidelines for Primary Prevention of Cardiovascular Disease and Stroke: 2002 Update. Blood pressure screening: adults age 18+. Tobacco use screening: all adults. Tobacco cessation interventions: adults who use tobacco products. Cholesterol screening: men aged 35 years and older and women aged 45 years and older.

Table 3.1
Summary Statistics

	Total		Patient Race								N
			White		Black		Hispanic		Asian		
	mean	se	mean	se	mean	se	mean	se	mean	se	
Patient Characteristics											
age	51	0.719	52	0.724	48	1.379	49	3.741	53	3.640	8160
female	59%	0.009	58%	0.009	64%	0.021	62%	0.021	56%	0.036	8160
private health insurance	58%	0.020	59%	0.018	54%	0.038	50%	0.083	44%	0.107	7987
medicare	28%	0.014	29%	0.014	23%	0.026	23%	0.053	34%	0.090	7987
medicaid	7%	0.010	5%	0.006	15% *	0.023	18% *	0.045	14%	0.042	7987
worker's comp	1%	0.002	1%	0.002	1%	0.006	1%	0.005	1%	0.006	7987
self pay	5%	0.007	4%	0.007	5%	0.014	6%	0.018	6%	0.021	7987
no charge	0%	0.001	0%	0.001	0%	0.001	0%	0.003	0%	0.002	7987
other health insurance	1%	0.003	1%	0.003	1%	0.006	2%	0.006	2%	0.015	7987
Physician Characteristics											
age	48	0.629	48	0.630	50	1.440	49	1.590	52	1.587	8160
female	19%	0.029	19%	0.031	23%	0.055	13%	0.048	11%	0.049	8160
white	77%	0.033	84%	0.026	65% *	0.070	44% *	0.122	38% *	0.116	8160
black	2%	0.008	1%	0.004	18% *	0.066	1%	0.004	0%	0.004	8160
hispanic	7%	0.024	3%	0.010	6%	0.030	44% *	0.137	2%	0.012	8160
asian	13%	0.023	12%	0.023	12%	0.044	12%	0.060	59% *	0.120	8160
foreign medical school	30%	0.033	25%	0.031	30%	0.065	65% *	0.077	64% *	0.111	8160
patient's primary care physician	88%	0.014	88%	0.015	90%	0.023	88%	0.038	93%	0.026	7835
concordant	74%	0.023	84%	0.026	18% *	0.066	44% *	0.137	59%	0.120	8160

* p<=.05
Primary care provider visits with patient and physician race nonmissing.
Standard errors are adjusted for the survey design.

Table 3.1, continued
Summary Statistics

	Total		Patient Race: White		Black		Hispanic		Asian		N
	mean	se	mean	se	mean	se	mean	se	mean	se	
Practice Characteristics											
percent private health insurance	57%	0.020	59%	0.017	55%	0.038	46%	0.076	41%	0.099	7835
percent medicare	28%	0.013	28%	0.014	25%	0.025	29%	0.052	32%	0.083	7835
percent medicaid	7%	0.010	5%	0.006	13% *	0.024	17% *	0.040	19% *	0.043	7835
percent uninsured	5%	0.007	5%	0.007	5%	0.011	5%	0.014	4%	0.016	7835
msa	79%	0.029	76%	0.033	82%	0.058	95% *	0.026	97% *	0.015	8160
solo practice	36%	0.035	34%	0.036	33%	0.071	45%	0.110	67% *	0.096	8160
County of Visit Characteristics											
percent white	77%	0.012	80%	0.011	69% *	0.023	67% *	0.027	64% *	0.042	8140
percent urban	78%	0.019	75%	0.021	82%	0.043	94% *	0.013	95% *	0.013	8140
percent persons in poverty	12%	0.005	11%	0.003	13%	0.007	18% *	0.025	11%	0.007	8140
median household income	42,398	809	42,407	784	41,362	1,277	39,966	2,655	52,015 *	2,042	8140
percent less than 9 years edu	8%	0.006	7%	0.003	7%	0.005	15% *	0.030	9%	0.008	8140
percent foreign born	10%	0.009	8%	0.006	10%	0.017	27% *	0.033	22% *	0.023	8140
percent non-english speakers	4%	0.005	3%	0.003	3%	0.006	13% *	0.018	9% *	0.011	8140
percent single parent households	11%	0.002	10%	0.002	12% *	0.004	12% *	0.006	10%	0.004	8140
Outcomes											
blood pressure	81%	0.018	81%	0.018	84%	0.023	86%	0.051	74%	0.087	7239
ask tobacco use	80%	0.020	80%	0.018	84%	0.029	74%	0.084	82%	0.046	7239
tobacco counsel	29%	0.023	29%	0.024	31%	0.057	20%	0.068	34%	0.075	1042
cholesterol women age 45+	13%	0.011	12%	0.011	10%	0.030	17%	0.053	10%	0.043	2950
cholesterol men age 35+	18%	0.015	16%	0.014	30% *	0.040	24%	0.069	12%	0.047	2557
time with physician	18	0.460	18	0.467	18	1.033	17	1.870	17	0.746	7509

* p<=.05
Primary care provider visits with patient and physician race nonmissing.
Standard errors are adjusted for the survey design.

In terms of practice characteristics, minority patients see physicians with significantly higher shares of Medicaid patients. Hispanic and Asian patients are more likely to see physicians in urban areas and Asians are more likely to see physicians in solo practice. We can see large and significant differences in county characteristics by patient race, particularly for Hispanics and Asians. Their visits are more likely to be in counties that are less white, more urban, and have more foreign-born and non-English speaking residents. Somewhat surprisingly, I find few significant differences in unadjusted outcomes across patient race, with only black men being significantly more likely to receive a cholesterol test.

3.3 Statistical Methods

This analysis aims to understand how racial/ethnic concordance contributes to differences in outcomes, particularly among the subgroup of physicians who see a substantial share of both white and minority patients. In order to assess the importance of the within-provider mechanisms, I control for geographic characteristics and restrict my analysis to providers who see both white and minority patients, using either physician fixed effects or the subsample of physicians with a mixed patient panel (25-75% white vs. minority). I use regression analysis to identify the incremental contribution of physician-patient racial/ethnic concordance on primary care outcomes while controlling for confounding variables including patient and physician race.[9] Ordinary least-squares (OLS) regression is used for the length of visit analysis and linear probability models are

[9] Note that the effect of concordance in observational data includes both the effect of "random matching" and the effect of individuals who desire concordance achieving it (selection).

used for the binary primary care screening outcomes. The analysis also accounts for the complex sampling design of the survey.

The estimating equation for the analysis of preventive screening outcomes is:

$$Y_{pmt} = \alpha + \beta_1(\text{race}_p) + \beta_2(\text{race}_m) + \beta_3(\mathbf{1}{*}\text{concordant}_{pm}) + \beta_4(X_p) + \beta_5(X_m) + \beta_6(\text{county}_m) + \beta_7(\text{year}_t) + \varepsilon_{pmt} \quad (1)$$

where Y_{pmt} is a visit-level outcome in year t, race_p is a set of indicator variables for the patient's race/ethnicity, race_m is a set of indicator variables for the provider's race/ethnicity, concordant_{pm} is an indicator variable equal to one if the physician-patient pair is concordant, X_p is a set of patient-level control variables (age, gender, insurance coverage), X_m is a set of physician-level controls (age and gender), county_m controls for the county where the visit takes place, and year_t is a set of year fixed-effects. β_3, the main coefficient of interest, measures the incremental contribution of concordance to the outcome, controlling for the main effects of patient and physician race. The impact of concordance, however, is a function of patient race, physician race and the incremental contribution of the match. Returning to the table in Section 3.1, the coefficients from the simplified regression:

$$Y_{pmt} = \alpha + \beta_1(\text{black}_p) + \beta_2(\text{black}_m) + \beta_3(\mathbf{1}{*}\text{concordant}_{pm}) + \varepsilon_{pmt} \quad (2)$$

would be distributed as follows.

	Physician Race	
Patient Race	White	Black
White	β_3	β_2
Black	β_1	$\beta_1 + \beta_2 + \beta_3$

For a black patient, the impact of changing from a white physician to a black physician is $\beta_2 + \beta_3$ – the average effect of a black physician plus the effect of the match. A key contribution of this analysis is to measure the contribution of these factors separately. Furthermore, since β_3 is the average effect of concordance across patient race, I also include interaction terms to allow the effect of concordance to vary.

$$Y_{pmt} = \alpha + \beta_1(race_p) + \beta_2(race_m) + \beta_3(\mathbf{1}*concordant_{pm}) + \beta_4(race_p*concordant_{pm}) + \beta_5(X_p) + \beta_6(X_m) + \beta_7(county_m) + \beta_8(year_t) + \varepsilon_{pmt} \quad (3)$$

I use two models to examine the contribution of concordance within the set of primary care providers that see both white and minority patients. First I run equations 1 and 3 on the subsample of physicians whose patient panel consists of 25-75 percent white patients (and correspondingly 75-25 percent minority patients). In this model, I am still able to estimate the importance of physician characteristics that are constant across patients, such as race and gender. Second, I run a similar model that includes physician fixed effects.

$$Y_{pmt} = \alpha + \beta_1(race_p) + \beta_2(MD_m) + \beta_3(\mathbf{1}*concordant_{pm}) + \beta_4(X_p) + \varepsilon_{pmt} \quad (4)$$

The physician fixed-effects, MD_m, control for the contributions of physician-level variables that are constant across patients (race, age, gender, county and year) as well as unobserved characteristics. In this model, β_3 is estimated based only on physicians who are concordant with some patients but not others.

The initial model for the analysis of visit length is:

$$Y_{pmt} = \alpha + \beta_1(race_p) + \beta_2(race_m) + \beta_3(\mathbf{1}*concordant_{pm}) + \beta_4(X_p) + \beta_5(X_m) + \beta_6(county_m) + \beta_7(specialty_m) + \beta_8(Z_{pm}) + \beta_9(year_t) + \varepsilon_{pmt} \quad (5)$$

where Y_{pmt} is the length of the visit in minutes, $specialty_m$ is a set of fixed-effects for physician specialty, Z_{pm} is a set of visit-level control variables (whether the patient has been seen before, the reason for the visit) and the other variables are the same as described above. Again, β_3 is the main coefficient of interest and interaction terms of concordance and patient race are also included. As with the analysis of preventive screening outcomes, I run subsequent models on the mixed patient panel subsample and then on the entire sample using physician fixed effects.

Section 4: Results

4.1 Distribution of Patients Across Primary Care Physicians

Before examining the contribution of racial/ethnic concordance among physicians who see both white and minority patients, I investigate the extent to which this occurs. To the degree that these patient populations see different providers, I examine how they differ in terms of physician, practice and geographic characteristics.

A brief examination of the distribution of patients across physicians indicates quite vividly the importance of examining concordance in the within-provider context. Figure 3.1 shows the distribution of patients across primary care physicians ranked in ascending order of the percent of their patient panel made up of minority patients. While white patients are essentially equally distributed across physicians, minority patients are highly concentrated among a smaller group of providers who see a high percentage of minority patients. Thirty six percent of physicians see 80 percent of black patients, while the corresponding figures for Hispanic and Asian patients are 23 and 29 percent of physicians, respectively. This picture illustrates that white and minority patients generally do not see the same primary care providers and suggests that across-provider variation may be an important source of disparities in the outpatient setting. However, a one-way analysis-of-variance shows that across-provider variation accounts for about 45 percent of the variance in blood pressure, asking about tobacco use and the length of the visit, 35 percent of the variance in tobacco use counseling, and 25 percent of the variance in rates of cholesterol screening. Variation across patients within providers accounts for 55, 65 and 75 percent, respectively.

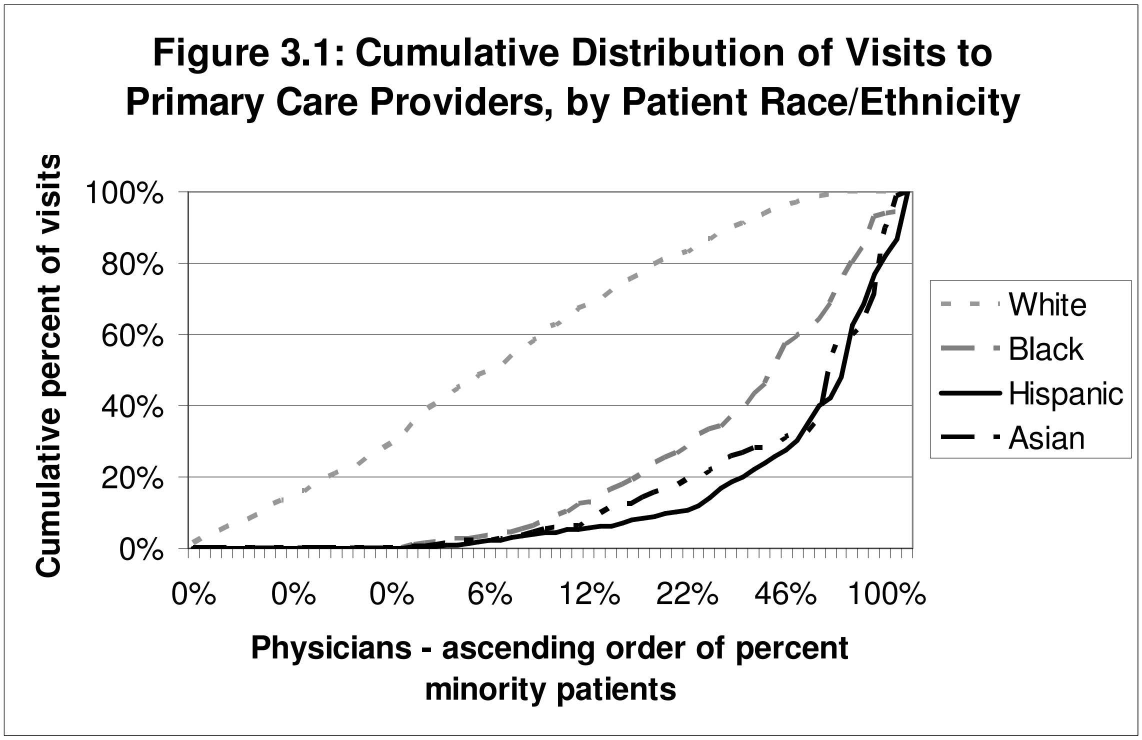
Figure 3.1: Cumulative Distribution of Visits to Primary Care Providers, by Patient Race/Ethnicity
Cumulative percent of visits
100%
80%
60%
40%
20%
0%
0%
0%
0%
6%
12%
22%
46%
100%
Physicians - ascending order of percent minority patients
White
Black
Hispanic
Asian

Table 3.2
Physician, Practice and Geographic Characteristics by Composition of Patient Panel

	Total	**Racial Composition of Patient Panel** >75% white	25%-75% white	<25% white
Physician and practice characteristics				
Age	48.3	47.6	49.8	50.9
	0.63	0.70	2.36	1.79
Female	19%	19%	23%	12%
	0.03	0.03	0.07	0.06
White	77%	86%	72%	26% *
	0.03	0.03	0.07	0.12
Black	2%	0%	5%	11%
	0.01	0.00	0.03	0.06
Hispanic	7%	2%	11%	34% *
	0.02	0.01	0.06	0.13
Asian	13%	12%	11%	28%
	0.02	0.02	0.04	0.11
Patient's primary care physician	88%	88%	88%	93%
	0.01	0.02	0.03	0.03
Solo practice	36%	34%	34%	53%
	0.03	0.04	0.08	0.12
Private practice (solo or group)	89%	91%	85%	84%
Free-standing clinic	6%	7%	4%	4%
Federally-qualified health center	1%	0%	7%	3%
Non-federal government clinic	2%	1%	2%	5%
Geographic characteristics				
CA	10%	8%	14%	25%
FL	3%	2%	3%	15%
LA	2%	1%	3%	12%
NY	3%	2%	3%	7%
TX	11%	9%	22%	15%
5-state total	29%	22%	45%	74%
MSA	79%	75%	83%	98% *
	3%	3%	7%	2%

* p<=.05
Primary care provider visits with patient and physician race nonmissing.
Patient panel composition is percent white relative to percent minority.
Standard errors are below means and are adjusted for the survey design.

Table 3.2, continued
Physician, Practice and Geographic Characteristics by Composition of Patient Panel

		Racial Composition of Patient Panel		
	Total	**>75% white**	**25%-75% white**	**<25% white**
County characteristics				
% pop white	77%	81%	69% *	62% *
	1.20	1.15	0.02	2.87
% pop urban	78%	73%	86% *	96% *
	1.75	2.31	0.04	1.57
% persons in poverty	12%	11%	14% *	17% *
	0.52	0.29	0.01	1.81
median hhd income	$42,398	$42,649	$41,954	$41,276
	$780	$854	$1,661	$2,441
% persons <9yrs edu	8%	7%	11% *	13% *
	0.57	0.31	0.01	2.15
% foreign born	10%	7%	16% *	27% *
	0.01	0.01	0.02	0.04
% non-english speaking	4%	3%	7% *	12% *
	0.00	0.00	0.01	0.02
% single-parent hhds	11%	10%	12% *	12% *
	0.00	0.00	0.00	0.01

* p<=.05
Primary care provider visits with patient and physician race nonmissing.
Patient panel composition is percent white relative to percent minority.
Standard errors are below means and are adjusted for the survey design.

To understand whether providers who have mixed patient panels differ from those who do not, I group physicians by the composition of their patient panel: more than 75 percent white, 25-75 percent white, and less than 25 percent white (more than 75 percent minority).[10] Table 3.2 shows that those physicians with many minority patients are much more likely to be members of minority groups themselves. Eleven percent are black, 34 percent Hispanic and 28 percent Asian, rates which are 2-5 times their share of the total sample of physicians. Practice characteristics do not vary significantly across the different patient panel groups, but geographic and county characteristics are quite different. Outpatient visits to physicians with many minority patients are highly concentrated in just a few states, notably California, Florida, Louisiana, New York and Texas, and are more likely to be in urban locations. Physicians with mixed or highly minority patient panels are in counties with higher percentages of the population in urban areas, in poverty, with less than 9 years of education, foreign-born, non-English speaking, and in single-parent households. Interestingly, median household income is not significantly different.

This combined evidence showing that minority patients generally see a select group of primary care physicians, and that these physicians differ both in their individual and geographic characteristics, supports previous work that highlights the importance of geography in racial/ethnic disparities in health care.[11] Not only are physicians who see mostly minority patients in more disadvantaged geographic areas, but evidence from the 2003 NAMCS survey suggests that they have more difficulty referring patients to

[10] This is based on the actual composition of physicians' patients in the data, rather than on physician self-report of the racial composition of their patients (as in Komaromy, et al. 1996 and Stinson and Thurston 2002).

[11] Bach, et al. 2004, Chandra and Skinner 2003, and Stinson and Thurston 2002.

specialists than physicians who see mostly white patients, controlling for type of health insurance coverage.[12] Given the geographic concentration of minority patients and physicians and the differences in socioeconomic and health system factors, the factors driving concordance, and its effects, are likely to be different than in the context where physicians see both white and minority patients. It is with this in mind that we turn to examining the impacts of concordance, focusing on physicians who see both white and minority patients and within-provider variation.

4.2 Racial/Ethnic Concordance and Quality of Care

Table 3.3 shows differences in preventive screening outcomes by patient race, after controlling for gender, age, health insurance status, and county fixed-effects. In the total sample, among men age 35 and higher blacks are 13 percentage points more likely to receive a cholesterol test than whites (column g). Restricting to providers with mixed patient panels, black adults are 9 percentage points more likely to be asked about tobacco use (column d). Health insurance coverage is consistently an important predictor of preventive screenings. Visit length is significantly shorter if the patient has been seen before. The results in this table suggest that, to the extent we see discrimination by patient race in these outpatient outcomes, it is likely to be the result of statistical discrimination toward black patients as opposed to prejudice.[13]

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[12] Twenty percent of physicians with high minority patient panels report difficulty referring uninsured patients for specialty consultations, compared to 10% with high white patient panels. Corresponding percentages are 27% vs. 6% for privately insured patients, 11% vs. 4% for Medicare patients, and 19% vs. 12% for patients with Medicaid. These referral questions were only asked in the 2003 NAMCS survey.

[13] Statistical discrimination in the health care context refers to increased clinical uncertainty (physicians' differential beliefs about the underlying illness severity of minority groups or physicians' difficulty reading minority patients' signals) and stereotyping (physicians' belief that minority patients are less likely to comply with treatment) (Balsa and McGuire 2003).

Table 3.3
Differences in Preventive Screening Outcomes and Length of Visit by Patient Race

	Blood Pressure Check Adults Age 18+		Ask Tobacco Use Adults Age 18+	
	Total Sample a	**25-75% Sample b**	**Total Sample c**	**25-75% Sample d**
Black	0.0333	0.0025	0.0226	0.0884 *
	0.0255	0.0429	0.0213	0.0355
Hispanic	0.0473	-0.0064	0.0050	0.0084
	0.0371	0.0546	0.0442	0.0436
Asian	-0.0519	0.0809	0.0810	0.0623
	0.0764	0.0571	0.0633	0.0474
Female	-0.0024	0.0605 *	0.0177	-0.0235
	0.0096	0.0247	0.0114	0.0201
Medicare	0.0001	-0.0140	0.0133	0.0304
	0.0212	0.0518	0.0229	0.0505
Medicaid	-0.0120	0.0230	0.0203	-0.0135
	0.0285	0.0657	0.0220	0.0557
Worker's comp	-0.0972	0.2547 *	-0.0916	-0.5705 ***
	0.0743	0.1089	0.0591	0.1125
Self pay	-0.1700 *	-0.0448	0.0263	0.0404
	0.0800	0.1223	0.0388	0.0777
No charge	-0.0511	-0.2795	0.0974 *	0.0703
	0.1461	0.2980	0.0463	0.0622
Other	0.0694	-0.0276	0.0468	0.0252
	0.0391	0.1189	0.0615	0.0847
2002	0.0047	0.0377	0.0646	0.1223
	0.0580	0.1355	0.0512	0.1056
2003	-0.0025	0.2198	0.0554	0.2463 *
	0.0561	0.1222	0.0563	0.1020
_cons	0.8894 ***	0.5457 **	0.7590 ***	0.4878 **
	0.0787	0.2063	0.0862	0.1634
r2	0.2949	0.1174	0.3250	0.1737
N_sub	7238	1862	7122	1815

* p<=.05, ** p<=.01, *** p<=.001
Standard errors are below the coefficients and are adjusted for the survey design. All models include patient age and county fixed effects except the 2575 sample regressions for blood pressure check and asking about tobacco use. These two models do not include county fixed effects due to problems with variance estimation. Omitted categories are white, male, private health insurance, family/general medicine, acute problem, 2001.

Table 3.3, continued
Differences in Preventive Screening Outcomes and Length of Visit by Patient Race

	Tobacco Counseling Adults Age 18+		Cholesterol Check Men Age 35+	
	Total Sample e	**25-75% Sample f**	**Total Sample g**	**25-75% Sample h**
Black	0.0278	0.0125	0.1338 ***	0.1253 **
	0.0656	0.0921	0.0382	0.0434
Hispanic	-0.1584	-0.0421	0.0553	0.1187
	0.1321	0.0971	0.0459	0.0646
Asian	0.0493	0.2604	0.0017	0.2362 *
	0.1124	0.1591	0.0510	0.1149
Female	-0.0555	-0.1776		
	0.0363	0.1058		
Medicare	-0.1064 *	0.2709	-0.1032 ***	-0.1637 **
	0.0467	0.1819	0.0301	0.0604
Medicaid	0.1416 *	0.2564	-0.0822 *	-0.0098
	0.0644	0.1312	0.0358	0.0807
Worker's comp	-0.1611		-0.2122 ***	-0.1433
	0.0917		0.0488	0.1032
Self pay	0.0490	0.0592	-0.1215 ***	-0.1041
	0.0691	0.1664	0.0361	0.1040
No charge	0.2243	0.5055 *	-0.1586 *	-0.1281 *
	0.1158	0.2281	0.0739	0.0601
Other	-0.0907	0.4744	-0.0160	-0.0996
	0.0870	0.2885	0.0714	0.0885
2002	0.0238	0.2628	0.0249	-0.0044
	0.0669	0.3422	0.0396	0.0685
2003	0.0564	0.5278 *	0.0369	0.1673 *
	0.0744	0.2062	0.0320	0.0652
_cons	0.3830 *	0.5232	-0.2076 *	-0.0230
	0.1506	0.2894	0.0825	0.1267
r2	0.3719	0.7754	0.2201	0.4253
N_sub	1028	157	2510	364

* p<=.05, ** p<=.01, *** p<=.001
Standard errors are below the coefficients and are adjusted for the survey design. All models include patient age and county fixed effects except the 2575 sample regressions for blood pressure check and asking about tobacco use. These two models do not include county fixed effects due to problems with variance estimation. Omitted categories are white, male, private health insurance, family/general medicine, acute problem, 2001.

Table 3.3, continued
Differences in Preventive Screening Outcomes and Length of Visit by Patient Race

	Cholesterol Check Women Age 45+		Time with Physician All Ages	
	Total Sample	**25-75% Sample**	**Total Sample**	**25-75% Sample**
	i	**j**	**k**	**l**
Black	-0.0320	0.0281	0.300	0.921
	0.0299	0.0509	0.492	0.644
Hispanic	-0.0233	-0.0335	0.594	-0.013
	0.0328	0.0606	0.842	0.552
Asian	-0.0243	-0.0191	-1.293	1.770
	0.0467	0.0991	0.913	0.990
Female			-0.212	0.235
			0.349	0.508
Medicare	-0.0263	-0.0739	0.085	1.133
	0.0191	0.0422	0.549	0.791
Medicaid	-0.0155	-0.0231	-0.110	-0.141
	0.0310	0.0630	0.733	0.793
Worker's comp	0.0468	-0.1801	0.353	2.936
	0.1191	0.1540	1.363	1.573
Self pay	-0.0875 ***	-0.1158 *	-2.009	2.340
	0.0207	0.0464	1.727	1.755
No charge	-0.2199 ***		1.969	-1.129
	0.0498		1.404	2.059
Other	-0.1119	-0.0695	-0.838	-2.370
	0.0936	0.1244	1.541	1.988
Seen before			-5.596 ***	-5.423 **
			1.374	1.957
Chronic, routine			1.112	2.078 *
			0.570	0.807
Chronic, flare up			0.831	1.681
			0.509	1.139
Follow-up			5.014 **	-3.857 *
			1.840	1.584
Episode			-0.253	-0.809
			0.420	0.622
Internal medicine			1.563	-4.751 *
			0.990	2.161
2002	0.0293	0.1362 *	-1.306	-0.110
	0.0341	0.0637	2.079	2.213
2003	0.0480	0.1668 **	-0.315	1.411
	0.0372	0.0570	2.527	2.301
_cons	-0.0325	0.0107	14.444 ***	7.192
	0.0451	0.2241	3.339	4.820
r2	0.1595	0.3193	0.3312	0.457
N_sub	2891	395	4554	864

* p<=.05, ** p<=.01, *** p<=.001
Standard errors are below the coefficients and are adjusted for the survey design. All models include patient age and county fixed effects except the 2575 sample regressions for blood pressure check and asking about tobacco use. These two models do not include county fixed effects due to problems with variance estimation. Omitted categories are white, male, private health insurance, family/general medicine, acute problem, 2001.

Table 3.4
Contribution of Racial/Ethnic Concordance to Differences in Preventive Screening Outcomes and Length of Visit, Providers with Mixed Patient Panels (25-75% white)

	Blood Pressure Check Adults Age 18+			**Ask Tobacco Use Adults Age 18+**		
	a	**b**	**c**	**d**	**e**	**f**
Black patient	0.0410	0.0306	0.0210	0.0245	0.0378	0.0382
	0.0330	0.0322	0.0458	0.0334	0.0381	0.0322
Hispanic patient	-0.0163	-0.0240	-0.0277	-0.0099	-0.0001	0.0412
	0.0223	0.0219	0.0487	0.0216	0.0256	0.0404
Asian patient	0.0422	0.0358	0.0725	-0.0462	-0.0376	0.0064
	0.0509	0.0506	0.0729	0.0555	0.0604	0.0754
Black physician	0.5949 ***	0.5925 ***	0.5538 ***	0.2632 ***	0.2668 ***	0.1973 **
	0.0447	0.0456	0.0571	0.0389	0.0384	0.0642
Hispanic physician	0.0088	-0.0046	-0.0099	0.3429 **	-0.0020	0.0171
	0.0454	0.0453	0.0556	0.1229	0.0434	0.0470
Asian physician	-0.4774 ***	-0.4805 ***	-0.4403 ***	0.5944 ***	0.5988 ***	0.6543 ***
	0.0306	0.0316	0.0508	0.0390	0.0402	0.0559
Concordant		-0.0164	-0.0190		0.0211	0.0461
		0.0178	0.0450		0.0260	0.0399
Black*Concordant			0.0788			0.1644
			0.0776			0.1022
Hispanic*Concordant			0.0161			-0.1255
			0.0872			0.0729
Asian*Concordant			-0.1386			-0.1400
			0.1217			0.1575
_cons	1.5075 ***	1.5215 ***	1.5239 ***	-0.1722	-0.1917	-0.2423 *
	0.1209	0.1290	0.1379	0.1089	0.1187	0.1019
r2	0.6660	0.6662	0.6671	0.6841	0.6844	0.6873
N_sub	1092	1092	1092	1069	1069	1069

* p<=.05, ** p<=.01, *** p<=.001
Standard errors are below the coefficients and are adjusted for the survey design.
All models contain patient sex, health insurance coverage, and age fixed effects; physician sex and age fixed effects; and year and county fixed effects. The time with physician models also include controls for whether the patient was seen before, the reason for the visit, and physician specialty. Omitted categories are white, male, private health insurance, family/general medicine, acute problem, 2001.

Table 3.4, continued
Contribution of Racial/Ethnic Concordance to Differences in Preventive Screening Outcomes and Length of Visit, Providers with Mixed Patient Panels (25-75% white)

	Tobacco Counseling Adults Age 18+			Cholesterol Check Men Age 35+		
	g	h	i	j	k	l
Black patient	-0.0320	0.0867	0.2519	0.1336 **	0.1255 *	-0.0242
	0.0835	0.1086	0.1785	0.0437	0.0604	0.0916
Hispanic patient	0.0504	0.1507	0.3946	0.1122	0.1043	-0.0930
	0.1001	0.1507	0.2028	0.0630	0.0840	0.1206
Asian patient	0.1962	0.2618	0.3457	0.2701 *	0.2666 *	0.1325
	0.1818	0.1705	0.2904	0.1187	0.1152	0.1870
Black physician	-0.2644	-0.2259	-0.1507	0.1713	0.1672	-0.0476
	0.2358	0.2127	0.2089	0.1631	0.1644	0.2156
Hispanic physicia	-1.0860 ***	-1.0121 ***	-0.7741	0.2103 *	0.2028 *	0.0043
	0.2859	0.2965	0.4302	0.0952	0.1012	0.1061
Asian physician	-1.4048 ***	-1.3291 ***	-1.2186 ***	0.2020	0.1970	0.0749
	0.2163	0.2190	0.3166	0.1324	0.1336	0.1763
Concordant		0.1571	0.3453		-0.0144	-0.2078 *
		0.1373	0.2515		0.0627	0.1011
Black*Concordant			-0.9942			0.3920 *
			0.5158			0.1915
Hispanic*Concordant			-1.4054 ***			0.4355
			0.3826			0.2733
Asian*Concordant			-0.1455			0.2641
			0.7181			0.2652
_cons	1.2523 ***	1.2297 ***	0.9419 *	-0.9528 **	-0.9484 **	-0.7634 *
	0.3526	0.3240	0.3933	0.3250	0.3253	0.3803
r2	0.8079	0.8107	0.8216	0.4602	0.4604	0.4662
N_sub	157	157	157	364	364	364

* p<=.05, ** p<=.01, *** p<=.001
Standard errors are below the coefficients and are adjusted for the survey design.
All models contain patient sex, health insurance coverage, and age fixed effects; physician sex and age fixed effects; and year and county fixed effects. The time with physician models also include controls fo whether the patient was seen before, the reason for the visit, and physician specialty. Omitted categories are white, male, private health insurance, family/general medicine, acute problem, 2001.

Table 3.4, continued
Contribution of Racial/Ethnic Concordance to Differences in Preventive Screening Outcomes and Length of Visit, Providers with Mixed Patient Panels (25-75% white)

	Cholesterol Check Women Age 45+			Time with Physician All Ages		
	m	n	o	p	q	r
Black patient	0.0359	0.1044	0.0274	1.063	1.835 **	1.541
	0.0562	0.0911	0.0457	0.659	0.686	0.864
Hispanic patient	-0.0184	0.0139	-0.0900	-0.036	0.348	0.074
	0.0638	0.0682	0.0679	0.561	0.624	1.040
Asian patient	-0.0341	0.0538	0.0033	1.501	1.824	1.330
	0.1153	0.1261	0.1282	0.936	0.981	1.054
Black physician	0.0126	0.1426	-0.0209	-14.353 ***	-14.570 ***	-15.063 ***
	0.0844	0.1257	0.1450	1.442	1.469	2.255
Hispanic physicia	-0.2354 **	-0.1449	-0.2074 **	6.025 ***	6.802 ***	6.643 ***
	0.0843	0.0775	0.0707	1.286	1.196	1.100
Asian physician	-0.3488 ***	-0.3670 ***	-0.2863 **	8.147 ***	8.633 ***	8.278 ***
	0.0954	0.1019	0.0924	1.401	1.415	1.546
Concordant		0.1126	0.0202		1.033	0.693
		0.0813	0.0538		0.689	1.135
Black*Concordant			0.1631			0.971
			0.4207			2.410
Hispanic*Concordant			0.2615			0.528
			0.1433			2.378
Asian*Concordant			-0.1911			1.280
			0.1571			2.860
_cons	0.3653 **	0.2309	0.3598 **	25.808 ***	25.148 ***	25.588 ***
	0.1378	0.1387	0.1247	3.912	3.791	4.126
r2	0.3300	0.3401	0.3440	0.4823	0.4840	0.4841
N_sub	395	395	395	864	864	864

* p<=.05, ** p<=.01, *** p<=.001
Standard errors are below the coefficients and are adjusted for the survey design.
All models contain patient sex, health insurance coverage, and age fixed effects; physician sex and age fixed effects; and year and county fixed effects. The time with physician models also include controls for whether the patient was seen before, the reason for the visit, and physician specialty. Omitted categories are white, male, private health insurance, family/general medicine, acute problem, 2001.

Table 3.4 presents estimates of the contribution of concordance based on the subsample of providers with mixed patient panels. After controlling for physician race, sex and age, black patients are no longer significantly more likely to be asked about tobacco use (column d) but black men remain much more likely to have a cholesterol screening (column j). Across the board, we can see that physician race is a much more important contributor to outcomes than patient race or concordance itself. Black physicians are 59 percentage points more likely than white physicians to check blood pressure while Asian physicians are 48 percentage points less likely (column a). Black, Hispanic and Asian physicians are 26, 34 and 59 percentage points more likely to screen for tobacco use, respectively (column d), and Hispanic and Asian physicians are 24 and 35 percentage points less likely to check women's cholesterol (column m).

For none of these outcomes does racial/ethnic concordance have a statistically significant incremental effect. Only for cholesterol testing among men does concordance appear to matter. Black concordant pairs are 39 percentage points more likely to screen cholesterol for men than white concordant pairs and Hispanic concordant pairs are much less likely to engage in tobacco cessation counseling. In terms of the length of the visit, black physician's visits are 14 minutes shorter than white physicians and Hispanic's and Asian's are 6 and 8 minutes longer, respectively. There is no significant incremental effect of concordance. These results highlight the importance of measuring the contribution of physician race separately from that of concordance and demonstrate why in some cases (i.e., Asian's blood pressure check and Asian women's cholesterol testing) minority patients may actually do worse with a concordant physician.

Table 3.5 presents results from estimating equation 4, which includes physician fixed effects. Controlling for factors that are fixed across each physician's patients (including physician race, gender, specialty, and geographic location), black men are 11 percentage points more likely to have cholesterol screened than white men (column j) and Asian patient have visits that are about 1 minute shorter than whites (column p). Again, concordance does not appear to have a large impact generally speaking. Black concordant pairs are much less likely to engage in counseling about tobacco use and Hispanic concordant pairs are more likely to screen men for cholesterol.

Racial/ethnic concordance increases cholesterol screening rates by 2-3 times among black and Hispanic men and appears to have a negative impact on counseling regarding tobacco use for blacks and Hispanics. For the other primary care outcomes included in this analysis, I generally find it is not a significant contributor. Table 3.6 shows the effect size bounds based on the 95 percent confidence intervals of the estimated coefficients from the models with physician fixed effects. Averaged over patient race (row 1), I can rule out positive effects of concordance of larger than 10 percent for blood pressure screening, tobacco use screening and the length of the visit. Broken out by patient race (rows 2-5), the upper bounds of positive effects on these outcomes are still relatively small. For rates of cholesterol screening among women, I am not able to rule out a large positive role of concordance.

Table 3.5
Contribution of Racial/Ethnic Concordance to Differences in Preventive Screening Outcomes and Length of Visit, Physician Fixed Effects

	Blood Pressure Check Adults Age 18+			Ask Tobacco Use Adults Age 18+		
	a	b	c	d	e	f
Black patient	0.0305	0.0265	0.0083	0.0095	0.0112	0.0042
	0.0242	0.0288	0.0422	0.0196	0.0243	0.0312
Hispanic patient	0.0035	0.0000	-0.0073	0.0252	0.0266	0.0268
	0.0168	0.0179	0.0420	0.0271	0.0239	0.0350
Asian patient	-0.0030	-0.0053	-0.0208	-0.0049	-0.0039	0.0054
	0.0391	0.0375	0.0567	0.0323	0.0317	0.0510
Concordant		-0.0054	-0.0210		0.0023	-0.0001
		0.0170	0.0424		0.0201	0.0374
Black*Concordant			0.1359			0.1033
			0.1097			0.1246
Hispanic*Concordant			-0.0090			-0.0171
			0.0853			0.0641
Asian*Concordant			0.0291			-0.0312
			0.0811			0.0797
_cons	0.8712 ***	0.8755 ***	0.8862 ***	0.8624 ***	0.8606 ***	0.8609 ***
	0.0522	0.0542	0.0659	0.0645	0.0643	0.0707
r2	0.4627	0.4628	0.4630	0.4728	0.4728	0.4729
N_sub	7238	7238	7238	7122	7122	7122

* p<=.05, ** p<=.01, *** p<=.001
Standard errors are below the coefficients and are adjusted for the survey design.
All models contain patient sex, health insurance coverage, and age and year fixed effects. The time with physician models also include controls for whether the patient was seen before and the reason for the visit (county fixed effects and physician specialty are included in the physician fixed effect). Omitted categories are white, male, private health insurance, family/general medicine, acute problem, 2001.

Table 3.5, continued
Contribution of Racial/Ethnic Concordance to Differences in Preventive Screening Outcomes and Length of Visit, Physician Fixed Effects

	Tobacco Counseling Adults Age 18+			Cholesterol Check Men Age 35+		
	g	**h**	**i**	**j**	**k**	**l**
Black patient	0.0162	0.0416	0.1382	0.1064 *	0.0878	-0.0572
	0.0738	0.1129	0.1240	0.0447	0.0552	0.1055
Hispanic patient	-0.1272	-0.1045	0.0895	0.0543	0.0383	-0.1569
	0.1389	0.1739	0.1494	0.0493	0.0583	0.1024
Asian patient	0.0958	0.1043	0.0937	0.0727	0.0631	-0.1231
	0.1524	0.1565	0.2101	0.0603	0.0598	0.1247
Concordant		0.0318	0.1480		-0.0254	-0.2068 *
		0.0965	0.1232		0.0391	0.1056
Black*Concordant			-0.9432 ***			0.1602
			0.2864			0.2532
Hispanic*Concordant			-0.8879			0.5062 **
			0.4763			0.2002
Asian*Concordant			0.1403			0.3100
			0.3811			0.1789
_cons	-0.3797 *	-0.4098	-0.5464 *	0.5291 ***	0.5506 ***	0.7450 ***
	0.1912	0.2133	0.2200	0.0745	0.0800	0.1178
r2	0.5030	0.5030	0.5082	0.2849	0.2851	0.2883
N_sub	1028	1028	1028	2510	2510	2510

* p<=.05, ** p<=.01, *** p<=.001
Standard errors are below the coefficients and are adjusted for the survey design.
All models contain patient sex, health insurance coverage, and age and year fixed effects. The time with physician models also include controls for whether the patient was seen before and the reason for the visit (county fixed effects and physician specialty are included in the physician fixed effect). Omitted categories are white, male, private health insurance, family/general medicine, acute problem, 2001.

Table 3.5, continued
Contribution of Racial/Ethnic Concordance to Differences in Preventive Screening Outcomes and Length of Visit, Physician Fixed Effects

	Cholesterol Check Women Age 45+			Time with Physician All Ages		
	m	n	o	p	q	r
Black patient	-0.0340	-0.0251	-0.0633	0.499	0.836	0.8609
	0.0348	0.0571	0.0822	0.566	0.607	0.6260
Hispanic patient	0.0150	0.0215	-0.0225	0.392	0.632	0.9135
	0.0418	0.0509	0.0843	0.625	0.749	0.9752
Asian patient	-0.0269	-0.0210	-0.0871	-1.232 *	-1.092 *	-1.3028
	0.0734	0.0683	0.0972	0.565	0.537	0.9243
Concordant		0.0121	-0.0359		0.430	0.5071
		0.0487	0.0793		0.377	0.7569
Black*Concordant			0.0540			1.1671
			0.3787			2.0628
Hispanic*Concordant			0.1136			-1.0219
			0.1642			1.2924
Asian*Concordant			0.1425			0.7295
			0.1707			1.4783
_cons	0.0065	-0.0051	0.0418	15.724 ***	15.332 ***	15.2610 ***
	0.0535	0.0702	0.0920	1.270	1.251	1.3924
r2	0.2163	0.2164	0.2167	0.4885	0.4885	0.4886
N_sub	2891	2891	2891	5867	5867	5867

* p<=.05, ** p<=.01, *** p<=.001
Standard errors are below the coefficients and are adjusted for the survey design.
All models contain patient sex, health insurance coverage, and age and year fixed effects. The time with physician models also include controls for whether the patient was seen before and the reason for the visit (county fixed effects and physician specialty are included in the physician fixed effect). Omitted categories are white, male, private health insurance, family/general medicine, acute problem, 2001.

Table 3.6
Confidence Intervals and Bounds on the Effect Size of Racial/Ethnic Concordance
Physician Fixed Effects Models

	Blood Pressure Check Adults Age 18+		**Ask Tobacco Use Adults Age 18+**		**Tobacco Counseling Adults Age 18+**	
Total						
Mean	81%		80%		29%	
CI Concordant (total)	-0.0794	0.0688	-0.0375	0.0420	-0.1591	0.2227
Bounds of effect size (%)	-10%	8%	-5%	5%	-55%	77%
By Patient Race						
Mean	81%		80%		29%	
CI Concordant (white)	-0.1048	0.0628	-0.0739	0.0736	-0.0957	0.3918
Bounds of effect size (%)	-13%	8%	-9%	9%	-33%	136%
Mean	84%		84%		31%	
CI Black*Concordant	-0.0805	0.3524	-0.1425	0.3491	-1.5099	-0.3766
Bounds of effect size (%)	-10%	42%	-17%	42%	-489%	-122%
Mean	86%		74%		20%	
CI Hispanic*Concordant	-0.1773	0.1593	-0.1435	0.1094	-1.8302	0.0543
Bounds of effect size (%)	-21%	19%	-19%	15%	-909%	27%
Mean	74%		82%		34%	
CI Asian*Concordant	-0.1309	0.1892	-0.1885	0.1260	-0.6136	0.8942
Bounds of effect size (%)	-18%	25%	-23%	15%	-181%	264%

Mean rates are based on the sample for which both patient and physician race are observed (Table 1). Confidence intervals come from the coefficients from the models that include physician fixed effects presented in Table 3.5.

Table 3.6, continued
Confidence Intervals and Bounds on the Effect Size of Racial/Ethnic Concordance
Physician Fixed Effects Models

	Cholesterol Check Men Age 35+		**Cholesterol Check Women Age 45+**		**Time with Physician All Ages**	
Total						
Mean	18%		13%		18	
CI Concordant (total)	-0.1025	0.0518	-0.0841	0.1083	-0.3140	1.1743
Bounds of effect size (%)	-58%	29%	-67%	86%	-2%	7%
By Patient Race						
Mean	16%		12%		18	
CI Concordant (white)	-0.4152	0.0016	-0.1925	0.1206	-0.9882	2.0023
Bounds of effect size (%)	-254%	1%	-155%	97%	-6%	11%
Mean	30%		10%		18	
CI Black*Concordant	-0.3398	0.6601	-0.6938	0.8017	-2.9078	5.2420
Bounds of effect size (%)	-112%	218%	-721%	833%	-16%	29%
Mean	24%		17%		17	
CI Hispanic*Concordant	0.1109	0.9015	-0.2107	0.4378	-3.5749	1.5311
Bounds of effect size (%)	46%	372%	-125%	261%	-21%	9%
Mean	12%		10%		17	
CI Asian*Concordant	-0.0432	0.6631	-0.1946	0.4796	-2.1907	3.6496
Bounds of effect size (%)	-36%	553%	-190%	468%	-13%	22%

Mean rates are based on the sample for which both patient and physician race are observed (Table 1). Confidence intervals come from the coefficients from the models that include physician fixed effects presented in Table 3.5.

Section 5: Conclusions

Discrimination may contribute to racial/ethnic disparities in health care use and health outcomes. Concordance between patients and physicians is hypothesized to reduce discrimination by fostering favorable prejudice, modification of negative stereotypes, increased clinical certainty, trust and compliance. While previous research on the impact of concordance has focused on patients' subjective views of their satisfaction with and quality of care, this analysis estimates the role of concordance in affecting more objective quality of care measures in the outpatient setting. In addition to examining a new set of outcomes, I addressed methodological weaknesses in the existing literature to separately measure the effects of patient race, physician race, and racial concordance. By restricting the analysis to physicians who see both white and minority patients, I also more accurately measure the impact of concordance on the mechanisms associated with discrimination, as opposed to health care system factors that are correlated with concordance in geographic areas with high concentrations of minority patients and physicians.

The results from this analysis suggest that, after controlling for patients' demographic characteristics and health insurance status and restricting to providers who see both white and minority patients, we do not see large significant differences in rates of preventive care screenings or in the length of primary care visits. To the extent that there are differences, they are consistent with statistical discrimination driven by differences in underlying prevalence (black men being more likely to have a cholesterol screening) as opposed to unfavorable prejudice. Generally speaking, physician race is a much more important predictor of outcomes than patient race or concordance. Two exceptions are that concordance increases cholesterol screening rates by 2-3 times among black and Hispanic men and appears to have a negative impact on tobacco cessation counseling among blacks and Hispanics. Minority physicians are more likely to provide high quality care on some measures and less likely on others. Given the relatively large magnitudes of the effects of physician race relative to the small contributions of concordance, for some outcomes minority patients may actually be worse off in concordant pairs. These results highlight the importance of addressing physician education and training to improve quality of care in addition to increasing the number of minority physicians.

The analysis of how patients are distributed across primary care physicians also suggests that across-provider variation is likely to be an important contributor to disparities. Since the policy options to address differences driven by these mechanisms are quite different from those that address within-provider variation, future research should identify the contribution of across-provider variation and the specific factors (practice style, geographic characteristics and local resources) that are most important.

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